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Vol. 56, No. 4

The Costs of Sovereign Default

EDUARDO BORENSZTEIN and UGO PANIZZA*

This paper empirically evaluates four types of costs that may result from an international sovereign default: reputational costs, international trade exclusion costs, costs to the domestic economy through the financial system, and political costs to the authorities. It finds that the economic costs are generally significant but short-lived, and sometimes do not operate through conventional channels. The political consequences of a debt crisis, by contrast, seem to be particularly dire for incumbent governments and finance ministers, broadly in line with what happens in currency crises. [JEL F34, F36, H63, G15]
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There is broad consensus in the economic literature that the presence of costly sovereign defaults is the mechanism that makes sovereign debt possible (Dooley, 2000). In the case of sovereign debt, creditor rights are not as strong as in the case of private debts. If a private firm becomes insolvent, creditors have a well-defined claim on the company's assets even if they may be insufficient to cover the totality of the debt. These legal rights are

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necessary for private debts to exist.[1] In the case of a sovereign debt, in contrast, the legal recourse available to creditors has limited applicability because many assets are immune from any legal action, and uncertain effectiveness because it is often impossible to enforce any favorable court judgment.[2] But the literature sustains that sovereign debt markets are still viable because, if defaults are costly in some way to the borrowing country, there will be an incentive to repay debts, regardless of the effectiveness of legal recourse. It is noteworthy that we use the term default to encompass any situation in which the sovereign does not honor the original terms of the debt contract, including voluntary restructurings where there is a loss of value for the creditors. This is entirely in line with the concept applied by credit-rating agencies.

There is much less agreement on what the costs of default actually are, let alone their magnitude. Traditionally, the sovereign debt literature has focused on two mechanisms: reputational costs, which in the extreme could result in absolute exclusion from financial markets, and direct sanctions such as legal attachments of property and international trade sanctions imposed by the countries of residence of creditors. The reputational cost of default has a well-established theoretical and historical tradition, with Eaton and Gersovitz (1981) presenting the canonical, formal model. An influential article by Bulow and Rogoff (1989a), however, casts doubts on the validity of the reputational cost, and points instead to direct sanctions—such as trade embargoes—as the only viable mechanism that makes governments repay their debts. While their argument may not be robust to other model specifications, there is a widespread body of literature based on the sanctions view.[3] But there is comparatively little work on assessing the empirical relevance of these mechanisms. An exception is Tomz (2007), who based on an extensive review of historical case studies, finds widespread evidence in favor of the importance of reputation in financial markets, in contrast to the view that seemed to prevail earlier (for example, Lindert and Morton, 1989).[4]

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[1]In fact, even in private markets debt contracts are not fully enforceable. Djankov and others (2007) show that creditor protection through the legal system is positively correlated with the development of the private credit market.

[2]Some recent litigation strategies against sovereigns in default appear to focus on becoming enough of a nuisance such that sovereigns would acquiesce to an out-of-court settlement, rather than seeking a direct enforcement of property rights. Those strategies, however, can succeed only if the plaintiffs hold a small fraction of the debt. For a detailed discussion of the law and economics of sovereign debt, see Panizza, Sturzenegger, and Zettelmeyer (forthcoming).

[3]Influential papers that base their results on the assumption that default causes a direct loss of output or trade access—in line with the sanctions view—includes Krugman (1988) and Sachs (1989).

[4]For recent reviews, see De Paoli, Hoggarth, and Saporta (2006); Hatchondo, Martinez, and Sapriza (2007); and Panizza, Sturzenegger, and Zettelmeyer (forthcoming).

More recently, recognizing that holders of government debt are not only foreign investors (in fact, perhaps a majority of investors in government bonds are domestic institutions and resident individuals in many cases nowadays) more attention has been paid to the consequences of default for the domestic economy, in particular the banking sector.

This channel is particularly relevant because, in many emerging economies, banks hold significant amounts of government bonds in their portfolios. Thus, a sovereign default would weaken their balance sheets and even create the threat of a bank run. To make matters worse, banking crises are usually resolved through the injection of government ‘‘recapitalization’’ bonds and central bank liquidity. But in a debt crisis, government bonds have questionable value and the domestic currency may not carry much favor with the public either. A corollary of the domestic economic costs of debt crises is that they may also involve a political cost for the authorities. A declining economy and a banking system in crisis do not bode well for the survival in power of the incumbent party and the policymaking authorities. Although such linkage has been noted in the case of currency devaluations, for example, it has not been explored in the case of debt defaults.

This paper evaluates empirically each one of the suspected mechanisms through which default costs may affect a sovereign government. It should be recognized at the outset that it is quite difficult to find econometrically sound ways to isolate the costs of default. For instance, while it is easy to find a negative correlation between default and growth, it is much more difficult to test whether this negative correlation is driven by the default episode or by a series of other factors that are the cause of both the debt default and an economic recession. Moreover, it is also hard to identify the direction of causality between growth and default.

Thus, this paper has more modest objectives. Rather than attempting to quantify precisely the costs of default on sovereign debt, the objective is to evaluate if there is some empirical basis for—or lack of evidence against—each one of the mechanisms that are believed to be relevant, and perhaps discard those mechanisms that appear to be less consistent with the data.

In addition to the traditional reputational and trade sanctions, the paper explores the significance of effects that operate through the domestic banking system and the political costs of default for the government.[5]

Identifying the channel and magnitude of the costs of sovereign default with some degree of precision would be important for a number of reasons. The ‘‘default point’’ for a sovereign should be the point at which the cost of servicing debt in its full contractual terms is higher than the costs incurred from seeking a restructuring of those terms, when these costs are comprehensively measured. An accurate measure of the default point is

[5]This paper does not explore the role of collateral. For a discussion of this issue see Dooley, Garber, and Folkerts-Landau (2007).

necessary, for example, to assess how "safe" a certain level of debt is, namely, how likely it is that an economic shock would trigger a situation of default.[6] In fact, it is not possible to compute the probability of default, or to price a sovereign bond without making a judgment about the default point.

From a policy perspective, an understanding of the channels through which default costs apply can help design initiatives to improve the functioning of international financial markets and lower the cost of borrowing for many sovereigns. For example, if the costs of default apply largely through international trade, a more open economy would have a higher default point than a more closed economy, other things equal, and would be less risky for lenders, which would result in lower borrowing costs.

This paper analyzes the incidence of four types of cost that may result from an international sovereign default: reputational costs, international trade exclusion costs, costs to the domestic economy through the financial system, and political costs to the authorities. We find that reputational costs, as reflected in credit ratings and interest rate spreads, are significant but appear to be short-lived; that despite evidence that trade and trade credit are negatively affected by default, controlling for trade credit does not seem to modify the effect of default on trade; that growth in the domestic economy suffers, and more so in cases where the causes for default seem less compelling, although this effect also seems to be short-lived; that default episodes seem to cause banking crises and not vice versa, but that—outside of banking crisis episodes—more credit dependent industries do not suffer more than other industries following a sovereign default; and that the political consequences of a debt crisis are dire for incumbent governments and finance ministers, broadly in line with what happens in currency crises.

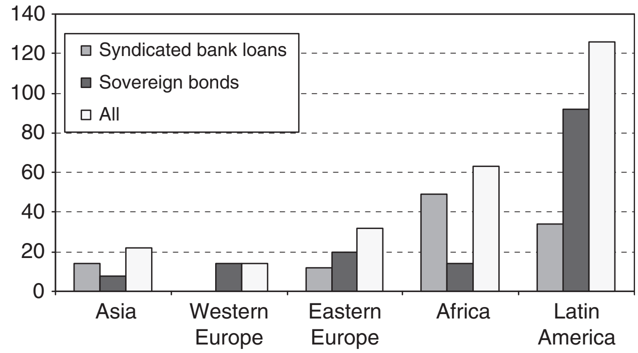
I. Two Hundred Years of Sovereign Default

Dating sovereign default episodes and measuring their duration is not a straightforward exercise. Table A1 uses four different sources to classify default episodes over the last 200 years.[7] Although there is substantial coincidence between the four sources, the match is far from perfect. There are, for instance, several episodes that are classified as defaults by Standard & Poor's but not classified as defaults by Beim and Calomiris (2000), and also a few

[6]This is analogous to the evaluation of the probability of default by a private company. Its default point, in theory, is the point at which existing liabilities equal the total market value of its assets, that is, its equity value is zero. See Merton (1974) and Kealhofer (2003). For an application to the sovereign case, see Gapen and others (2005).

[7]The first four columns of the table use data from Standard and Poor's and include all defaults on sovereign bonds and bank loans. Columns 5 and 6 are from Beim and Calomiris (2000) and also include defaults on suppliers' credit. Column 7 is from Sturzenegger and Zettelmeyer (2006) and is based on primary data from Beim and Calomiris (2000), and Lindert and Morton (1989). The last column uses data from Detragiache and Spilimbergo (2001). The definitions of default episodes applied by each one of these sources are presented in Table A1.

Figure 1. Number of Defaults (1824–2004)



Note: This figure plots the geographical distribution of sovereign defaults that took place over a period 1824–2004. The figure also divides default episodes between default on sovereign bonds and sovereign syndicated bank loans. Default episodes are identified using the Standard & Poor's definition of sovereign default (see Table A1).

episodes that are classified as defaults by Beim and Calomiris (2000) and not by Standard and Poor's. There are also differences in the methodology used to measure the length of a default episode. Beim and Calomiris (2000), for instance, find fewer but longer lasting default episodes because they tend to merge into a unique episode defaults that occurred within five years. The methodology used by Detragiache and Spilimbergo (2001), instead, leads to code as defaults several episodes that are not classified as defaults by Standard and Poor's.[8] Largely on the basis of its completeness, the rest of the paper will use Standard and Poor's classifications as reported in the first four columns of Table A1. Moreover, alternative definitions of default include debt rescheduling with official creditors which, in our view, involves a different cost/benefit analysis on the part of the debtor governments, because financial relationships with multilateral institutions or other governments are based on different principles from those of private markets. Having said this, the results of the paper are robust to using the definition of default adopted by Detragiache and Spilimbergo (2001) or to including the Paris Club rescheduling events studied by Rose (2005). Furthermore, while some papers (e.g. Kraay and Nehru, 2006; and Pescatori and Sy, 2007) include episodes of near default in their definition of debt crisis, we do not include these episodes in our empirical exercises because we are interested in the cost of actual default, instead of the cost of debt crises that do not result in an actual default.

Figure 1 shows the number of default episodes by geographical area for the period from 1824 to 2004. Latin America is the region with the highest

[8]This is the case, for instance, of Nigeria, Zambia, and Sierra Leone in the 1970s; Egypt and El Salvador in the 1980s; and Sri Lanka, Thailand, Korea, and Tunisia in the 1990s.

Table 1. Default Episodes

	Period	Africa			Asia			Eastern Europe			Latin America			Western Europe			All		
		Bond	Bank	All	Bond	Bank	All	Bond	Bank	All	Bond	Bank	All	Bond	Bank	All	Bond	Bank	All
Number of episodes	1824–40	0	0	0	0	0	0	1	0	1	14	0	14	4	0	4	19	0	19
Average length		0.0	0.0	0.0	0.0	0.0	0.0	52.0	0.0	52.0	21.4	0.0	21.4	24.8	0.0	24.8	23.7	0.0	23.7
Number of episodes	1841–60	0	0	0	0	0	0	0	0	0	5	0	5	1	0	1	6	0	6
Average length		0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	8.0	0.0	8.0	6.0	0.0	6.0	7.7	0.0	7.7
Number of episodes	1861–80	3	0	3	1	0	1	0	0	0	14	0	14	1	0	1	19	0	19
Average length		10.0	0.0	10.0	5.0	0.0	5.0	0.0	0.0	0.0	14.1	0.0	14.1	2.0	0.0	2.0	12.4	0.0	12.4
Number of episodes	1881–1900	0	0	0	0	0	0	2	0	2	16	0	16	1	0	1	19	0	19
Average length		0.0	0.0	0.0	0.0	0.0	0.0	2.0	0.0	2.0	3.4	0.0	3.4	9.0	0.0	9.0	3.6	0.0	3.6
Number of episodes	1901–20	5	0	5	1	0	1	2	0	2	11	0	11	1	0	1	20	0	20
Average length		1.6	0.0	1.6	13.0	0.0	13.0	2.5	0.0	2.5	5.3	0.0	5.3	1.0	0.0	1.0	4.3	0.0	4.3
Number of episodes	1921–40	1	0	1	4	0	4	8	0	8	20	0	20	6	0	6	39	0	39
Average Length		3.0	0.0	3.0	7.3	0.0	7.3	12.6	0.0	12.6	9.3	0.0	9.3	6.7	0.0	6.7	9.2	0.0	9.2
Number of episodes	1941–60	0	0	0	1	0	1	2	0	2	1	0	1	0	0	0	4	0	4
Average length		0.0	0.0	0.0	10.0	0.0	10.0	13.5	0.0	13.5	1.0	0.0	1.0	0.0	0.0	0.0	9.5	0.0	9.5
Number of episodes	1961–70	1	0	1	0	0	0	0	0	0	1	0	1	0	0	0	2	0	2
Average length		15.0	0.0	15.0	0.0	0.0	0.0	0.0	0.0	0.0	1.0	0.0	1.0	0.0	0.0	0.0	8.0	0.0	8.0
Number of episodes	1971–80	0	5	5	0	3	3	0	0	0	0	7	7	0	0	0	0	15	15
Average length		0.0	13.6	13.6	0.0	16.0	16.0	0.0	0.0	0.0	0.0	4.9	4.9	0.0	0.0	0.0	0.0	10.0	10.0
Number of episodes	1981–90	1	33	34	0	6	6	0	5	5	5	24	29	0	0	0	6	68	74
Average length		2.0	8.5	8.4	0.0	10.0	10.0	0.0	5.6	5.6	3.6	8.7	7.8	0.0	0.0	0.0	3.3	8.5	8.1
Number of episodes	1991–2004	3	11	14	1	5	6	5	7	12	5	3	8	0	0	0	14	26	40
Average length		2.0	5.1	4.4	1.0	2.2	2.0	1.4	5.7	3.9	1.6	4.0	2.5	0.0	0.0	0.0	1.6	4.6	3.5

Note: This table lists the number and average duration of sovereign default episodes that took place in the period 1824–2004 by geographical region and time period. The table also divides default episodes between defaults on sovereign bonds and sovereign syndicated bank loans. Default episodes are identified using the Standard & Poor's definition of sovereign default (see Table A1).

number of default episodes at 126, Africa, with 63 episodes, is a distant second. The Latin American "lead" is, however, largely determined by the fact that Latin American countries gained independence and access to international financial markets early in the 19th century, while most African countries continued to be European colonies for another 100 or 150 years Among the developing regions, Asia shows the lowest number of defaults. Table 1 groups the various default episodes by time period and geographical area. Besides reporting the number of episodes, the table also reports the average length of the episodes.

As noted by Sturzenegger and Zettelmeyer (2006), default episodes tend to happen in clusters and usually follow lending booms. The first cluster of defaults happened in the period that spans from 1824 to 1840 and followed a lending boom driven by the newly acquired independence of most Latin American countries. Of 19 default episodes recorded during this period, 14 involved Latin American countries. The other five default episodes involved Greece, Portugal, and Spain (three episodes). The average length of the default episodes of this period (more than 20 years) suggests difficult restructuring processes.

The following period (1841–60) was relatively tranquil and comprised only six default episodes. However, a lending boom developed at this time, which soon resulted in a new series of default episodes (Lindert and Morton, 1989; Sutter, 2003). The period from 1861 to 1920 was characterized by 58 default episodes, including 41 episodes in Latin America and eight in Africa.[9] Resolution of default improved dramatically in speed, with the length of the average default period dropping to less than five years by 1881–1920.

The next wave of defaults was associated with the Great Depression and the Second World War. The 1921–40 period was punctuated by 39 default episodes. Again, more than half of these defaults happened in Latin American countries and more than one-third of them (16 episodes) in Europe. This is the last period in which we observe debt default episodes among western European countries.

By the end of the war, most developing countries had completely lost access to the international capital market. As a consequence, over the period that goes from 1941 to 1970 we observe very few default episodes (six episodes in total).[10] Lending to developing countries restarted timidly in the 1960s, but exploded after the oil shock of 1973 creating the need of recycling the earnings of oil-producing countries. One feature that differentiated the lending boom of the 1970s from previous ones is the vehicle used to extend

[9]This is also the period in which we observe the first default on bank loans (Russia in 1918).

[10]Of these six episodes, two were related to World War II (Hungary in 1941 and Japan in 1942), and other two were largely politically motivated defaults by communist countries (Czechoslovakia in 1959 and Cuba in 1960). The remaining two were Costa Rica (1962) and Zimbabwe (1965).

credit to developing countries. While in previous episodes developing countries borrowed by issuing bonds, in the 1970s most of the lending to developing countries took the form of syndicated bank loans. While the lending instrument was different, the fate of the lending boom did not differ, and the tranquil period was soon followed by a chain of defaults. Already in the 1970s, we observe 15 episodes of defaults on syndicated bank loans. The "debt crisis," however, did not erupt until the Mexican payment suspension of August 1982, which was soon followed by more than 70 default episodes (34 episodes involving African countries and 29 involving Latin American countries).

As in previous cases, credit to developing countries (including to countries that did not experience debt service disruptions) died out in the aftermath of the crisis and did not restart until the end of the restructuring process. The average default lasted approximately nine years, which suggests that restructuring syndicated bank loans was more cumbersome than restructuring international bonds. Eventually the defaulted bank loans were restructured by issuing new, partly collateralized, bonds that took the name of Brady Bonds (after the name of U.S. Treasury Secretary Nicholas Brady who was main architect of the restructuring process).

The Brady Plan played a key role in creating a bond market for debt issued by emerging market countries and, together with low interest rates in the United States, contributed to a new lending boom to emerging market countries (see Calvo, Leiderman, and Reinhart, 1993). The defaults that followed this new lending boom are recent history. Over the 1991–2004 period, we observed 40 defaults (14 on bonds and 26 on syndicated bank loans). Most of the syndicated bank loan defaults took place in Africa, where the bond instrument had not become widely used yet, while most of the bond defaults took place among Latin American issuers.

II. Default and GDP Growth

As a first stab at the issue at hand, we examine the effect of default on GDP growth. While this may not distinguish between competing theories of default costs, it can say something about the significance and lag structure of the costs. In addition, we are interested in exploring if the GDP costs are higher for countries that default in circumstances that seem less easily identified as an insolvency problem, what could in principle identify cases of "strategic" default.

Note that in this paper we will follow the traditional literature and focus on defaults on external debt. To the best of our knowledge, Reinhart and Rogoff's (2008) is the only attempt to document what happens around domestic debt defaults. They show that domestic defaults tend to happen during deep recessions and that, in the year of the default, GDP is 8 percent lower than GDP four years before the default. However, they also show that GDP starts recovering in the year after the default and that as early as three years after the default, GDP is back at the level it had four years before the default. In fact, when they compare domestic defaults with external defaults, Reinhart and Rogoff show that the former are characterized by a sharper contraction in the run up to the default, but also by a sharper recovery in the post-default period.

Before going into the details of the estimation, it should be acknowledged that there are unresolved endogeneity problems in the relationship between default and GDP growth, and this paper does not provide any breakthroughs in this regard. The problem is relevant because the theoretical literature has also noted the causation from weak growth to default. While the early literature based on Eaton and Gersovitz (1981) focused on ‘‘strategic’’ defaults (in the sense that defaults took place in good times when the country could easily have paid), recent work by Aguiar and Gopinath (2006) and Rochet (2006) shows that models that add persistent shocks to a simple Eaton and Gersovitz’s (1981) framework yield procyclical borrowing and (nonstrategic) default episodes. Arellano (2008) shows that even in the presence of i.i.d. endowment shocks, it is possible to generate a region of risky borrowing in which defaults take place after a negative shock. Mendoza and Yue (2008) adopt an alternative modeling strategy and show that defaults may lead to an inefficient reallocation of labor because they limit the ability of private agents to obtain the working capital necessary to buy imported inputs. Kohlscheen and O’Connell (2007) reach similar conclusions by focusing on the role of trade credit. Tomz and Wright (2007) show that a calibrated model based on Aguiar and Gopinath (2006) predicts that almost all defaults should happen during bad times. Next, they take the model to the data by using a large number of sovereign default episodes between 1820 and 2004 and show that about two-thirds of defaults happen when output is below trend. They argue that the difference between the predictions of the model and the data can be explained by political shocks and exogenous changes in the availability of international credit.

We follow Chuan and Sturzenegger (2005), who estimates several cross-section and panel growth regressions and find that default episodes are associated with a reduction in growth of approximately 0.6 percentage points. If the default coincides with a banking crisis, the effect is much larger and growth decreases by 2.2 percentage points. In Table 2, we present results

Table 2. Default and Growth, Panel, 1972–2000

	(1) GROWTH	(2) GROWTH	(3) GROWTH	(4) GROWTH
INV_GDP	1.211	1.152	1.205	1.146
	(8.63)***	(8.08)***	(8.58)***	(8.04)***
POP_GR	−0.120	−0.119	−0.121	−0.118
	(1.22)	(1.22)	(1.24)	(1.20)
GDP_PC70s	−0.121	−0.124	−0.121	−0.125
	(7.25)***	(7.34)***	(7.24)***	(7.37)***
SEC_ED	0.014	0.018	0.014	0.018
	(1.62)	(2.03)**	(1.63)	(2.03)**
POP	0.004	0.004	0.004	0.004
	(6.32)***	(6.72)***	(6.30)***	(6.66)***

Table 2 *(concluded)*

	(1) GROWTH	(2) GROWTH	(3) GROWTH	(4) GROWTH
GOV_C1	2.965	2.974	2.970	3.000
	(2.91)***	(2.89)***	(2.89)***	(2.89)***
CIV_RIGTH	−0.026	−0.033	−0.026	−0.035
	(0.37)	(0.45)	(0.37)	(0.49)
DTOT	−0.270	−0.111	−0.277	−0.082
	(0.22)	(0.10)	(0.23)	(0.07)
OPEN	2.149	2.156	2.151	2.146
	(3.50)***	(3.50)***	(3.49)***	(3.48)***
SSA	−0.859	−0.832	−0.839	−0.788
	(2.84)***	(2.70)***	(2.73)***	(2.54)**
LAC	−0.399	−0.430	−0.367	−0.355
	(1.60)	(1.70)*	(1.45)	(1.39)
TRANS	−0.064	−0.266	−0.071	−0.268
	(0.10)	(0.44)	(0.11)	(0.44)
BK_CR	−1.087	−1.068	−1.092	−1.080
	(4.64)***	(4.53)***	(4.65)***	(4.57)***
DEF	−1.239	−1.184	−1.282	−1.370
	(4.32)***	(3.82)***	(4.38)***	(4.06)***
DEF_B		−1.388		−1.291
		(2.11)**		(1.93)*
DEF_B1		0.481		0.916
		(0.87)		(1.49)
DEF_B2		0.337		0.495
		(0.63)		(0.82)
DEF_B3		0.994		1.242
		(1.55)		(1.90)*
END_DEF			−0.665	−1.135
			(1.14)	(1.77)*
END_DEF1			0.002	0.003
			(0.00)	(0.01)
END_DEF2			0.122	−0.384
			(0.22)	(0.70)
Constant	1.387	1.474	1.389	1.471
	(2.16)**	(2.28)**	(2.16)**	(2.28)**
Observations	2,048	1,985	2,048	1,985
R-squared	0.22	0.22	0.22	0.22

Note: This table shows the results of a set of regressions where the dependent variable is annual real per capita GDP growth and the explanatory variables are investment divided by GDP (INV_GDP); population growth (POP_GR); GDP per capita in 1970 measured in U.S. dollars (GDP_PC70s); percentage of the population that completed secondary education (SEC_ED); total population (POP); lagged government consumption over GDP (GOV_C1); an index of civil rights (CIV_RIGHT); the change in terms of trade (DTOT); the degree of openness (OPEN, defined as exports plus imports divided by GDP); a dummy variable taking a value of one in presence of a banking crisis (BK_CR); three regional dummies for sub-Saharan Africa (SSA), Latin America and Caribbean (LAC), and transition economies (TRANS); a dummy variable that takes a value of one each year that a country is in default and zero otherwise (DEF); and dummy variable that takes a value of one in the first year of a default episode and zero otherwise (DEF_B); (DEF_B1, DEF_B2, and DEF_B3 are the first, second and third lags of DEF_B). Robust *t*-statistics in parentheses. * significant at 10 percent; ** significant at 5 percent; *** significant at 1 percent.

from several regressions aimed at estimating the relationship between default and growth. In all regressions we use an unbalanced panel that includes up to 83 countries for the 1972–2000 period, and estimate the following model:

$$GROWTH_{i,t} = \alpha + \beta X_{i,t} + \gamma DEFAULT_{i,t} + \varepsilon_{i,t}, \quad (1)$$

where $GROWTH_{i,t}$ is per capita annual real GDP growth in country i and year t, X is a matrix of controls,[11] and *DEFAULT* is a set of dummy variables tracking default episodes. In column 1, the variable *DEF* takes a value of one each year that a country is in default and zero otherwise. We find that, on average, default is associated with a decrease in growth of 1.2 percentage points per year. This figure is consistent with Chuan and Sturzenegger's (2005) finding that default has a negative effect on growth that ranges between 0.5 and 2 percentage points.

We next explore the dynamic structure of the impact of default. In column 2, we augment the regressions with a variable that takes a value of one at the beginning of the default episode (*DEF_B*) and three lags of this variable (*DEF_B1, DEF_B2*, and *DEF_B3*). We find that the impact of default seems to be short-lived. We estimate a large effect in the first year of the default episode (with a drop in growth of 2.6 percentage points), and we find no statistically significant effect of the lagged default variables. This is consistent with results in Levy, Yeyati, and Panizza (2005), who, using quarterly data, find that crises precede defaults, and that defaults tend to occur at the trough of the recession.

As a check on the validity of the above result, we test whether the estimated negative effect of default is in fact an artifice of the rebound in growth that tends to occur in the post-default years. To control for this possibility, in columns 3 and 4 we augment the regressions with a dummy variable that takes a value of one when a country exits from default (*END_DEF*) and two lags of this variable.[12] We find that these dummy variables are not statistically significant and do not affect the estimated effect of default in the original regressions.

The direction of causality in the relationship between sovereign defaults and growth raises some questions. While the previous regressions suggest a robust association between debt defaults and low growth, they are only indicative of a correlation between the two variables. In fact, debt defaults are usually a consequence of some economic shocks, such as terms-of-trade

[11]Our set of controls includes the investment over GDP ratio (*INV_GDP*), population growth (*POP_GR*), GDP per capita in the early 1970s (*GDP_PC70s*), percentage of the population that completed secondary education (*SEC_ED*), total population (*POP*), lagged government consumption over GDP (*GOV_C1*), an index of civil rights (*CIV_RIGHT*), the change in terms of trade (*DTOT*), the degree of openness (*OPEN*), a dummy variable taking a value of one in presence of a banking crisis (*BK_CR*), and three regional dummies for sub-Saharan Africa (*SSA*), Latin America and Caribbean (*LAC*), and transition economies (*TRANS*). Substituting country fixed for the regional dummies does not change the results.

[12]That is, if a country was in default from 1982 to 1986, *END_DEF* takes a value of one in 1987.

shocks, sudden stops, currency crises, and so on that also hurt growth in some fashion. Moreover, the *anticipation* of a default episode (rather than the default) may carry substantial costs (Levy, Yeyati, and Panizza, 2005). While the regressions in Table 2 control for some of these effects (for instance, they control for banking crises), they cannot account for all the variables that jointly affect the probability of a sovereign default and an economic recession. Hence, lower growth might not be the consequence of default but of other factors that also affect debt sustainability.

Identifying the causal effect of default on growth would require an instrument for default (that is, a variable that affects the probability of default without having a direct effect on GDP growth). Unfortunately, such instrument has not been found and it may not exist. Here we set a more modest objective and use a two-stage approach to attempt to disentangle the effect of the "predictable" component of default and the unpredictable one. That is, we try to decompose the correlation between default and economic growth in two parts: the effect owing to all the variables which can be used to predict the probability of default and a residual effect, which we interpret as the decision of default itself, over and above its causes. More precisely, the default dummy can be statistically divided into two components:

$$default_{i,t} = pred\_def_{i,t} + v_{i,t}, \tag{2}$$

where $pred\_def_{i,t,}$ denotes the predicted probability of default obtained by running a logit regression of $default_{i,t}$ on a set of standard predictors of default, and $v_{i,t}$ is the error term of the logit model.[13]

Within this setup, $pred\_def_{i,t}$ captures the predicted effect of default and proxies for the fact that an increase in the probability of default may have a direct effect on growth, but $v_{i,t}$ captures the unpredictable components of default. After having estimated the anticipated and unanticipated component of default, we can include these two variables in a set of regressions similar to those of Table 2 and gauge their distinct effect on growth. As we predict default using a nonlinear model, this strategy is similar but not identical to directly adding to the original growth regression all the variables used to predict default. Again, the objective is not to identify the causal effect of default but just split the correlation between default and growth between a predictable and unpredictable component of default.

Table 3 presents the main results. As the sample of Table 3 is smaller than that of Table 2 (843 vs. 2,048 observations),[14] we start by reestimating the basic model of Table 2 for the restricted sample and check whether there are any differences in the estimated cost of default and we find that the results are

[13]To predict default we use model similar to that of Manasse, Roubini, and Schimmelpfennig (2003). Full regression results are available upon request.

[14]This is due to the fact that it does not make much sense to estimate the probability of default for industrial countries and, hence, Table 3 only includes developing countries. Furthermore, estimating the probability of default requires variables that are not available for all the countries included in the regressions reported in Table 2.

Table 3. Default and Growth, Panel, 1972–2000

	(1) GROWTH	(2) GROWTH	(3) GROWTH	(4) GROWTH
INV_GDP	1.607	1.584	1.635	1.584
	(5.11)***	(5.00)***	(4.58)***	(5.03)***
POP_GR	−0.331	−0.337	−0.319	−0.338
	(1.35)	(1.37)	(1.16)	(1.38)
GDP_PC70s	−0.259	−0.269	−0.300	−0.275
	(1.38)	(1.43)	(1.53)	(1.46)
SEC_ED	0.036	0.037	0.039	0.037
	(1.56)	(1.59)	(1.63)	(1.60)
POP	0.006	0.006	0.005	0.006
	(5.36)***	(5.29)***	(4.12)***	(5.25)***
GOV_C1	3.402	3.281	3.084	3.299
	(2.95)***	(2.76)***	(2.45)**	(2.75)***
CIV_RIGTH	−0.090	−0.093	−0.050	−0.092
	(0.71)	(0.73)	(0.36)	(0.72)
DTOT	−2.271	−2.333	−2.133	−2.342
	(1.20)	(1.23)	(1.16)	(1.24)
OPEN	1.764	1.816	1.677	1.818
	(1.52)	(1.55)	(1.31)	(1.55)
SSA	−0.542	−0.510	−0.637	−0.520
	(1.16)	(1.08)	(1.23)	(1.07)
LAC	−0.508	−0.457	−0.381	−0.460
	(1.41)	(1.26)	(1.03)	(1.26)
TRANS	−2.443	−2.437	−2.216	−2.430
	(2.53)**	(2.53)**	(2.02)**	(2.50)**
BK_CR	−1.364	−1.328	−1.188	−1.324
	(3.81)***	(3.73)***	(3.36)***	(3.71)***
DEF	−1.043			
	(3.15)***			
DEF_PR		−1.440	−1.246	−1.443
		(2.30)**	(1.89)*	(2.29)**
DEF_U		−0.930	−1.037	−0.930
		(2.46)**	(2.69)***	(2.32)**
DEF_PRB			−13.700	
			(2.13)**	
DEF_PRB1			0.330	
			(0.07)	
DEF_PRB2			3.506	
			(0.99)	
DEF_PRB3			−0.929	
			(0.24)	
DEF_UB			0.000	
			(0.00)	
DEF_UB1			1.098	
			(1.71)*	
DEF_UB2			0.860	
			(1.32)	
DEF_UB3			0.898	
			(1.17)	
END_DEF				−0.237
				(0.36)

Table 3 *(concluded)*

	(1) GROWTH	(2) GROWTH	(3) GROWTH	(4) GROWTH
END_DEF1				0.266
				(0.49)
END_DEF2				0.149
				(0.24)
Constant	2.629	2.660	–0.435	2.662
	(1.95)*	(1.97)**	(0.30)	(1.97)**
Observations	843	843	726	843
R-squared	0.26	0.26	0.28	0.26

Note: This table shows the results of a set of regressions where the dependent variable is annual real per capita GDP growth and the explanatory variables are investment divided by GDP (INV_GDP); population growth (POP_GR); GDP per capita in 1970 measured in U.S. dollars (GDP_PC70s); percentage of the population that completed secondary education (SEC_ED); total population (POP); lagged government consumption over GDP (GOV_C1); an index of civil rights (CIV_RIGHT); the change in terms of trade (DTOT); the degree of openness (OPEN, defined as exports plus imports divided by GDP); a dummy variable taking a value of one in presence of a banking crisis (BK_CR); three regional dummies for sub-Saharan Africa (SSA), Latin America and Caribbean (LAC), and transition economies (TRANS); a dummy variable that takes a value of one each year that a country is in default and zero otherwise (DEF); the predicted component of DEF (DEF_PR); the unexpected component of DEF (DEF_U); the predicted component of DEF measured in the first year of a default episode (DEF_PRB); and the unexpected component of DEF measured in the first year of a default episode (DEF_UB) (DEF_PRB1, DEF_PRB2, and DEF_PRB3 are the first, second and third lags of DEF_PRB; DEF_UB1, DEF_UB2, and DEF_UB3 are the first, second and third lags of DEF_UB); a dummy variable that takes a value of one in the first year after the end of a default episode (END_DEF ; END_DEF1 and END_DEF2 are the first and second lags of END_DEF). Robust *t*-statistics in parentheses. * significant at 10 percent; ** significant at 5 percent; *** significant at 1 percent.

basically unchanged (column 1, Table 3). In particular, we find that the effect of default is a bit smaller but, at 1 percent, still sizable, and it is still highly statistically significant.

The split between anticipated and unanticipated components of default reveals that both variables are statistically significant. The estimate reported in column 2 of the anticipated effect (*DEF_PR*), at 1.4 percent, is slightly larger than the unanticipated component (*DEF_U*), which is close to 1 percent. This suggests that the default decision itself may involve significant collateral costs for the domestic economy.

In column 3, we estimate the dynamic structure of the anticipated and unanticipated components of default.[15] We find that the anticipated default effect (*DEF_PRB*) is on impact negative, quite large, and statistically

[15]In order to estimate the probability of the beginning of the default episode, we used the logit described in Table A2 but restricted the dependent variable to take value one only in the first year of a default episode.

significant (we investigated whether the large coefficient of *DEF_PRB* was due to the presence of outliers but were unable to find evidence in this direction).

In contrast, while we find that the unanticipated component of default is still large and statistically significant, we find no significant negative effect in the first year. In the last column, we augment the regression in column 2 with *END_DEF* and its two lags and we find that the results are unchanged.[16]

We note that an alternative interpretation of the effect of the unexpected portion of the default variable is that it captures the cost of "unjustified" defaults, under the assumption that the magnitude of the costs of default to a country depends on whether the default was unavoidable or resulted from a weak willingness to pay. Much of the sovereign debt literature emphasizes the distinction between "ability" to pay and "willingness" to pay. The markets would punish debtors in the latter case, but will be more forgiving in the former case (see Grossman and van Huyck, 1988).[17] From this perspective, the specification above can be interpreted as a measure of the degree to which a default was justified by fundamental economic conditions.

One caveat with the previous analysis is that the variables included in the first stage are likely to be a subset of the information available to governments and markets. If one assumes that Grossman and van Huyck (1988) are right, such errors in our first stage regression would lead to an overestimation of the cost of anticipated defaults and an underestimation of the costs of unpredictable defaults.

We now turn to the investigation of the specific channels through which default may have a negative impact on growth.

III. Default and Reputation

As argued at the beginning of this paper, whether reputation has a significant effect or not plays a key role on the timing and the circumstances under which a sovereign will initiate a debt restructuring action.

In the seminal paper by Eaton and Gersovitz (1981), international lending is sustained by the fact that defaults are associated with a permanent exclusion from future borrowing. This assumption, however, was soon criticized because the threat of a permanent exclusion from the capital market is not time-consistent. Moreover, Bulow and Rogoff (1989b) argued that, even in the presence of such a threat, a defaulter could smooth

[16]One problem with the regressions of Tables 2 and 3 is that they are based on annual information and hence they cannot capture the precise timing of the default. Levy, Yeyati, and Panizza (2005) study the impact of default on growth by looking at quarterly data for emerging economies and find that output contractions precede defaults, and that the trough of the contraction coincides with the quarter of default.

[17]Alternatively one could try to identify the "avoidable" or unjustified defaults directly, but there are few cases that could clearly be labeled as resulting from lack of willingness to pay. Nearly all unilateral sovereign debt repudiation cases have stemmed from communist revolutions or other radical political postures, and the economic downturns probably resulted more from those political changes than from the debt defaults themselves.

consumption by purchasing insurance or investing a portion of its wealth abroad. For these reasons, Bulow and Rogoff (1989a) argued that positive international lending cannot be sustained without some form of direct punishment (such as trade embargoes).

However, a series of more recent papers argued that reputational concerns can sustain positive lending even in the absence of a threat of permanent exclusion from future borrowing. In Kletzer and Wright (2000), lenders collusion is guaranteed by the fact that the original lender can punish new lenders by forgiving the defaulter if the defaulter stops servicing the loans granted by new lenders. Wright (2003) shows that, even in the presence of contracts à la Bulow-Rogoff, the existence of syndicated lending generates incentives to collude in punishing default. The key idea of Kletzer and Wright's (2000) and Wright's (2003) models is that, rather than triggering permanent exclusion from credit markets, a default leads to a new financial relationship in terms at which the defaulter's utility is the same as that would result from permanent exclusion.[18]

Studies that provide empirical evidence in support to the "reputation view" include those by English (1996) and Tomz (2007). English (1996) focuses on defaults by U.S. states in the 19th century and argues that, because foreign creditors could not impose trade embargoes on the U.S. states, states that paid back their debt did so for reputational reasons alone, and not because of the threat of sanctions. Tomz (2007) describes the evolution of sovereign debt over the last three centuries and presents several stylized facts which are consistent with the reputational view. For instance, he shows that unjustified defaults during World War II (these are defaults by countries that did not participate in the war) triggered an exclusion from the international capital market which was twice as long as the exclusion suffered by "expected" defaulters. He also shows that "surprising" payers (that is, countries that were expected to default in the 1930s but did not default) were rewarded by investors with lower spreads when these countries re-accessed the international capital market. He also deconstructs the sanction view and argues that there is no evidence that defaults ever led to sanctions. He argues that the conventional argument that in the 1930s Argentina repaid its debt to avoid a trade embargo from the United Kingdom (Díaz-Alejandro, 1983) is not correct and provides evidence suggesting that Argentina repaid its foreign debt in order to strengthen its reputation of good debtor.[19] Although Mitchener and Weidenmier (2005) find that military pressure or political control were common responses to default episodes in the gold standard period, Tomz (2007) argues that gunboat diplomacy was driven by civil wars, territorial conflicts, and tort claims and not by default episodes.

[18]An ever more recent literature focuses on reputation vis-à-vis domestic agents (see Panizza, Sturzenegger, and Zettelmeyer, forthcoming, for a survey).

[19]Argentine Finance Minister Alberto Hueyo stated: "To honor existing commitments is always highly honorable, but to do it when everyone is failing to and at times of hardship … is a thousand times more valuable" (quoted in Tomz, 2007).

Although the evidence presented by English (1996) and Tomz (2007) in favor of the reputation view is persuasive, it remains unclear whether reputational costs alone are enough to justify the existence of the sovereign debt market. For instance, Arellano and Heathcote (2008) show that models in which if the only cost of default is *permanent* exclusion from future borrowing yield maximum sustainable debt levels which are much lower than the debt levels we observe in reality.

Regardless of the reasons that led Argentina or the U.S. states to repay their debts, there is by now agreement on the fact that default does not lead to a permanent exclusion from the international capital market. Although there is some capital market exclusion period following a default, countries that defaulted in the last three decades have regained access to international capital markets fairly quickly. Gelos, Sahay, and Sandleris (2004) find that countries that defaulted in the 1980s were able to regain access to international credit in about four years. Richmond and Dias (2008) study all defaults that took place between 1980 and 2005 and find that, on average, defaulters regain partial market access after 5.7 years and full market access after 8.4 years. Levy Yeyati (forthcoming) shows that countries that defaulted in the 1970–2004 period received lower net transfers in the years that followed the default episode but that the effects were not very large (they range between 0.1 and 1 percentage point of GDP). Thus, the evidence suggests that, while countries lose access during default, once the restructuring process is fully concluded, financial markets do not discriminate, in terms of access, between defaulters and nondefaulters. External factors and the mood of foreign investors seem to be far more important than default history in determining access to the international capital market. One example of this behavior can be found by observing that in the period that goes from the 1930s to the 1960s all Latin American countries were excluded from the world capital market, and this exclusion reached both countries that defaulted in the 1930s and countries, like the case of Argentina commented above, which had made a successful effort to avoid default. The recent lending booms and default experiences also provide evidence in the same direction. Several countries that had defaulted in the 1980s were able to attract large capital flows in the 1990s and countries that defaulted in the late 1990s regained access to the international capital market almost immediately after their debt restructurings. In fact, Richmond and Dias (2008) show that external financial market conditions are the most important factor in determining the speed with which defaulters are able to re-access the international capital markets.[20]

[20]An under-researched topic concerns the relationship between the size of the haircut and the conditions under which defaulters re-access to capital market (Panizza, Sturzenegger, and Zettelmeyer, forthcoming, and Trebesch, 2009, include discussions of this issue).

There is some evidence suggesting that markets also discriminate in terms of *cost* of credit, in the sense that default history is positively correlated with borrowing costs. What is not clear, however, is whether this effect is long lasting or not. In what follows, we review the existing literature and provide some new evidence.

Studies that measured the impact of default on borrowing costs have focused on both indirect and direct measures. The main indirect measure in this line of work is a country's credit rating. This is a relevant measure because credit ratings tend to be highly correlated with borrowing costs. Cantor and Packer (1996) were among the first to highlight the link between default history and credit ratings. In their study, they collect data for approximately 50 countries and regress credit ratings in 1995 on a set of eight explanatory variables, and find that this relatively small set of independent variables explains more than 90 percent of the variance in credit ratings.[21] They also find that a dummy variable that takes value one for countries that defaulted after 1970 is highly significant and associated with a drop of two notches in a country's credit rating. Along similar lines, Reinhart and others (2003) find that a history of default is associated with lower ratings assigned by *Institutional Investor*.

One important question that the literature does not seem to address is whether default has a long-term impact on credit ratings. That is, how long is the markets' memory? To answer this question, we estimate the following cross-country model:

$$RATING_i = \alpha + \beta X_i + \gamma DEFAULT_i + \varepsilon_i, \tag{3}$$

where *RATING* measures average credit ratings over the 1999–2002 period, *X* is a set of explanatory variables also measured over the 1999–2002 period and *DEFAULT* is the variable measuring previous history of default.[22]

We measure credit ratings by converting Standard and Poor's foreign-currency long-term credit ratings into numerical values (20 corresponds to AAA, 19 to AA+, 18 to AA, and so forth, all the way down to selective default rating, *SD*, which is assigned a value of zero). In selecting the explanatory variables we follow Cantor and Packer (1996) and include the log of GDP per capita *(LGDP_PC*), GDP growth (*GDPGR*), the log of inflation (*LINF*), the central government balance scaled by GDP (*CG_BAL* takes positive values for fiscal surpluses and negative values for deficits), the external current account balance scaled by GDP (*CA_BAL*), external debt over exports (EXDEXP), and a dummy variable that takes value one for

[21]It is remarkable that GDP per capita by itself explains 80 percent of the variance of credit rating, a fact not highlighted in the original paper (thanks to Kevin Cowan for pointing this out).

[22]We also estimated the model using average ratings for the 2000–04 period, and the set of explanatory variables averaged over the 1990–2000 period. The results did not change.

industrial countries (*IND*).[23] In column 1 of Table 4 we follow Cantor and Packer (1996) and measure the history of default with a dummy variable that takes value one if country *i* has defaulted over the 1970–2002 period and zero otherwise. Most variables have the expected sign and are statistically significant (the exceptions are GDP growth which has the wrong sign but is not statistically significant and the current account balance which has the expected sign but is not statistically significant). As in Cantor and Packer (1996), we find that this limited set of control variables explains more than 90 percent of the cross-country variance of credit ratings (the R^2 of the regression is 0.91). We also find that default history is negatively correlated with credit ratings. In particular, our point estimates indicate that default history leads to a drop in credit rating of 1.7 notches, slightly lower than the estimate of Cantor and Packer (2.5 notches).

In column 2, we add two control variables that have been used in previous studies. The first variable is public debt over GDP (*DEBT_GDP*) and the second is the index of original sin (*OR_SIN*), developed by Eichengreen, Hausmann, and Panizza (2005). Both variables have the right sign and are statistically significant. While we lose 16 observations, the results are essentially unchanged.

In column 3, we augmented the regression of column 2 with the standard deviation of the terms of trade (*SDTOT*) of the period 1991–2002. This variable has the right sign but is not statistically significant; the other results do not change. In column 4, we use a specification similar to the one of column 1 but substitute the default dummy with seven dummy variables aimed at tracking default history (*DEF1800* takes value 1 for countries that defaulted in the 19th century and zero otherwise; *DEF1900_50* takes value 1 for countries that defaulted over the 1900–50 period; *DEF1950_70* takes value 1 if countries that defaulted over the 1950–70 period; and so forth for the remaining four dummies). The results indicate that defaults episodes do not have a long-term impact on credit ratings. In fact, only defaults in the 1995–2002 period are significantly correlated with credit ratings over the 1999–2002 period. Panizza, Sturzenegger, and Zettelmeyer (forthcoming) use a similar model and show that the effect of default on credit rating lasts only for three years.

Next, we look at the direct impact of default on borrowing costs. Empirical studies of the effect of default on borrowing cost can be divided in three groups: (1) papers that do not find any effect of default on borrowing cost; (2) papers that find a long-lasting but small effect of defaults on

[23]Using external debt over GDP yields identical results. Our data for external debt come from the World Bank's GDF. As this data set only includes data for developing countries, we set *EXDEXP* equal to zero for industrial countries (therefore *EXDEXP* can be thought of as the following interaction $EE*(1-IND)$ where *EE* is a latent variable that contains data on external debt for industrial countries). In all our estimations we drop countries that were in default over the entire 1999–2004 period. The results are robust to keeping these countries in the sample.

borrowing costs; and (3) papers that find a temporary and rapidly decaying effect of default on borrowing cost.

The first group of papers includes work by Lindert and Morton (1989) and Chowdhry (1991) who find that countries that defaulted in the 19th century and in the 1930s did not suffer higher borrowing cost in the 1970s, and more recent work by Ades and others (2000) who find that default history had no significant effect on sovereign spreads in the late 1990s.

The second group of papers includes Eichengreen and Portes (1995) who focus on bonds issued in the 1920s and find that recent defaults were associated with an increase in spreads of approximately 20 basis points but that earlier defaults had no impact on borrowing cost, and Ozler (1993) who focuses on sovereign bank loans extended over the 1968–81 period and finds a small but statistically significant effect of default in the 1930s. While Ozler's findings suggest that default history has a long-term impact, it is worth

Table 4. Default and Credit Ratings, Cross Section Regression, 1999–2002

	(1) RATING	(2) RATING	(3) RATING	(4) RATING
LGDP_PC	1.627	1.418	1.215	1.366
	(4.69)***	(3.83)***	(3.20)***	(3.47)***
GDPGR	−1.968	−4.273	−5.324	−4.888
	(0.42)	(1.06)	(1.07)	(0.91)
LINF	−0.707	−0.817	−0.727	−0.932
	(3.48)***	(3.88)***	(3.04)***	(3.65)***
CG_BAL	14.131	6.899	8.079	9.411
	(2.61)**	(1.26)	(1.50)	(1.20)
CA_GDP	3.011	−2.800	−1.679	−1.697
	(0.64)	(0.74)	(0.40)	(0.41)
EXDEXPGDF	−0.834	−0.776	−0.750	−0.761
	(2.67)***	(3.05)***	(2.03)**	(2.13)**
IND	2.549	2.685	2.839	2.847
	(2.63)**	(2.97)***	(2.96)***	(2.66)**
DEFAULT	−1.669	−1.486	−1.855	
	(3.10)***	(2.86)***	(3.57)***	
DEBT_GDP		−0.022	−0.020	−0.020
		(2.99)***	(2.16)**	(2.73)***
OR_SIN		−1.368	−1.212	−1.143
		(2.42)**	(1.84)*	(1.56)
SDTOT			−4.102	
			(0.70)	
DEF1800				0.620
				(1.09)
DEF1900_50				−0.017
				(0.03)
DEF1950_70				0.426
				(0.56)
DEF1970_80				−0.043
				(0.06)

Table 4 *(concluded)*

	(1) RATING	(2) RATING	(3) RATING	(4) RATING
DEF1980_90				−1.049
				(1.35)
DEF1990_95				0.080
				(0.08)
DEF1995_02				−1.897
				(2.79)***
Constant	0.394	4.181	6.077	2.313
	(0.14)	(1.28)	(1.88)*	(0.73)
Observations	68	59	55	68
R-squared	0.91	0.94	0.95	0.92

Note: This table shows the results of a set of regressions where the dependent variable is average credit rating measured over the period 1999–2002 (20 corresponds to AAA, 19 to AA+, 18 to AA, and so forth, all the way down to selective default rating, SD, which is assigned a value of zero) and the explanatory variables are: the log of GDP per capita (measured in constant U.S. dollars, LGDP PC); real GDP growth (GDPGR); log inflation (LINF); central government budged balance as a share of GDP (CG BAL); current account balance as a share of GDP (CA BAL); total external debt as a share of exports (EXDEXP); a dummy variable that takes a value of one for the advanced economies and zero for all other countries (IND); a dummy variable that takes a value of one if the country defaulted in the period 1970–2002 and zero otherwise (DEFAULT); a dummy variable that takes a value of one if the country defaulted in the period 1800–1900 and zero otherwise (DEF1800); a dummy variable that takes a value of one if the country defaulted in the period 1900–1950 and zero otherwise (DEF1900 50); a dummy variable that takes a value of one if the country defaulted in the period 1950–1970 and zero otherwise (DEF1950 70); a dummy variable that takes a value of one if the country defaulted in the period 1970–1980 and zero otherwise (DEF197 80); a dummy variable that takes a value of one if the country defaulted in the period 1980–1990 and zero otherwise (DEF1980 90); a dummy variable that takes a value of one if the country defaulted in the period 1990–1995 and zero otherwise (DEF1990 95); and a dummy variable that takes a value of one if the country defaulted in the period 1995–2002 and zero otherwise (DEF1995 02). All explanatory variables (except the default dummies) are measured over the period 1999–2002. Robust *t*-statistics in parentheses. * significant at 10 percent; ** significant at 5 percent; *** significant at 1 percent.

noting that her estimates do not seem to cluster the standard errors and, back-of-the-envelope, calculations suggest that clustering would substantially reduce the explanatory power of default in the 1930s. Dell'Ariccia, Schnabel, and Zettelmeyer (2002) also find that defaults have a long-lasting effect and show that countries that participated in the Brady exchange suffered higher borrowing costs in the late 1990s. They also show that the effect of the Brady exchange on borrowing costs increased after the Russian crisis of 1998.

The third group of papers includes recent work by Flandreau and Zumer (2004) who focus on the 1880–1914 period and find that default is associated with a jump in spreads of about 90 basis points in the year that follows the end of a default episode but that the effect of default on spreads declines very rapidly over time.

Table 5 reports a set of simple regressions aimed at explaining emerging market sovereign spreads over the 1997–2004 period. We use an unbalanced panel of up to 31 countries to regress the yearly average of EMBI global spreads over a set of standard controls and a set of variables that track default history (in all regressions we drop the observations for countries that are in default in the current year). The controls include the log of GDP per capita (*LGDP_PC*), the log of inflation (*LINF*), the fiscal balance scaled by GDP (*CG_BAL*), the current account balance scaled by GDP (*CA_BAL*), and the ratio of external debt over exports (*EXDEXP*). The default variables include a dummy taking a value of one if country's *i* last default was in year $t-1$ (*DEF_1YR*), a dummy variable taking a value of one if country's *i* last default was in year $t-2$ (*DEF_2YRS*), a dummy variable taking a value of one if country's *i* last default was between year $t-3$ and year $t-5$ (*DEF3_5YRS*), a dummy variable taking a value of one if country's *i* last default was between year $t-6$ and year $t-10$ (*DEF6_10YRS*), and a dummy variable taking a value of one if country's *i* last default was between year $t-11$ and year $t-25$ (*DEF11_25YRS*). The excluded dummy is the one for countries that defaulted before year $t-25$ or never defaulted.[24]

Column 1 uses a random effects model that allows for region fixed effects and year fixed effects. We find that default in year $t-1$ has a large and statistically significant effect on spreads amounting to 400 basis points. The effect of default the following year is still sizable, 250 basis points, but not statistically significant. Longer-lasting effects are small and not statistically significant. Taken at face value, these results suggest that investors react strongly but have short memory—a result that is consistent with what Flandreau and Zumer (2004) found for the Gold Standard period. Column 2 uses a fixed effect model. As the five default dummies are collinear with the country fixed effects, we drop *DEF11_25YRS*. Hence, the results for the default dummies should be interpreted as differences with respect to countries that did not default after year $t-10$. The results are similar to those of the random effect model of column 1. Columns 3 and 4 repeat the models of column 1 and 2 but do not control for *CG_BAL* and *CA_BAL* (this allows us to include two extra countries in the sample). The results do not change significantly. In columns 5 to 8, we control for the effect of credit ratings. In columns 5 and 6 we use the residual of a rating regression that includes all the control variables (excluding default history) used in Table 4.[25] While we find that ratings have a large and statistically significant effect on spreads (a one notch change in ratings is associated with a jump in spreads of 50 basis points), our finding that default episodes have a short-lived impact on spreads does not change. In fact, we find that when we control for credit

[24]The results are essentially identical if we add a dummy variable for countries that defaulted between year t–26 and t–50.

[25]In the case of column 5 we obtain the residuals by running a random effect model and in the case of column 6 we obtain the residuals by running a fixed effects model.

Table 5. Defaults and Bond Spreads, Panel Regression, 1997–2004

	(1) EMBIG	(2) EMBIG	(3) EMBIG	(4) EMBIG	(5) EMBIG	(6) EMBIG	(7) EMBIG	(8) EMBIG
LGDP_PC	−200.578	−1424.802	−218.969	−1237.708	−216.274	−1663.319	−47.260	−1172.255
	(4.08)***	(4.97)***	(4.70)***	(4.94)***	(3.32)***	(5.53)***	(0.90)	(3.89)***
LINF	46.061	25.281	55.359	31.052	54.589	33.042	36.787	26.325
	(2.95)***	(1.55)	(3.70)***	(1.96)*	(2.97)***	(1.70)*	(2.15)**	(1.35)
CG_BALW	−446.783	635.532			−718.671	99.209	373.806	237.916
	(0.68)	(0.85)			(1.06)	(0.13)	(0.59)	(0.31)
CA_GDPW	665.056	−342.523			794.891	−454.604	465.100	−757.808
	(1.73)*	(0.81)			(1.93)*	(0.98)	(1.20)	(1.53)
EXDEXPGDF	166.770	207.386	169.660	213.708	192.966	246.435	96.262	189.341
	(5.27)***	(4.53)***	(5.71)***	(4.80)***	(5.52)***	(5.13)***	(3.24)***	(3.88)***
DEF1YEAR	412.863	307.746	433.912	305.783	389.342	249.764	267.770	249.175
	(3.39)***	(2.52)**	(3.95)***	(2.68)***	(3.04)***	(2.04)**	(2.42)**	(2.03)**
DEF2YRS	246.746	188.244	267.262	162.114	238.877	145.339	134.276	144.640
	(2.10)**	(1.63)	(2.52)**	(1.49)	(1.99)**	(1.26)	(1.33)	(1.25)
DEF3_5YRS	122.262	61.572	169.914	68.725	105.895	14.997	4.983	14.682
	(1.28)	(0.70)	(1.92)*	(0.81)	(1.07)	(0.17)	(0.06)	(0.16)
DEF6_10YRS	112.608	39.982	123.758	45.416	104.661	32.995	14.330	32.061
	(1.31)	(0.64)	(1.53)	(0.73)	(1.17)	(0.53)	(0.21)	(0.51)
DEF11_25YRS	116.623		123.956		101.621		12.180	
	(1.25)		(1.38)		(1.06)		(0.17)	
RATING_RES					−40.583	−52.359		
					(2.03)**	(2.55)**		

Table 5 *(concluded)*

RATING							−62.546 (4.99)***	−51.225 (2.49)**
Constant	1375.104 (3.60)***	11054.177 (4.92)***	1523.830 (4.19)***	9563.936 (4.87)***	1671.025 (3.38)***	13401.740 (5.56)***	1316.136 (3.74)***	9674.835 (4.18)***
Observations	150	150	162	162	144	144	144	144
Number of cc	29	29	31	31	27	27	27	27
R-squared		0.56		0.53		0.58		0.58
Region fixed effects	Yes		Yes		Yes		Yes	
Country fixed effects		Yes		Yes		Yes		Yes
Years fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Note: This table shows a set of regressions where the dependent variable measures yearly average sovereign spreads for up to 31 emerging market countries included in the JP Morgan EMBI global index. The explanatory variables are the log of GDP per capita (measured in constant U.S. dollars, LGDP PC); log inflation (LINF); central government budged balance as a share of GDP (CG_BAL); current account balance as a share of GDP (CA_BAL); total external debt as a share of exports (EXDEXP); a dummy variable taking a value of one if a country's last default was in year t−1 (DEF1YEAR); a dummy variable taking a value of one if a country's last default was in year t−2 (DEF2YRS); a dummy variable taking a value of one if a country's last default was between year t−3 and year t−5 (DEF3 5YRS); a dummy variable taking a value of one if a country's last default was between year t−6 and year t−10 (DEF6 10YRS); and a dummy variable taking a value of one if a country's last default was between year t−11 and year t−25 (DEF11 25YRS; The excluded dummy is the one for countries that defaulted before year t−25 or never defaulted); credit rating (RATING); and the residual of a rating regression that includes all the control variables (excluding default history) used in Table 4 (RATING RES). Absolute value of z-statistics in parentheses. * significant at 10 percent; ** significant at 5 percent; *** significant at 1 percent.

ratings the effect on spreads becomes even more short-lived. This is probably due to the fact that part of the reputational cost of default is reflected in lower ratings. (Panizza, Sturzenegger, and Zettelmeyer, forthcoming, use a slightly different specification and confirm this result.)

In columns 7 and 8, we substitute residual ratings with actual ratings and again find similar results.

IV. Default and International Trade

While the idea that defaults may lead to some form of trade retaliation has been around for a long time (see, for instance, Díaz-Alejandro, 1983), the empirical evidence on a link between default and trade is much more recent. Rose (2005) tests the hypothesis that defaults have a negative effect on trade by including an indicator variable for Paris Club debt renegotiations in a standard gravity trade model that uses bilateral trade data covering 200 countries over the 1948–97 period. He finds that Paris Club debt renegotiations are associated with a decline in bilateral trade that lasts for 15 years and amounts to approximately 8 percent per year.[26] In Borensztein and Panizza (2006), we use industry-level data and find that sovereign defaults are particularly costly for export-oriented industries. However, unlike Rose (2005) we find that the effect of default on exports tends to be short-lived. One question that is not addressed by either Rose (2005) or Borensztein and Panizza (2006) concerns the channel through which default affects trade.

In principle, the reduction in trade following a debt default could come from restrictive measures imposed by the country of residence of the investors. This is the assumption often made by the theoretical debt literature. However, there is little historical record of countries imposing quotas or embargos on a country that falls in default. The current structure of international capital markets, where investors are increasingly anonymous bondholders who may switch from long to short positions in minutes, makes this traditional assumption more implausible nowadays. There is, however, a more likely scenario. The deterioration in the credit quality of exporting firms after the default (that results from the risk of imposition of capital or exchange controls) could make trade credit less available and more expensive. This would, in fact, have consequences similar to those of retaliatory measures. This is precisely the idea of Kohlscheen and O'Connell (2007) who build a model of sovereign debt in which trade credit reduces the transaction costs associated with international trade and defaults are costly because they lead to a collapse in trade credit.[27] They look at 12 default episodes that took place between 1992 and 2001 and show that 11 of these episodes were followed by a collapse in trade credit much larger than the

[26] See also Martinez and Sandleris (2008).

[27] Kohlscheen and O'Connell (2007) accumulation of international reserves is justified by the fact that, by allowing countries to survive without trade credit, reserves put defaulters in a stronger bargaining position during the renegotiation of defaulted debt.

collapse in trade (the exception is Ivory Coast in 2000).[28] In the sample of Kohlscheen and O'Connell (2007), the median reduction of trade credit was 35 percent two years after the default, and 51 percent four years after the default. However, Kohlscheen and O'Connell do not provide a formal econometric test of the relationship between default and trade credit, which is what we attempt in this section.

We study the relationship between default and trade credit using Organization for Economic Development and Cooperation (OECD) data on net trade credit extended by OECD countries to developing countries and economies in transition. According to the OECD definition, trade credit measures loans for the purpose of trade which are not represented by a negotiable instrument. One problem with the OECD data set is that it only includes loans issued or guaranteed by the official sector and hence it may underestimate total trade credit. With this caveat in mind, we test the trade credit channel using and unbalanced panel to estimate the following equation:

$$NTC_{i,t} = \alpha DEFAULT_{i,t} + \beta X_{i,t} + \mu_i + \varepsilon_{i,t}, \tag{4}$$

where $NTC_{i,t}$ is net trade credit (new trade credit flows minus repayments) scaled by international trade (measured as exports plus imports) in country i in year t, $DEFAULT_{i,t}$ is a default dummy that takes a value of one if country i is in default in year t, $X_{i,t}$ is a set of controls (X includes log inflation, log GDP, the change in terms of trade, the change in the real exchange rate, a variable measuring the level of democracy, and lagged trade), and μ_i is a set of country fixed effects (we also experimented with year fixed effects and our results were unchanged).[29] We scale trade credit by trade to implicitly control for the decline in trade associated with defaults. Expressing trade credit as a share of total trade allows an interpretation of the coefficients of the regressions which is similar to the concept of elasticity. For instance, a negative value of α indicates that default episodes lead to a decrease in trade credit greater than the overall decline in trade.[30]

We start by estimating our baseline model and find that the default dummy has a negative and statistically significant effect on trade credit (column 1 of Table 6). In column 2, we explore the dynamic effect of default by augmenting the model with two dummies that take a value of one in the first and second year of the default episode (*DEF_EP* take value one in the first year of the default episode and *DEF_EP1* is a one-year lag of *DEF_EP*).

[28]There exists some evidence on the relationship between currency crisis and trade credit. Love and Zaidi (2004) and World Bank (2004) find that, in the case of East Asia, the 1997 crisis had a negative impact on trade credit, albeit smaller than that on total bank lending.

[29]In order to make sure that our results are not driven by outliers, we dropped all observations for which the dependent variable had a z-score greater than 5.

[30]In particular: $\alpha = \left(\frac{C_d}{C} - \frac{T_d}{T}\right)\frac{C}{T}$ (where C is trade credit and T trade, C_d and T_d measure the effect of default on trade and trade credit). See Love, Preve, and Sarria-Allende (2005) for a similar interpretation.

Table 6. Default and Trade Credit

	(1) NEC	(2) NEC	(3) NEC	(4) NEC	(5) NEC	(6) NEC
Estimation Method		Fixed Effects			Arellano and Bond	
DEFAULT	−0.800	−0.800	−0.800	−0.900	−0.134	0.011
	(5.85)***	(5.74)***	(5.95)***	(5.85)***	(4.88)***	(0.39)
LINF	0.000	0.000	0.000	0.000	−0.032	−0.038
	(0.14)	(0.21)	(0.11)	(0.17)	(7.51)***	(10.24)***
LGDP	0.600	0.600	0.400	0.400	0.007	−0.029
	(2.72)***	(2.72)***	(1.53)	(1.55)	(0.08)	(0.34)
DTOT	0.000	0.000	0.100	0.100	−0.599	−0.533
	(0.12)	(0.03)	(0.38)	(0.28)	(9.53)***	(7.92)***
DRER	−0.100	−0.100	−0.300	−0.300	−0.266	−0.259
	(1.74)*	(1.73)*	(2.50)**	(2.50)**	(4.78)***	(4.69)***
DEMOC	0.000	0.000	0.000	0.000	0.011	0.012
	(1.31)	(1.42)	(1.13)	(1.26)	(14.99)***	(17.32)***
DEF_EP		0.500		0.500		−0.439
		(1.75)*		(1.76)*		(8.10)***
DEF_EP1		−0.400		−0.400		−0.449
		(1.40)		(1.35)		(6.87)***
TRADE_1			0.300	0.300	0.337	0.300
			(1.31)	(1.28)	(7.20)***	(4.92)***
NEC_1					7.296	7.911
					(37.68)***	(36.16)***
Constant	−14.200	−14.200	−16.700	−16.500	−0.008	0.002
	(2.70)***	(2.70)***	(2.94)***	(2.92)***	(0.80)	(0.26)
Observations	1,060	1,060	1,059	1,059	872	872
Number of cc	99	99	99	99	96	96
R-squared	0.07	0.07	0.07	0.07		

Note: This table shows the results of a set of regressions where the dependent variable measures is net trade credit (new trade credit flows minus repayments) scaled by international trade (measured as exports plus imports) in country *i* in year *t*. The explanatory variables are a dummy variable that takes a value of one if the country is in default in year *t* (DEFAULT); log inflation (LINF); log GDP (LGDP), change in terms of trade (DTOT); change in the real exchange rate (DRER); and index of democracy (DEMOC); a dummy variable that takes a value of one in the first year of a default episode (DEF_EP); the lagged value of DEF_EP (DEF_EP1), lagged trade over GDP (TRAD_1) and lagged trade credit (NEC_1). Absolute value of *t*-statistics in parentheses. * significant at 10 percent; ** significant at 5 percent; *** significant at 1 percent.

We find that the effect of default is smaller in the first year of the default episode (this is probably due to the fact that defaults do not always happen at the beginning of the year) and larger (although the coefficient is not statistically significant) in the second year. In columns 3 and 4, we control for lagged trade and find that including this variable does not affect our baseline estimates.

There are at least two problems with the estimations of columns 1 to 4. Fist, they do not allow for persistence in the left-hand side variable. Second, they do not recognize that most variables included in the model are endogenous. Columns 5 and 6 deal with these issues by using the Arellano and Bond (1991) GMM difference estimator which allows to consistently estimate a fixed effect model that includes the lagged dependent variable. Under certain conditions, this class of general method of moments (GMM) estimators also allows to deal with endogeneity by instrumenting the explanatory variables with their lagged values. Column 5 replicates the model of column 3 adding the lagged dependent variable and using the Arellano and Bond estimator. We find that the coefficient of the default dummy remains negative and statistically significant but drops from –0.8 to –0.13. Column 6 reproduces the model of column 4 adding the lagged dependent variable and using the Arellano and Bond (1991) estimator. In this case, we find that the effect is negative and large only in the first and second year of the default. This result suggests that default does have a negative effect on trade credit but that this effect is short-lived.[31] Moreover, it is not clear whether default episodes affect trade credit more than other forms of credit. Arteta and Hale (2008) use firm-level data and suggest that this is not the case. In particular, they find that sovereign defaults reduce foreign credit access to nonexporters more than to exporters.

To probe the issue further, we run a set of regressions in which we look at whether controlling for trade credit affects the relationship between default and bilateral trade. Formally, we estimate the following gravity model:

$$LTR_{i,j,t} = \mu_{i,j} + \gamma X_{i,j,t} + \alpha DEF\_NS_{i,j,t} + \beta TC\_NS_{i,j,t} + \varepsilon_{i,j,t}, \tag{5}$$

where $LTR_{i,j,t}$ is the log of bilateral trade between country i and country j at time t, $\mu_{i,j}$ is a country pair fixed effect and $X_{i,j,t}$ is a set of controls.[32] $DEF\_NS_{i,j,t}$ is a dummy variable that takes value one if in year t either country i or country j is in default (as usual, we measure default using Standard and Poor's data) and the i, j pair consists of a developing and industrial country. This strategy, which is similar to the one used by Rose (2005) in his robustness analysis (Table 4 in Roses' paper), assumes that if there is some retaliation for default that operates through trade credit, this retaliation should mainly affect trade between high-income and low-income countries because the former are the likely creditors. $TC\_NS_{i,j,t}$ measures

[31] Note that our data for trade credit only cover suppliers of trade credit based in the OECD countries. If a default were to cause a diversion of trade towards non-OECD countries (a fact consistent with Rose's finding), we would be interpreting a change in trade credit pattern as a reduction in credit.

[32] We use the same set of controls used by Rose (2005) in his fixed effect regressions (log of total GDP, log of GDP per capita, regional trade agreement dummy, colony dummy, and currency union dummy) but also augment the regressions with a variable measuring default interacted with average trade between country i and country j.

total trade credit received by the developing country in the pair. In particular, when one of the two countries in the pair is a developing country and the other is an industrial country, $TC\_NS_{i,j,t}$ is set to be equal to the log of the stock of official trade credit received by the developing country in year t, and it takes value 0 if the i, j pair consists of either two industrial countries or two developing countries.

Although trade credit is endogenous with respect to trade, and β should not be given any causal interpretation and only interpreted as the correlation between $TC\_NS_{i,j,t}$ and $LTR_{i,j,t}$, this exercise is interesting because if it were true that the effect of default operates through trade credit, we should find that controlling for trade credit should reduce the correlation between default and trade.

In column 1 of Table 7, we reproduce the basic result of Rose (2005) and show that defaults are associated with a large and statistically significant decline in bilateral trade flows between advanced and emerging or developing economies. In column 2, we assume that country pairs with large, well-established trade relationships should be able to cope better with disruptions arising from default episodes, and control for this possibility by augmenting the regression with a variable that interacts the default dummy with the log of average trade between country i and j (*DEF_AVT*, where the average is measured using all periods for which data are available).

As expected, we find that *DEF_AVT* has a positive and statistically significant coefficient and that including this variable in the regression increases the point estimates of *DEF_NS*. In column 3, we estimate the same model of column 2 but restrict the sample to be the same to the one for which we have data on trade credit. Qualitatively, the results are unchanged. In particular, *DEF_NS* remains negative and statistically significant. Quantitatively, the impact of default is much smaller in the restricted sample.[33]

In column 4, we augment the regression with $TC\_NS_{i,j,t}$ and measure trade credit with the log of the total stock of trade credit to country i (where country i is the developing country in the pair) in year t. As expected, this variable is positive and statistically significant. It is also quantitatively important indicating that the elasticity of trade to trade credit is approximately 7 percent. While this coefficient cannot be interpreted in terms of causality, what is interesting is that controlling for trade credit does not affect the relationship between default and trade. In particular, the coefficients of *DEF_NS* and *DEF_AVT* in column 4 are identical to those of column 3. Columns 5 and 6 repeat the experiment by focusing on total nonbank trade credit and total bank trade credit. The results are basically unchanged.

[33]Running these regressions using imports as the trade measure yields less significant results (not shown here).

Table 7. Default and Trade: Does Trade Credit Matter?

	(1) LTR	(2) LTR	(3) LTR	(4) LTR	(5) LTR	(6) LTR
DEF_NS	−0.206	−0.319	−0.054	−0.054	−0.047	−0.104
	(16.46)***	(25.21)***	(1.68)*	(1.66)*	(1.47)	(3.00)***
LGDP	0.315	0.353	0.393	0.393	0.393	0.392
	(40.18)***	(45.03)***	(38.75)***	(38.76)***	(38.73)***	(38.39)***
LGDP_PC	0.323	0.262	0.145	0.144	0.145	0.149
	(27.51)***	(22.28)***	(9.75)***	(9.70)***	(9.75)***	(9.93)***
RTA	0.108	0.104	0.179	0.179	0.179	0.178
	(13.28)***	(12.86)***	(15.77)***	(15.76)***	(15.75)***	(15.62)***
CURCOL	0.332	0.388	−0.095	−0.096	−0.095	−0.091
	(3.80)***	(4.46)***	(0.38)	(0.38)	(0.38)	(0.36)
CUSTRICT	0.669	0.665	0.647	0.647	0.647	0.647
	(13.39)***	(13.38)***	(10.21)***	(10.21)***	(10.21)***	(10.15)***
DEF_AVT		0.141	0.176	0.176	0.176	0.178
		(47.37)***	(44.87)***	(44.87)***	(44.84)***	(44.88)***
LTC_TOTNS				0.073		
				(4.55)***		
LTC_NBNKNS					0.055	
					(3.59)***	
LTC_BNKNS						0.037
						(3.29)***
Constant	−10.215	−11.02	−11.338	−11.41	−11.392	−11.38
	(46.75)***	(50.53)***	(37.59)***	(37.78)***	(37.71)***	(37.39)***
Observations	234,457	234,457	151,371	151,371	151,243	147,057
Number of country pairs	12,150	12,150	11,885	11,885	11,883	11,687
R-squared	0.11	0.12	0.08	0.08	0.08	0.08

Note: This table shows the results of a set of gravity regressions where the dependent variable is the log of bilateral trade between country *i* and country *j* at time *t* (LTR) and the explanatory variables are a dummy variable that takes value one if one of the two countries is in default and the pair contains an advanced and a developing country (DEF_NS); the product of the log of GDP of the two countries (LGDP); the product of the log of GDP per capita of the two countries (LGDP_PC); a dummy variable that takes a value of one for pair of countries that are in a regional trade agreement (RTA); a dummy variable for colonies (CURCOL); a dummy variable for countries that belong to a currency union or use the same currency (CUSTRICT); average trade between the two countries (DEF_AVT); total trade credit received by the developing country in the pair (LTC_TOTNS); total nonbank trade credit received by the developing country in the pair (LTC_NBNKNS); and total bank trade credit received by the developing country in the pair (LTC_BNKNS). Absolute value of *t*-statistics in parentheses. * significant at 10 percent; ** significant at 5 percent; *** significant at 1 percent.

V. Default and the Domestic Banking System

Sovereign defaults affect not only external creditors but also domestic bondholders. Although data on the breakdown of bondholders by country of residence is scant, some recent default events suggest that domestic residents tend to account for a sizable portion of the holdings, perhaps a majority in some cases. This means that a sovereign default can have serious consequences for the domestic private sector. In particular, when domestic banks hold large amounts of government debt, the domestic financial sector may be put under significant stress by the default (Beim and Calomiris, 2000; Sturzenegger and Zettelmeyer, 2006).

Our strategy is to test if sovereign defaults lead to banking crises or a domestic credit crunch. This may happen for several reasons. First of all, default episodes may cause a collapse in confidence in the domestic financial system and may lead to bank runs, resulting in banking crises or at least a credit crunch. Second, even in the absence of a bank run, default episodes would have a negative effect on banks' balance sheet, especially if holdings of the defaulted article are large, and lead banks to adopt more conservative lending strategies. Finally, default episodes are often accompanied by a weakening of creditor rights or at least more uncertainty about them, which may also have a negative effect on bank lending.

To investigate the possible effect of sovereign defaults on banking crises, we build an index of banking crises using data from Glick and Hutchinson (2001), Caprio and Kingelbiel (2003), and Dell'Ariccia, Detragiache, and Rajan (2005).[34] Our data include 149 countries for the 1975–2000 period and 3,874 observations. In this sample, there are 111 banking crises (yielding an unconditional probability of observing a crisis of 2.9 percent) and 85 default episodes (yielding an unconditional probability of observing a default of 2.2 percent). In order to check whether defaults predict currency crisis, we compute the probability of having a banking crisis in year t conditional on having a debt default in year t or year $t-1$ (this is similar to the test in Kaminsky and Reinhart, 1999). The results indicate that the probability of having a banking crisis conditional on default is 14 percent, an 11 percentage point increase with respect to the unconditional probability (Table 8). The statistical significance of the difference between conditional and unconditional probability is quite high.

As banking crises tend to involve large fiscal costs, it is also possible that the direction of causality is reversed, namely that banking crises cause default episodes. However, the probability of a default conditional on having a banking crisis is only two percentage points higher than the unconditional probability, and the difference is not statistically significant at conventional confidence levels. These results should be taken with an appropriate degree of caution because we have relatively few cases of "twin" crisis and, as we work with annual data, we lose some precision in the measure of the relative timing of banking crises and default episodes. However, the results suggest that default episodes may increase the probability of a banking crisis much more than the other way round.

[34]We code a country-year as a banking crisis if one of the following conditions apply: either Glick and Hutchinson (1999) define the episode as a major banking crisis, or Caprio and Klingebiel (2003) define the episode as a systemic crisis, or the country year is included in the list in Dell'Ariccia, Detragiache, and Rajan (2005).

Table 8. Probabilities of Default and Banking Crisis

Unconditional probability of a banking crisis (111 episodes)	2.9
Probability of a banking crisis conditional on a default	14.1
P-value on a test P(BC/DEF) > P(BC)	0.0
Unconditional probability of a sovereign default (85 episodes)	2.2
Probability of a default conditional on a banking crisis	4.5
P-value on a test P(DEF/BC) > P(DEF)	0.1

Sources: Authors' calculations based on data from Glick and Hutchinson (2001); Caprio and Kingelbiel (2003); and Dell'Ariccia, Detragiache, and Rajan (2005).

Note: The top panel of the table shows the unconditional probability of banking crises (measured as the number of banking crises divided by number of observations in the sample) and the probability of a banking crisis conditional on a default (measured as the number of banking crises in the first or second year of a default episode divided by number of default episodes times two). The bottom panel of the table shows unconditional and conditional probabilities of a default episode (conditional and unconditional probabilities are calculated using the same methods used in the top panel of the table).

To test whether default episodes generate a credit crunch, we use a methodology similar to the one originally developed by Rajan and Zingales (1998), and recently applied by Dell'Ariccia, Detragiache, and Rajan (2005) to investigate the cost of banking crises. The basic idea is to use data at the industry level to test whether defaults have a larger negative impact on sectors that require more external finance.

Following Dell'Ariccia, Detragiache, and Rajan (2005), we pose the following specification:

$$VAGR_{i,j,t} = a_{i,j} + b_{i,t} + c_{j,t} + \alpha SHVA_{i,j,t-1} + \beta DEF_{i,t} * EXT_j + \varepsilon_{i,j,t}, \quad (6)$$

The dependent variable in equation (6) measures real value added growth for industry j in country i at time t. The controls comprise a set of country-industry fixed effects ($a_{i,j}$), a set of country-year fixed effects ($b_{i,t}$), a set of industry-year fixed effects ($c_{i,t}$), and the lagged ratio of sector j's value added over total manufacturing production ($SHVA$). Fixed effects control for country-specific, industry-specific, and time-invariant country-industry specific shocks, and hence capture most of the factors that are likely to affect the performance of a given industry and greatly attenuate omitted variable biases. $SHVA$ controls for convergence and mean reversion (possibly due to errors in variables). Our variable of interest is the interaction between a default dummy (DEF) and the index of external financial dependence (EXT) assembled by Rajan and Zingales (1998) and later used by Dell'Ariccia, Detragiache, and Rajan (2005).[35]

In the above setup, β measures whether value added growth in sectors that require more external financing is affected differentially by default

[35]Note that the definition of external-finance-dependent industries is based on data for advanced economies.

Table 9. Default and Industry Value-Added Growth

	(1) VAGR	(2) VAGR	(3) VAGR	(4) VAGR	(5) VAGR
DEF × EXT	0.009	0.021	0.009	0.019	
	(0.74)	(1.41)	(0.73)	(1.30)	
DEF_b × EXT		−0.027		−0.022	
		(1.22)		(0.99)	
DEF_b1 × EXT		−0.033		−0.031	
		(1.51)		(1.42)	
DEF_b2 × EXT		−0.013		−0.013	
		(0.60)		(0.61)	
SHVA	−1.251	−1.25	−1.253	−1.252	−1.253
	(15.18)***	(15.17)***	(15.21)***	(15.19)***	(15.21)***
BK_CR × EXT			−2.277	−2.164	−2.282
			(2.29)**	(2.16)**	(2.29)**
Constant	0.154	0.152	0.157	0.156	0.3
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Observations	15,872	15,872	15,872	15,872	15,872
R-squared	0.46	0.46	0.46	0.46	0.46

Note: This table shows the results of a set of regressions where the dependent variable is value-added growth in industry j, country i at time t. The explanatory variables are country-year fixed effects; industry-year fixed effects; country-industry fixed effects; a dummy variables that takes a value of one if the country is in default interacted with the industry-level Rajan and Zingales index of dependence on external finance (DEF × EXT); a dummy variables that takes a value of one in the first year of a default episode interacted with the industry-level Rajan and Zingales index of dependence on external finance (DEF_b × EXT; DEF_b1 and DEF_b2 are the first and second lags of DEF_b); the lagged share of the value added of industry j over total valued added in country i at time $t-1$. (SHVA); and a dummy variables that takes a value of one during banking crises interacted with the industry-level Rajan and Zingales index of dependence on external finance (BK_CR × EXT). Absolute value of t-statistics in parentheses. All regressions exclude top and bottom 5 percent observations in the dependent variable.

episodes. A negative value of β would provide evidence in support of the hypothesis that default episodes lead to a credit crunch in the banking sector.

The results of estimating this model, reported in Table 9, do not provide much support for the credit crunch hypothesis.[36] In column 1, we focus on all the years in which the country is in default (*DEF*). The coefficient has the wrong (positive) sign, although it is not statistically significant. In column 2, we use three dummy variables taking a value of one in the first, second and third year of a default episode, and find that these variables tend to have the

[36]We use the same sample restriction used in Dell'Ariccia, Detragiache, and Rajan (2005). In particular, we focus on the 1980–2000 period and restrict the sample to all the countries that observed at least a banking crisis or a default over this period. We drop from the sample the top and bottom 5 percent of observations. The last column of Table 10 uses a specification that is identical to the one used by Dell'Ariccia, Detragiache, and Rajan (2005) and obtains results which are similar (although not identical) to those obtained by those authors.

right (negative) sign but that they are never statistically significant (neither individually nor jointly). In columns 3 and 4, we augment the regressions of columns 1 and 2 with the interaction between banking crisis and external dependence (the same variable used by Dell'Ariccia, Detragiache, and Rajan, 2005) and find that our results are unchanged. We conclude that, unlike banking crises, defaults do not seem to have a special effect on industries that depend more on external finance.

VI. Political Implications of Default

Sometimes, politicians and bureaucrats seem to go to a great length to postpone what seems to be an unavoidable default. In the case of Argentina, for instance, it is reported that even Wall Street bankers had to work hard to persuade the policymaking authorities to accept reality and initiate a debt restructuring (Blustein, 2005). Why the reluctance? There seems to be evidence that defaults do not bode well for the survival in office of finance ministers and the top executive politicians.

High political costs have two important implications. On the positive side, a high political cost would increase the country's willingness to pay and hence its level of sustainable debt. On the negative side, politically costly defaults might lead to "gambles for redemption" and possibly amplify the eventual economic costs of default if the gamble does not pay off and results in larger economic costs. Delaying default might be costly for at least three reasons: (1) noncredible restrictive fiscal policies are ineffective in avoiding default and lead to output contractions; (2) delayed defaults may prolong the climate of uncertainty and high interest rates and thus have a negative effect on investment and banks' balance sheets; and (3) delayed default may have direct harmful effects on the financial sector.[37]

This suggests that a politician concerned about his/her political survival faces a tradeoff that is somewhat different from the one affecting the country itself, say, the representative citizen.

This contrast can be illustrated in a simple formal framework as follows.[38]

Assume that a country is entering a period of crisis and the policymaker needs to decide whether to default now or attempt to implement some sort of emergency program with a small chance of success. The social cost of current default is D_0. If the measures are successful (with probability $\prod$) there will be no future default (and hence no cost), but if the measures are not successful there will be a delayed default with a cost of D_1 (with $D_1 > D_0$). Hence, trying to avoid default is optimal if and only if $D_0 > (1-\prod)D_1$. This inequality can

[37]This might happen for at least two reasons. Firstly, in the attempt to avoid default, banks might be forced to increase their holdings of government bonds, which later collapse in value, and secondly, the climate of uncertainty and the weakening of the banks' financial position may trigger a deposit run.

[38]This framework is inspired in Sturzenegger and Zettelmeyer (2006, Chapter 11).

be rewritten as

$$\Pi > \frac{D_1 - D_0}{D_1},$$

implying that trying to avoid default is socially optimal only if the probability of success is greater than the percent difference between the cost of defaulting today and the cost of defaulting in the future (we assume zero discount rate).

It is now interesting to ask how self-interested politicians can lead to a deviation from the social optimum. Let us assume that the default decision is made by a policymaker who obtains a rent from being in power and that this policymaker knows that in case of default he/she will lose his/her job with probability θ. Let us assume that the policymaker's objective function is to maximize his/her own utility function, which is given by $U = (1-\Phi)R + \Phi W$, where R represents the rents from being in power, W is a measure of social welfare, and Φ $(0 \leq \Phi \leq 1)$ is the weight that the politician puts on social welfare. In this setup, the politician will decide to attempt to avoid default if:

$$(1-\Phi)R\vartheta(1+\vartheta) - \Phi D_0 < (1-\Phi)R(1+\Pi)$$

$$+ (1-\Pi)((1-\Phi)\vartheta R - \Phi D_1),$$

where $\vartheta = (1-\theta)$. This inequality can be rewritten as:

$$\Pi > \frac{D_1 - D_0}{D_1} - \frac{R(1-\Phi)}{\Phi D_1}(1 + \Pi - \vartheta(\Pi + \vartheta)). \tag{8}$$

This inequality implies that politicians who are altruistic (meaning $R = 0$, or equivalently, $\Phi = 1$) will just maximize social welfare, which is given by the first term of the right-hand side.

The same happens if defaults are not politically costly (that is when $\theta = 0$ and $\vartheta = 1$). However, in the presence of politically costly defaults, politicians who care about their own careers (that is, politicians with $R > 0$ and $\Phi < 1$) will try to delay default even when that is detrimental to social welfare. In fact, the above equation suggests that politicians who do not care about social welfare $(\Phi = 0)$ will try to postpone default even if the probability of success is zero.

There is no empirical literature on the political costs of default, but there exists a related literature on the political cost of sharp devaluations. In particular, Cooper (1971) was the first to illustrate the political cost of devaluations by showing that devaluations more than double (from 14 to 30 percent) the probability of a political crisis and a government change within the next 12 months. Frankel (2005) updates Cooper's (1971) data and finds that over the 1971–2003 period devaluations increased the probability of a change in the chief of the executive in the following 12 months by approximately 45 percent (from 20 to 29 percent).[39] Frankel (2005) also

[39]The impact of the crisis is even higher when the window is restricted to six months. In this case the probability of a change in the executive goes from 12 to 23 percent, an increase of nearly 100 percent.

checked whether devaluations affect the probability of a change of the minister of finance or governor of the central bank (whoever held the position of governor of the IMF) and found that devaluations are associated with a 63 percent increase in the probability of replacement of this official (from 36 to 58 percent).

Applying a similar methodology, we find that defaults have a broadly similar political cost. Table 10 lists all democracies that defaulted over the 1980–2003 period.[40] The table also reports the share of votes of the ruling coalition in the elections that preceded and followed the default. Of 19 countries for which we have data on electoral results before and after defaults, we find that the ruling coalitions lost votes in 18 countries (the exception is Ukraine). We also find that, on average, ruling governments in countries that defaulted observed a 16 percentage point decrease in electoral support, and that in 50 percent of the cases (11 of 22 episodes) there was a change in the chief of the executive either in the year of the default episode or in the following year. This is more than twice the probability of a change of the chief of the executive in normal times reported by Frankel (2005).

We also investigate changes in the top economic officials by looking for changes in the country's IMF governor (who is typically the finance minister but in some cases the governor of the central bank). The first column of the upper panel of Table 11 shows that in tranquil years there is a 19.4 percent probability of observing a change of the IMF governor, but after a default, the probability jumps to 26 percent (the difference is statistically significant with a *p*-value of 0.04). Interestingly, defaults on bank loans do not seem to matter (column 2) but bond defaults are particularly perilous for finance ministers. In the latter case, the probability of turnover more than doubles to over 40 percent. To check for the possibility that our results are driven by changes in political and economic institutions (for example, an increase in the ease of government turnovers), we split the sample into two subperiods. Interestingly, we do not find large differences between the two subperiods and, if anything, find that defaults seemed to have a higher political cost in the 1980s than in the 1990s. The second panel of the table uses an 18-month window to measure turnover. The results are similar to those of upper panel, but here the impact of bond defaults is even more dramatic, with more than 90 percent of finance ministers losing their job in the 18 months following a default episode (the turnover in tranquil times is 47 percent using this extended window).

In Table 12, we divide the sample according to the political regime, between dictatorships and democracies. Somewhat surprisingly, we find that the political cost of defaulting on bank loans is higher in dictatorships but the

[40]The table does not include dictatorships or countries that were transitioning towards democracy at the time of default.

Table 10. Defaults and Elections

	Year of Default	Election before Default		Election before Default		Change in Votes	Change in the Chief Executive
		Year	Votes	Year	Votes		
Argentina	2001	1999	37.50	2003	16.90	–20.60	Yes
Bolivia	1989	1985	26.42	1989	19.64	–6.78	Yes
Costa Rica	1981	1978	39.66	1982	25.79	–13.87	Yes
Costa Rica	1983	1982	45.03	1986	41.73	–3.30	
Dominican Republic	1982	1978	37.47	1982	32.85	–4.62	Yes
Ecuador	1982	1979	18.25	1984	8.31	–9.94	
Ecuador	1999	1998	18.98	2002	NA	NA	Yes
Guatemala	1989	1985	23.56	1990	8.48	–15.08	
Jamaica	1981	1980	40.67	1983	0.00	–40.67	
Jamaica	1987	1983	89.86	1989	43.32	–46.54	
Moldova	1998	1996	NA	2000	NA	NA	Yes
Paraguay	2003	1998	43.29	2003	23.88	–19.41	Yes
Peru	1980	1980	27.71	1985	5.65	–22.06	Yes
Peru	1984	1980	27.71	1985	5.65	–22.06	Yes
Trinidad and Tobago	1988	1986	NA	1992	NA	NA	
Ukraine	1998	1994	21.55	1999	25.60	4.05	
Uruguay	1987	1984	35.39	1989	25.74	–9.65	
Uruguay	1990	1989	33.03	1994	27.18	–5.85	Yes
Uruguay	2003	1999	29.30	2004	9.11	–20.19	
Venezuela	1983	1978	39.96	1983	27.85	–12.11	Yes
Venezuela	1990	1988	42.23	1993	13.69	–28.54	
Venezuela	1995	1993	13.18	1998	0.00	–13.18	

Sources: Inter-American Development Bank, *Democracies in Development*; and International Parliamentary Union: www.ipu.org/parline-e/parlinesearch.asp.

Note: The last column of the table lists all the cases in which the chief of the executive changed in the year of the default or in the year after the default. Column (f) shows the change in votes (measured in percentage points) received by the ruling part in the first election after the default episode (this is equal to (c)-(e)). Column (g) lists all the cases in which the chief of the executive changed in the year of the default or in the year after the default.

cost of defaulting on sovereign bonds is higher in democracies. When we pull all defaults together, we find a higher turnover of economic policymakers in dictatorships. This may suggest that dictators find it easier to blame and fire their minister of finance. The second panel shows that using 18-month windows does not affect the basic finding described above.

While the patterns described above are consistent with the presence of a political cost of default, defaults tend to happen in periods of economic and political turmoil and the correlation between default episodes and the dismissal of top economic officials could be driven by a common shock (for instance, a negative output shock). We are well aware of this problem (which,

Table 11. Default and the Probability of Replacing the Minister of Finance by Type of Default

Probability of Replacing the Minister of Finance	All Defaults	Defaults on International Bank Loans	Defaults on Sovereign Bonds	All Defaults 1977–89	All Defaults 1990–2004
			One year later		
Tranquil years	19.40	19.50	19.50	17.80	20.70
After a default	25.70	24.20	40.00	24.60	28.60
Difference	6.40	4.60	20.50	6.80	7.90
P-value	0.04	0.16	0.01	0.07	0.18
			18 months later		
Tranquil years	47.30	47.40	47.40	43.30	50.90
After a default	57.70	55.60	92.30	55.10	64.40
Difference	10.40	8.20	44.80	11.80	13.50
P-value	0.01	0.05	0.00	0.01	0.07

Sources: IMF, *Annual Report,* various issues; and authors' calculations.

Note: All the *p*-values refer to a two-tails test. This table shows the probability of replacing the Minister of Finance in tranquil years and in years that follow a default episode.

in fact, is also present in the analyses of Cooper, 1971; and Frankel, 2005) and thus we make no claim of causality.[41]

The previous discussion emphasized that the political cost of default may decrease social welfare because it generates incentives to gamble for redemption. However, the presence of political costs may also have a positive effect on social welfare because, like all other costs of default, it provides incentives to repay and thus increases the level of sustainable debt. This is exactly the channel emphasized by the political science literature on "leader-specific punishment." In particular, McGillivray and Smith (2000, 2004, 2005, 2006) show that, if leaders are replaceable, punishment targeted against individual politicians increases the probability of international cooperation. In this setup, defaulters may be punished only as long as the leader that defaulted remains in power. Thus, from the country's point of view, it would be optimal to default and immediately replace the leader. Knowing this, the leader will do his or her best not to default. Rose and Spiegel (2007) provide some evidence that countries willing to cooperate on noneconomic relationships (such as environmental treaties) are also more likely to cooperate on economic relationships. There are cases, however, in which the domestic-audience cost may have exactly the opposite effect and provide politicians with incentives to repudiate external debt. For instance,

[41]This is why we did not attempt to estimate a formal econometric model but just showed that the well-known result that sharp devaluations may be politically costly also applies to episodes of sovereign default.

Table 12. Default and the Probability of Replacing the Minister of Finance by Type of Default and Government

Probability of Replacing the Minister of Finance	Defaults on International Bank Loans	Defaults on International Bank Loans	Defaults on Sovereign Bonds	Defaults on Sovereign Bonds	All Defaults	All Defaults
	Democracies	Dictatorships	Democracies	Dictatorships	Democracies	Dictatorships
			One year later			
Tranquil years	21.90	17.60	21.70	18.00	21.80	17.50
After a default	23.30	27.00	44.40	33.30	24.70	28.70
Difference	1.40	9.40	22.70	15.30	2.80	11.20
P-value	0.79	0.04	0.02	0.17	0.54	0.01
			18 months later			
Tranquil years	51.10	44.90	50.60	45.50	50.80	44.80
After a default	53.50	60.80	94.40	87.50	57.00	61.50
Difference	2.40	15.90	43.80	42.00	6.20	16.70
P-value	0.69	0.00	0.00	0.00	0.29	0.00

Sources: IMF, *Annual Report,* various issues; Alvarez and others (1999); and authors' calculations.

Note: This table shows the probability of replacing the Minister of Finance in tranquil years and in years that follow a default episode. The data are organized by type of government (democracies versus dictatorships). All the *p*-values refer to a two-tails test.

Tomz (2002) argues that the domestic-audience effect prevented Argentina from defaulting in 1999 but had the opposite effect in 2001. It is also plausible that the recent partial default of Ecuador was motivated to some extent by domestic-audience pressure.

VII. Conclusions

This paper has investigated the empirical basis of the costs of sovereign defaults in its different versions. Table 13 provides a synthesis of the results presented in a way which may be useful to authors who need evidence on the various costs of default in order to calibrate theoretical models. The findings suggest that default costs are significant, but short-lived. Reputation of sovereign borrowers that fall in default, as measured by credit ratings and spreads, is tainted, but only for a short time. While there is some evidence that international trade and trade credit are negatively affected by episodes of default, we could not trace it to the volume of trade credit, as the default literature suggests. Debt defaults seem to cause banking crises, and not vice-versa, but we found weak evidence to suggest the presence of default-driven credit crunches in domestic markets. Finally, defaults seem to shorten the life expectancy of governments and officials in charge of the economy in a significant way.

The results suggest that default costs remain somewhat vaguely defined, and difficult to quantify. On the positive side, we found a fairly sensible estimate of the effect on credit ratings and bond spreads, and we call attention to the sharp increase in government turnovers following debt crises. On the negative side, the result regarding how international trade credit affects the link between trade and default and the finding that default episodes do not appear to affect bank lending do not seem to be very plausible. Perhaps the most robust and striking finding is that the effect of defaults is short-lived, as we almost never can detect effects beyond one or two years.

A relatively unexplored avenue is the decision-making process of policymakers concerning the timing of defaults (see, however, Alichi, 2008). Defaults tend to be widely anticipated and happen at times when the domestic economy is quite weak. This may happen for two widely different reasons. Self-interested policymakers may try postponing defaults even at increasing economic cost, as the evidence presented in this paper suggests a clearly higher political turnover following a debt default. A different possibility is that policymakers postpone default to ensure that there is broad market consensus that the decision is unavoidable and not strategic. This would be in line with the model in Grossman and Van Huyck (1988) whereby "strategic" defaults are very costly in terms of reputation—and that is why they are never observed in practice—while "unavoidable" defaults carry limited reputation loss in the markets. Hence, choosing the lesser of the two evils, policymakers would postpone the inevitable default decision in order to avoid a higher reputational cost, even at a higher economic cost during the delay.

Table 13. Summary of Evidence

Effect of Default on:	Immediate Effect	Long-Run Effect
GDP growth	• Negative effect ranging between 0.6 and 2.5 percentage points • If quarterly data are used no statistically significant effect in the quarter after the default	No statistically significant effect after the year in which the default takes place
Exclusion from capital market	Almost full exclusion	There is no permanent exclusion. Access is regained between four and eight years after the default.
Credit rating	Negative effect of about one notch in the three years after the default	No statistically significant effect after three years
Borrowing cost	• 250 to 400-basis-point increase in the two years after the default • This includes the effect of credit rating downgrade (the effect on spread is smaller and often not statistically significant if credit ratings are included in the regression)	No statistically significant effect after two years
Trade	• Net decrease of bilateral trade (about 8 percent) • Negative effect on export-oriented industries	• The negative effect on bilateral trade lasts for approximately 15 years. • The effect on export-oriented industries lasts for two to three years
Trade credit	Decline of 0.5 percentage points in the year of the default and the year after the default	Some evidence of an effect up to four years after the default
Banking crises	An increase in the probability of a banking crisis of approximately 11 percentage points	
Politicians and policymakers	• A 16 percent decrease in support of the ruling party in the first election after a default. • A 50 percent increase in the probability of replacing the head of the executive • A 33 percent increase in the probability of replacing the minister of finance or the head of the central bank.	

Note: This table is based on the various articles surveyed in this paper and on the various econometric experiments included in the paper.

APPENDIX I

Table A1. Private Lending to Sovereign: Default and Rescheduling

		Standard and Poor's (1824–2004)				Beim and Calomiris (1800–1992)		Sturzenegger and Zettelmeyer (1874–2003)	Detragiache and Spilimbergo (1973–91)
		Foreign currency bond debt		Foreign currency bank debt					
Region	Country	Beginning of period	End of period	Beginning of period	End of period	Beginning of period	End of period	Beginning of period	Beginning of period
Africa	Algeria			1991	1996				1991
Africa	Angola			1985	2004	1988	1992	1988	
Africa	Burkina Faso			1983	1996				1982
Africa	Burundi								1986
Africa	Cameroon								1979
Africa	Cameroon			1985	2003	1989	1992	1989	1985
Africa	Cape Verde			1981	1996				
Africa	Central African Rep.			1981					
Africa	Central African Rep.			1983	2004				
Africa	Congo			1983	2004	1986	1992	1986	
Africa	Congo, Dem. Rep.					1961			
Africa	Congo, Dem. Rep.			1976	2004	1976	1992	1976	1975
Africa	Côte d'Ivoire			1983	1998	1984	1992	1984	1987
Africa	Côte d'Ivoire	2000	2004						
Africa	Egypt	1876	1880			1816	1880	1876	
Africa	Egypt					1984	1992	1984	1986
Africa	Ethiopia			1991	1999				1987
Africa	Gabon					1978			

Africa	Gabon			1986	1994	1986	1992	1986	
Africa	Gabon			1999	2004				
Africa	Gambia			1986	1990	1986	1988	1986	
Africa	Ghana					1969	1974		
Africa	Ghana			1987					
Africa	Guinea			1986	1988				
Africa	Guinea			1991	1998	1985	1992		
Africa	Guinea-Bissau			1983	1996				
Africa	Kenya			1994	2004				1990
Africa	Lesotho								1990
Africa	Liberia	1875	1898			1875	1898	1874	
Africa	Liberia	1912							
Africa	Liberia	1914	1915						
Africa	Liberia	1917	1918						
Africa	Liberia	1919	1923			1912	1923	1912	
Africa	Liberia	1932	1935			1932	1935		
Africa	Liberia			1987	2004	1980	1992	1980	
Africa	Madagascar			1981	2002	1981	1992	1981	1980
Africa	Malawi			1982		1982	1988	1982	
Africa	Malawi			1988					1987
Africa	Mauritania			1992	1996				
Africa	Morocco	1903	1904			1903	1904		
Africa	Morocco			1983					
Africa	Morocco			1986	1990	1983	1990	1983	1985
Africa	Mozambique			1980					
Africa	Mozambique			1983	1992	1984	1992	1984	
Africa	Niger			1983	1991	1983	1991	1983	1984
Africa	Nigeria								1972
Africa	Nigeria	1986	1988						
Africa	Nigeria	1992		1982	1992	1983	1991	1983	1986
Africa	Nigeria	2002							
Africa	São Tomé and Principe			1987	1994				
Africa	Senegal			1981	1985	1981	1992	1981	1984

Table A1 *(continued)*

Region	Country	Standard and Poor's (1824–2004)				Beim and Calomiris (1800–1992)		Sturzenegger and Zettelmeyer (1874–2003)	Detragiache and Spilimbergo (1973–91)
		Foreign currency bond debt		Foreign currency bank debt					
		Beginning of period	End of period	Beginning of period	End of period	Beginning of period	End of period	Beginning of period	Beginning of period
Africa	Senegal			1990					1989
Africa	Senegal			1992	1996				
Africa	Seychelles			2000	2002				
Africa	Sierra Leone								1972
Africa	Sierra Leone			1983	1984				
Africa	Sierra Leone			1986	1995	1977	1992	1977	
Africa	South Africa			1985	1987	1985	1992	1985	
Africa	South Africa			1989					
Africa	South Africa			1993					
Africa	Sudan			1979	2004	1979	1992	1979	1976
Africa	Tanzania			1984	2004	1984	1992	1984	
Africa	Togo			1979	1980				
Africa	Togo			1982	1984				
Africa	Togo			1988					
Africa	Togo			1991	1997	1979	1992	1979	
Africa	Tunisia	1867	1870			1867	1870		
Africa	Tunisia								1991
Africa	Uganda			1980	1993	1981	1992	1981	
Africa	Zambia								1978
Africa	Zambia			1983	1994	1983	1992	1983	
Africa	Zimbabwe	1965	1980			1965	1980		
Africa	Zimbabwe			2000	2004				

Asia	Bangladesh								1978
Asia	Bangladesh								1991
Asia	China	1921	1936						
Asia	China	1939	1949			1921	1949		
Asia	Indonesia			1998	1999				1998
Asia	Indonesia			2000					
Asia	Indonesia			2002					
Asia	Iran			1978	1995	1992			
Asia	Iraq			1987	2004	1990	1992		
Asia	Japan	1942	1952			1942	1952		
Asia	Jordan			1989	1993	1989	1992	1989	
Asia	Korea								1998
Asia	Korea, Dem. Rep.			1974	2004				
Asia	Myanmar			1997	2004				
Asia	Pakistan	1999		1998	1999			1981	
Asia	Philippines			1983	1992	1983	1992	1983	1984
Asia	Sri Lanka								1992
Asia	Thailand								1998
Asia	Turkey	1876	1881			1876	1881	1876	
Asia	Turkey	1915	1928			1915	1932	1915	
Asia	Turkey	1931	1932						
Asia	Turkey	1940	1943			1940	1943	1940	
Asia	Turkey					1959			
Asia	Turkey					1965			
Asia	Turkey			1978	1979	1978	1982	1978	
Asia	Turkey			1982					
Asia	Vietnam			1985	1998	1985	1992	1985	1984
Asia	Yemen			1985	2001				
Europe	Albania			1991	1995	1990	1992		
Europe	Austria					1802	1816		
Europe	Austria	1868	1870			1868	1870	1868	
Europe	Austria	1914	1915			1914	1915	1914	
Europe	Austria	1932	1933			1932	1952	1932	
Europe	Austria	1938							

Table A1 *(continued)*

		Standard and Poor's (1824–2004)				Beim and Calomiris (1800–1992)		Sturzenegger and Zettelmeyer (1874–2003)	Detragiache and Spilimbergo (1973–91)
		Foreign currency bond debt		Foreign currency bank debt					
Region	Country	Beginning of period	End of period	Beginning of period	End of period	Beginning of period	End of period	Beginning of period	Beginning of period
Europe	Austria	1940	1952						
Europe	Bosnia and Herzegovina			1992	1997				
Europe	Bulgaria	1916	1920			1915	1920	1915	
Europe	Bulgaria	1932				1932	1992	1932	
Europe	Bulgaria			1990	1994	1990	1992		
Europe	Croatia			1992	1996				
Europe	Czechoslovakia	1938	1946			1938	1946		
Europe	Czechoslovakia	1959	1960			1952	1959		
Europe	Germany	1932	1938					1932	
Europe	Germany	1939	1953			1932	1953		
Europe	Germany East					1949	1992		
Europe	Greece	1826	1878			1826	1878	1824	
Europe	Greece	1894	1897			1894	1897	1893	
Europe	Greece	1932	1964			1932	1964		
Europe	Hungary	1932	1937			1932	1967		
Europe	Hungary	1941	1967					1931	
Europe	Italy	1940	1946			1940	1946	1940	
Europe	Macedonia			1992	1997				
Europe	Moldova	1998						2002	
Europe	Moldova	2002							

Europe	Netherlands					1802	1814		
Europe	Poland	1936	1937						
Europe	Poland	1940	1952			1936	1952	1936	
Europe	Poland			1981	1994	1981	1992	1981	
Europe	Portugal	1837	1841			1834	1841	1834	
Europe	Portugal	1850	1856			1850	1856		
Europe	Portugal	1892	1901			1892	1901	1892	
Europe	Romania							1915	
Europe	Romania	1933	1958			1933	1958	1933	
Europe	Romania			1981	1983	1982	1987	1981	
Europe	Romania			1986					
Europe	Russia/USSR					1839		1839	
Europe	Russia/USSR					1885			
Europe	Russia/USSR	1918				1917	1918	1917	
Europe	Russia/USSR			1991	1997	1991	1992		
Europe	Russia/USSR	1998	2000					1998	
Europe	Serbia and Montenegro			1992	2004				
Europe	Slovenia			1992	1996				
Europe	Spain	1824	1834			1820		1820	
Europe	Spain					1831	1834	1831	
Europe	Spain					1851			
Europe	Spain	1837	1867			1867	1872	1867	
Europe	Spain	1827	1882			1882		1882	
Europe	Ukraine	1998	2000					1998	
Europe	Yugoslavia	1895				1895		1895	
Europe	Yugoslavia	1933	1950			1933	1960	1933	
Europe	Yugoslavia	1992		1983	1991	1983	1992	1983	
LAC	Antigua and Barbuda			1996	2004				
LAC	Argentina	1828	1857					1830	
LAC	Argentina	1890	1893			1890	1893	1890	
LAC	Argentina					1956	1965		
LAC	Argentina	1989		1982	1993	1982	1992	1982	1983

Table A1 *(continued)*

		Standard and Poor's (1824–2004)				Beim and Calomiris (1800–1992)		Sturzenegger and Zettelmeyer (1874–2003)	Detragiache and Spilimbergo (1973–91)
		Foreign currency bond debt		Foreign currency bank debt					
Region	Country	Beginning of period	End of period	Beginning of period	End of period	Beginning of period	End of period	Beginning of period	Beginning of period
LAC	Argentina	2001	2004	2001	2004			2001	
LAC	Bolivia	1875	1879			1875	1879	1874	
LAC	Bolivia	1931	1948			1931	1957	1931	
LAC	Bolivia			1980	1984	1980	1992	1980	
LAC	Bolivia	1989	1997	1986	1993				
LAC	Brazil	1826	1829			1826	1829	1826	
LAC	Brazil	1898	1901			1898	1910	1898	
LAC	Brazil	1902	1910						
LAC	Brazil	1914	1919			1914	1919	1914	
LAC	Brazil	1931	1933					1931	
LAC	Brazil	1937	1943			1931	1943		
LAC	Brazil					1961	1964		
LAC	Brazil			1983	1994	1983	1992	1983	
LAC	Chile	1826	1842			1826	1842	1826	
LAC	Chile	1880	1883			1880	1883	1879	
LAC	Chile	1931	1947			1931	1948	1931	
LAC	Chile					1965			
LAC	Chile					1972	1975		1973
LAC	Chile			1983	1990	1983	1990	1983	
LAC	Colombia	1826	1845						
LAC	Colombia	1850	1861			1826	1861	1826	
LAC	Colombia	1873				1873			

LAC	Colombia	1880	1896						
LAC	Colombia	1900	1904			1880	1904	1879	
LAC	Colombia	1932	1934					1900	
LAC	Colombia	1935	1944			1932	1944	1932	
LAC	Colombia								1985
LAC	Costa Rica	1828	1840			1828	1840	1827	
LAC	Costa Rica	1874	1885			1874	1885	1874	
LAC	Costa Rica	1895	1897						
LAC	Costa Rica	1901	1911			1895	1911	1895	
LAC	Costa Rica	1932	1952			1932	1953	1937	
LAC	Costa Rica	1962							
LAC	Costa Rica			1981					
LAC	Costa Rica	1984	1985	1983	1990	1981	1990		
LAC	Cuba	1933	1934			1933	1934	1933	
LAC	Cuba	1960				1960	1963		
LAC	Cuba			1982	2004	1982	1992	1982	
LAC	Dominica			2003	2004			2003	
LAC	Dominican Rep.							1869	
LAC	Dominican Rep.	1872	1888						
LAC	Dominican Rep.	1892	1893						
LAC	Dominican Rep.	1897							
LAC	Dominican Rep.	1899	1907			1872	1907	1899	
LAC	Dominican Rep.	1931	1934			1931	1934	1931	
LAC	Dominican Rep.								1976
LAC	Dominican Rep.			1982	1994	1982	1992	1982	
LAC	Ecuador	1826	1855			1832	1855	1832	
LAC	Ecuador	1868	1890						
LAC	Ecuador	1894	1898			1868	1898	1868	
LAC	Ecuador	1906	1908			1906	1955		
LAC	Ecuador	1909	1911					1911	
LAC	Ecuador	1914	1924					1914	
LAC	Ecuador	1929	1954					1931	
LAC	Ecuador			1982	1995	1982	1992	1982	1983
LAC	Ecuador	1999	2000					1999	

Table A1 *(continued)*

		Standard and Poor's (1824–2004)				Beim and Calomiris (1800–1992)		Sturzenegger and Zettelmeyer (1874–2003)	Detragiache and Spilimbergo (1973–91)
		Foreign currency bond debt		Foreign currency bank debt					
Region	Country	Beginning of period	End of period	Beginning of period	End of period	Beginning of period	End of period	Beginning of period	Beginning of period
LAC	El Salvador	1828	1860			1828	1860	1827	
LAC	El Salvador	1898							
LAC	El Salvador	1921	1922			1921	1922	1921	
LAC	El Salvador	1932	1935						
LAC	El Salvador	1938	1946			1932	1946	1931	
LAC	El Salvador								1984
LAC	El Salvador								1995
LAC	Guatemala	1828	1856			1828	1856	1828	
LAC	Guatemala	1876	1888			1876	1888	1876	
LAC	Guatemala	1894						1894	
LAC	Guatemala	1899	1913			1894	1917		
LAC	Guatemala	1933	1936			1933	1936	1933	
LAC	Guatemala	1989							1985
LAC	Guyana			1979					
LAC	Guyana			1982	2004	1982	1992		
LAC	Haiti			1982	1994				1983
LAC	Honduras	1828	1867			1828	1867	1827	
LAC	Honduras	1873	1925			1873	1925	1873	
LAC	Honduras							1914	
LAC	Honduras								1976
LAC	Honduras			1981	2004	1981	1992	1981	1982

LAC	Jamaica			1978	1979	1978	1990		
LAC	Jamaica			1981	1985				
LAC	Jamaica			1987	1993				
LAC	Mexico	1828	1830					1827	
LAC	Mexico	1833	1841						
LAC	Mexico	1844	1850			1828	1850		
LAC	Mexico	1854	1864						
LAC	Mexico	1866	1885			1859	1885	1867	
LAC	Mexico	1914	1922			1914	1922	1914	
LAC	Mexico	1928	1942			1928	1942		
LAC	Mexico			1982	1990	1982	1990	1982	
LAC	Nicaragua	1828	1874			1828	1874	1828	
LAC	Nicaragua	1894	1895			1894	1895	1894	
LAC	Nicaragua	1911	1912					1911	
LAC	Nicaragua	1915	1917			1911	1917		
LAC	Nicaragua	1932	1937			1932	1937	1932	
LAC	Nicaragua			1979	2004	1980	1992	1980	1978
LAC	Panama	1932	1946			1932	1946	1932	
LAC	Panama	1987	1994	1983	1996	1983	1992	1982	
LAC	Panama								1987
LAC	Paraguay							1827	
LAC	Paraguay	1874	1885			1874	1885	1874	
LAC	Paraguay	1892	1895			1892	1895		
LAC	Paraguay	1920	1924			1920	1924	1920	
LAC	Paraguay	1932	1944			1932	1944	1932	
LAC	Paraguay			1986	1992	1986	1992	1986	1984
LAC	Paraguay	2003	2004						
LAC	Peru	1826	1848			1826	1848	1826	
LAC	Peru	1876	1889			1876	1889	1876	
LAC	Peru	1931	1951			1931	1951	1931	
LAC	Peru					1968	1969		
LAC	Peru			1976					
LAC	Peru			1978					
LAC	Peru			1980					

Table A1. *(concluded)*

		Standard and Poor's (1824–2004)				Beim and Calomiris (1800–1992)		Sturzenegger and Zettelmeyer (1874–2003)	Detragiache and Spilimbergo (1973–91)
		Foreign currency bond debt		Foreign currency bank debt					
Region	Country	Beginning of period	End of period	Beginning of period	End of period	Beginning of period	End of period	Beginning of period	Beginning of period
LAC	Peru			1984	1997	1978	1992	1978	
LAC	Peru							1983	1983
LAC	Trinidad and Tobago			1988	1989	1989	1989		1988
LAC	Uruguay	1876	1878			1876	1878	1876	
LAC	Uruguay	1891				1891		1891	
LAC	Uruguay	1915	1921			1915	1921	1915	
LAC	Uruguay	1933	1938			1933	1938	1933	
LAC	Uruguay			1983	1985				
LAC	Uruguay			1987					
LAC	Uruguay			1990	1991	1983	1991	1983	
LAC	Uruguay	2003						2003	
LAC	Venezuela	1826	1840			1832	1840	1832	
LAC	Venezuela	1848	1859					1847	
LAC	Venezuela	1860	1862						

LAC	Venezuela	1865	1881			1848	1881	1864	
LAC	Venezuela	1892				1892		1892	
LAC	Venezuela	1898	1905			1898	1905		
LAC	Venezuela			1983	1988	1982	1990	1982	1984
LAC	Venezuela			1990					
LAC	Venezuela	1995	1997						

Note: *Standard and Poor's* generally defines sovereign default as the failure to meet a principal or interest payment on the due date (or within the specified grace period) contained in the original terms of the debt issue. In particular, each issuer's debt is considered in default in any of the following circumstances: (1) for local and foreign currency bonds, notes and bills, when either scheduled debt service is not paid on the due date, or an exchange offer of new debt contains terms less favorable than the original issue; (2) for central bank currency, when notes are converted into new currency of less than equivalent face value; and (3) for bank loans, when either scheduled debt service is not paid on the due date, or a rescheduling of principal and/or interest is agreed to by creditors at less favorable terms then the original loan. Such rescheduling agreements covering short- and long-term debt are considered defaults even where, for legal or regulatory reasons, creditors deem forced rollover of principal to be voluntary.

Beim and Calomiris only includes private lending through bonds, supplier's credits, or banks loans. The data set does not include every instance of technical default on bond or loan covenants. An extended period (six months or more) was identified where all or part of interest and/or principal payments due were reduced or rescheduled. Some of the defaults and rescheduling involved outright repudiation (a legislative or executive act of government liability), while others were minor and announced ahead of time in a conciliatory fashion by debtor nations. The end of each period of default or rescheduling was recorded when full payments resumed or restructuring was agreed upon. Periods of default or rescheduling within five years of each other were combined. Where a formal repudiation was identified, its date served as the end of the period of default and the repudiation is noted in notes, where no clear repudiation was announced the default was listed as persisting. Voluntary refinancing (Colombia, 1985, and Algeria, 1992) were not included.

Sturzenegger and Zettelmeyer: Unless otherwise noted, all defaults are federal or central government defaults. Defaults of U.S. southern states in early 1840s are not shown in the table. Defaults on wars, revolutions, occupations and the collapse of the Soviet Union etc. are excluded, except when they coincide with a cluster. In the event of sequence rescheduling, the year listed refers to the initial default or rescheduling.

Detragiache and Spilimbergo: An observation is classified as a debt crises if either or both of the following conditions occur: (1) there are arrears of principal or interest on external obligations towards commercial creditors (banks or bondholders) of more than 5 percent of total commercial debt outstanding; and (2) there is a rescheduling or debt restructuring agreement with commercial creditors as listed in the GDF.

APPENDIX II

Table A2. Logit model for the probability of default

	(1)	(2)
	DEF	DEF_B
Total debt to GDP_1 × Dummy 70s	1.973	6.058
	(0.87)	(2.89)
Total debt to GDP_1 × Dummy 80s	0.081	−5.920
	(0.03)	(2.60)
Total debt to GDP_1 × Dummy 90s	−0.583	−5.935
	(0.26)	(2.75)
Short term debt_1 × Dummy 70s	−1.381	0.751
	(0.33)	(0.29)
Short term debt_1 × Dummy 80s	0.855	−1.448
	(0.20)	(0.55)
Short term debt_1 × Dummy 90s	1.287	−1.178
	(0.30)	(0.43)
Short term interest payments to GDP_1 × Dummy 70s	5.552	3.337
	(2.31)	(1.95)
Short term interest payments to GDP_1 × Dummy 80s	−5.441	−2.693
	(2.25)	(1.54)
Short term interest payments to GDP_1 × Dummy 90s	−4.418	−1.975
	(1.78)	(1.14)
External debt service to reserves_1 × Dummy 70s	4.843	5.154
	(4.42)	(3.20)
External debt service to reserves_1 × Dummy 80s	−1.234	−6.162
	(0.50)	(3.52)
External debt service to reserves_1 × Dummy 90s	−2.268	−5.778
	(1.22)	(3.52)
Current account balance to GDP_1 × Dummy 70s	0.409	0.210
	(0.49)	(0.20)
Current account balance to GDP_1 × Dummy 80s	−0.286	−0.460
	(0.36)	(0.40)
Current account balance to GDP_1 × Dummy 90s	0.158	0.033
	(0.19)	(0.03)
Exports plus imports to GDP_1 × Dummy 70s	0.115	−0.267
	(0.32)	(0.64)
Exports plus imports to GDP_1 × Dummy 80s	−0.396	0.431
	(1.01)	(1.02)
Exports plus imports to GDP_1 × Dummy 90s	−0.182	0.257
	(0.48)	(0.57)
Concessional debt to total debt_1 × Dummy 70s	−5.356	−4.456
	(1.85)	(1.56)
Concessional debt to total debt_1 × Dummy 80s	4.166	3.482
	(1.39)	(1.21)
Concessional debt to total debt_1 × Dummy 90s	6.343	4.413
	(2.09)	(1.39)
US real treasury bill rate_1	0.184	0.127
	(2.13)	(1.29)
Real GDP growth_1	−0.063	−0.066
	(3.91)	(3.36)

Table A2 *(concluded)*

	(1) *DEF*	(2) DEF_B
Volatility on inflation_1	0.093	0.017
	(3.16)	(0.76)
Inflation higher than 50 percent_1	0.164	−0.258
	(0.49)	(0.68)
Year of presidential election_1	−0.006	−0.553
	(0.03)	(1.34)
Civil Liberties Index_1	−1.132	−0.028
	(1.22)	(0.03)
US real treasury bill rate_1 × Total debt to GDP_1	−0.101	−0.187
	(0.76)	(1.19)
US real treasury bill rate_1 × Short term debt_1	0.049	0.276
	(0.38)	(1.55)
US real treasury bill rate_1 × Short term interes payments to GDP_1	0.001	−0.126
	(0.01)	(1.62)
US real treasury bill rate_1 × External debt service to reserves_1	−0.558	0.192
	(1.54)	(1.47)
US real treasury bill rate_1 × Current account balance to GDP_1	−0.047	−0.076
	(0.92)	(1.12)
Dummy 80s	1.868	1.667
	(0.85)	(1.29)
Dummy 90s	−0.140	0.600
	(0.06)	(0.37)
Constant	−2.290	−4.532
	(1.05)	(3.12)
Observations	1416	1416
R-square	0.313	0.147

Note: The numbers between brackets are t-statistics.

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On the Probabilistic Approach to Fiscal Sustainability: Structural Breaks and Non-Normality

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This paper modifies several assumptions in the probabilistic approach to fiscal sustainability proposed by Celasun, Debrun, and Ostry (2007). First, we allow for structural breaks in the vector autoregression model for the macroeconomic variables. Second, in the Monte-Carlo simulations, we draw directly from the empirical distribution of the shocks instead of drawing from a normal distribution, thus allowing for asymmetries and thick tails. Third, we circumvent the use of a fiscal reaction function by focusing attention instead on debt-stabilizing balances, to produce more "agnostic" debt projections. The paper illustrates how these methodological modifications have significant impacts on the results for specific country cases. [JEL C11, O47]
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The joint World Bank-International Monetary Fund (IMF) debt-sustainability framework (DSF) produces baseline projections for debt dynamics under different macroeconomic scenarios, and conducts stress tests specifying large changes in the variables involved, keeping all else constant.

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These shocks are applied, in turn, to interest rates, exchange rates, or growth—without changing the values of the remaining endogenous variables. Celasun, Debrun, and Ostry (2007; CDO henceforth) discuss the limitations of this approach and propose a probabilistic scenario analysis. This paper proposes further modifications to the CDO framework.

Using a stochastic framework, CDO draw constellations of values for the variables driving the debt dynamics. For this purpose, CDO estimate a vector autoregression (VAR) model for the nonfiscal components of the system. Under the assumption of joint normality, they create fan charts for the debt-to-GDP ratio. These fan charts are produced by drawing from the fitted residual multivariate normal distribution while iterating forward the law of motion of the debt-to-GDP ratio—along an estimated fiscal reaction function. Among others, Budina and van Wijnbergen (2009), Tanner and Samake (2008), and Bandiera and others (2007) apply this framework to several country cases, and discuss how it improves on the standard World Bank-IMF DSF.

This paper builds on this work, and explores some methodological modifications to the CDO debt sustainability analysis, motivated by observed empirical regularities of the variables involved. First, in the VAR specification, we allow for structural breaks in the data-generation mechanism, through the application of Markov-Switching models. Second, in the Monte-Carlo simulations, the assumption of normally distributed shocks is relaxed and bootstrapping techniques are used to draw directly from their empirical distributions. This allows for a better risk assessment as it accounts for thicker tails and asymmetries in the debt projections. Third, the estimation of the fiscal reaction function is avoided by using instead the debt-stabilizing balance each period when producing baseline projections.

This specification is *not* intended to predict fiscal behavior, but rather to serve as an agnostic and sensible reference when assessing debt sustainability. It addresses the question of what would be the predictive density for the debt-to-GDP ratios if (1) the government's balance at each period were precisely the debt-stabilizing balance while (2) interest rates, GDP growth, and the exchange rate exhibited joint behavior similar to past experience—possibly differentiating between normal (tranquil) and turbulent (nontranquil) times. However, we do not take this as a prediction of fiscal behavior, but rather as a reference. Moreover, we avoid here predicting whether the times ahead for each particular country are likely to be tranquil or nontranquil, and we leave it to the analyst to choose her weights when averaging the two scenarios.

I. Debt Sustainability Framework

Using d for debt and b for the government balance (expressed as percentages of GDP), g for real GDP growth, α^f for the share of foreign-denominated government debt, and β^f as the share of the tradable sector in GDP, the debt dynamics are given by

$$d_t = \frac{1+i_t}{(1+g_t)(1+\pi_t)} d_{t-1} - b_t \tag{1}$$

$$i_t = \alpha^h i_t^h + \alpha^f [i_t^f + e_t(1 + i_t^f)] \tag{2}$$

$$\pi_t = \beta^h \pi_t^h + \beta^f [\pi_t^f + e_t \beta^f (1 + \pi_t^f)], \tag{3}$$

where e_t is the rate of nominal depreciation of the local currency, i_t is the nominal interest rate, and π_t is GDP deflator (Ley, 2009). Equations (1)–(3) can be combined to obtain the familiar expression driving public-debt dynamics

$$\begin{aligned} d_t &= \frac{1+i_t}{(1+g_t)(1+\pi_t)} d_{t-1} - b_t \\ &= \frac{1 + \alpha^h i_t^h + \alpha^f [i_t^f + e_t(1+i_t^f)]}{(1+g_t)\{1 + \beta^h \pi_t^h + \beta^f [\pi_t^f + e_t \beta^f (1+\pi_t^f)]\}} d_{t-1} - b_t. \end{aligned} \tag{4}$$

The World Bank-IMF DSF analyses the behavior of Equation (4) under some scenario assumptions for $\{g_t, e_t, i_t^f, i_t^h, \pi_t\}$, and conducts stress tests. For the stress tests, large shocks are applied, in turn, to interest rates, exchange rates, or growth—keeping other endogenous variables fixed—and the subsequent behavior of the debt dynamics is projected. Ideally, Equation (4) would be embedded in a macroeconomic model suitable for policy analysis, but in practice these are not generally available. CDO discuss the limitations of World Bank-IMF DSF approach and implement a probabilistic scenario analysis that addresses some of its shortcomings. In particular, CDO acknowledge the interdependencies among the macroeconomic variables and draw values from an estimated joint distribution.

The CDO stochastic framework requires an empirical specification for the macroeconomic variables in the right-hand side of Equation (4). First, a VAR for the aforementioned input variables, $\{g_t, e_t, i_t^f, i_t^h\}$, is estimated. Second, a fiscal reaction function is specified and estimated, yielding the statistical relationship between the primary balances, past levels of debt, and the output gap. Finally, in Monte-Carlo simulations, shocks for the macroeconomic variables are drawn from a joint normal distribution, parameterized using the VAR-estimated covariance matrix. These shocks are in turn used in the forward iteration of a debt equation and the fiscal reaction function. The resulting fan charts thus provide a probability distribution for future debt-to-GDP ratios, which are contrasted against the DSF criteria.

In this paper, we propose methodological modifications addressing three aspects of the empirical formulation, which we think that may be open to improvement.

(1) *Structural breaks*. Estimation conducted for emerging market economies often includes episodes that feature financial crises—for example, Mexico in 1995, Asia in 1997, Turkey in 2001, and Argentina in 2002. During such periods, significant volatility in interest rate spreads, exchange rates and GDP growth have historically been observed. In this context, estimation of a standard unvarying VAR is problematic, as the assumption of stationarity of the data generating processes cannot be ensured. In this paper we allow for structural breaks, which are identified through a Markov-Switching VAR (SVAR). Further estimation is then conducted over periods exhibiting constancy in the data generating processes, which ensures consistent covariance estimates.

(2) *Normality*. The assumption of normality with regard to the underlying shock distribution constitutes a severe limitation of the Monte-Carlo analysis in CDO. The empirical behavior of economic variables, especially during period of financial distress, displays asymmetries and thick tails. Consequently, any inferred symmetry of the CDO debt projections may merely be driven by the normality assumption and thus conceals the true degree and shape of risk exposure. We replace the normality assumption by direct sampling from the joint distribution of the shocks. Identification of stable subsamples allows for the application of bootstrapping techniques. Consistent resampling of the data and repeated estimation implies that the resulting residuals converge to their true underlying counterpart. This then allows for the analysis of any non-normality such as asymmetry and for a better specification of the tail behavior of the shocks.

(3) *Fiscal reaction function*. CDO makes inference with regard to the dynamics between primary balances, lagged debt, and the output gap by estimating a fiscal reaction function, which, we think, is subject to severe model uncertainty. In this paper we avoid the use of a fiscal reaction function to predict the fiscal outturn and instead we specify, in each period, the fiscal balance to be the debt-stabilizing balance. As noted before, this specification is *not* intended to predict fiscal behavior, but rather to serve as an agnostic and sensible reference when assessing debt sustainability. We are interested in the effect of $\{g_t, e_t, i_t^f, i_t^h\}$ on the debt-to-GDP ratios, and this way we minimize the role of bad or good (predicted) fiscal behavior—for example, in this context, mounting debt cannot possibly be driven by the (debt-stabilizing) fiscal balance. By drawing repeated shocks from the empirical error distribution, in each period we calculate the implied time-varying debt-stabilizing balances. Alternatively, one could instead do a *fiscal-gap analysis* specifying in each period the required balance to bring the debt-to-GDP ratio to some desired level by some particular date.

Other methodological differences introduced here include the amendment of the debt equation in order to account for tradables, which are proxied by the export-to-GDP ratio. Furthermore, we do not assume that interest rate shocks affect the entire debt stock, but rather differentiate between its maturity structure such that merely the short-term and a uniformly distributed fraction of the medium to long-term debt is rolled over each period. Finally, with regard to data, our estimation is based on the fiscal panel and the input variables provided by CDO, except for the definition of the interest rate for foreign-denominated debt. Whereas CDO employ the deflated U.S. treasury bill rates as a proxy for borrowing costs in foreign markets, we incorporate interest rate spreads in the form of the Emerging Markets Bond Index, as we believe that this is more appropriate for emerging market countries.

II. Methodological Modifications

Markov-Switching VAR

The SVAR, proposed by Hamilton (1989) and Krolzig (1997), differs from its linear counterpart by allowing the structural coefficients and the covariance matrix to be dependent on an unobserved state variable S_t which is assumed to follow a Markov chain. This flexibility is important in our context, as we may expect, a priori, the joint distribution of the shocks to be nonconstant across our sample period. We estimate the following SVAR for the macroeconomic variables, $x_t = (g_t, e_t, i_t^f, i_t^d)$

$$\begin{aligned} x_t &= \upsilon_{S_t} + \beta^1_{S_t} x_{t-1} + \ldots + \beta^q_{S_t} x_{t-q} + \varepsilon_t, \\ &\quad \varepsilon_t \sim \mathrm{iid}(0, \Omega_{S_t}), \quad S_t \in \{S_1, S_2\}. \end{aligned} \tag{5}$$

Furthermore, estimation of this model allows for inference with regard to the probabilities of being in the respective states of the world, such that any structural breaks are identified by regime switches. The issue of predicting—going forward outside the sample period—which regime may apply falls outside the scope of this paper. Predicting currency or fiscal crises is tricky—for example, on the prediction of sovereign debt distress, see Manase, Roubini, and Schimmelpfennig (2003).

Bootstrap Draws

As discussed above, this paper replaces the assumption of normally distributed shocks imposed in CDO by instead drawing directly from their empirical joint distribution. To this end we employ nonparametric bootstrapping techniques on the stationary subperiods exhibiting constancy in terms of the data generating processes.

Within the general bootstrapping framework, the data is replicated through resampling and subsequent estimation is repeated using these resulting pseudo series, which exhibit the same statistical properties as the original sample. The parameters of interest are estimated from each bootstrap series, yielding their respective empirical probability distributions. In the context of time-series data, random resampling leads to inconsistent estimators in the case of dependency over time as well as across individual variables (Singh, 1981). In order to account for this, in this paper, replication is conducted using blocks of consecutive observations, as proposed by Künsch (1989). Equal probability of sample selection is ensured by applying the circular block bootstrap by Politis and Romano (1992).

With regard to the selection of block length, there is a trade-off between the approximation of the observed data characteristics and the randomness of the resampling mechanism. As the block length converges to the actual sample size, the statistical properties of the replicated series approach those of the original data, but at the cost of imposing ever increasing sequential restrictions, thus reducing the variability of the sampling process. Optimality is determined by an algorithm proposed by Politis and White (2004) whereby the length is based on the degree of autocorrelation exhibited by the data. In our application, we draw 1,000 pseudo series and sets of the residuals from their empirical distribution—which provides a consistent estimate of the true joint shock distribution.

III. Empirical Results

In the estimation of the SVAR,[1] we assume the existence of two data generating processes. Figure 1 presents evidence of structural breaks in the data, which suggests that estimation over the entire sample period potentially yields inconsistent parameter and covariance estimates, and incorrect debt projections.

In Argentina, a normal or tranquil period between 1993 and 2002 is identified, after which statistically significant regime switches coincide with the observed period of financial instability. As a result, in what follows, we conduct estimation separately for these two distinct states of the world.

With regard to Brazil, significant spillovers are found during the Mexican and Argentinean crises in 1995 and 2002, respectively, between which a subperiod of tranquility is defined, albeit a marginally significant break in 1997.

In South Africa, estimation is carried out across two stable periods, the former ranging from 1980 to 1987, and the latter from 1987 to 2001.

Figure 2 displays the quantiles of the residuals against the quantiles of a normal distribution—if the residuals were drawn from a normal distribution, the scatter would be a straight line. Figure 2 presents strong evidence of non-normality, especially during turbulent times. Moreover, the p-values for all joint normality tests are smaller than 0.001.

[1]The data used here covers the following periods—Argentina, 1993:Q1–2005:Q2; Brazil, 1995:Q1–2004:Q3; South Africa, 1980:Q1–2004:Q2. The identified breaks are (a) Argentina break, 2001:Q3, and (b) South Africa break, 1986:Q1; crisis end, 2001:Q1.

Figure 1. SVAR: Estimated Probability of Turbulent Times (S_2)

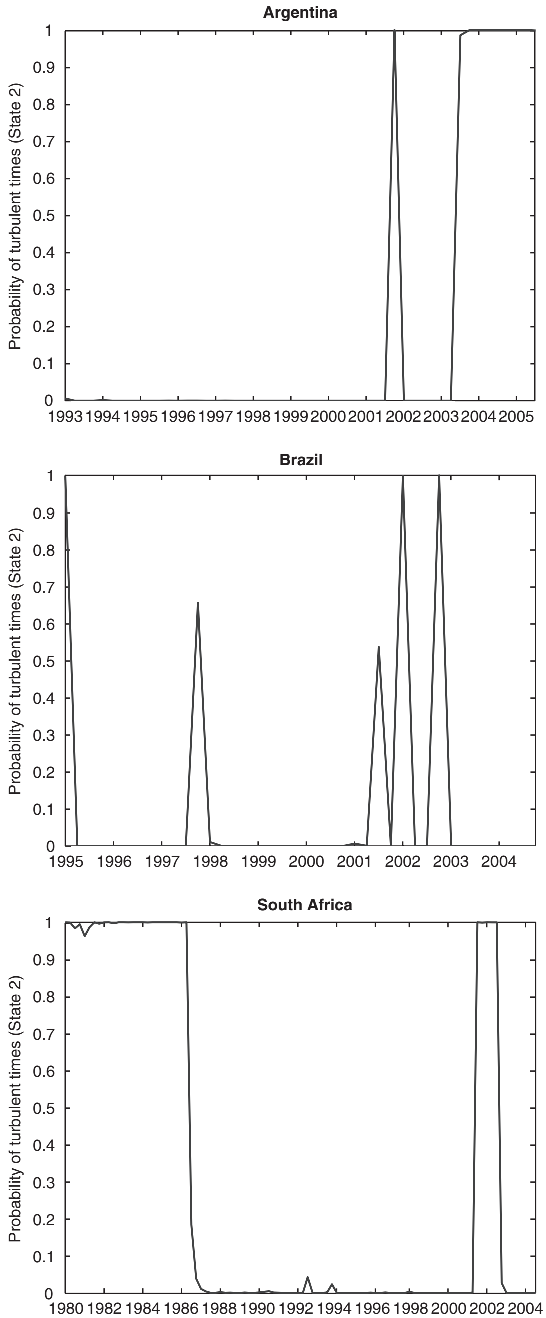
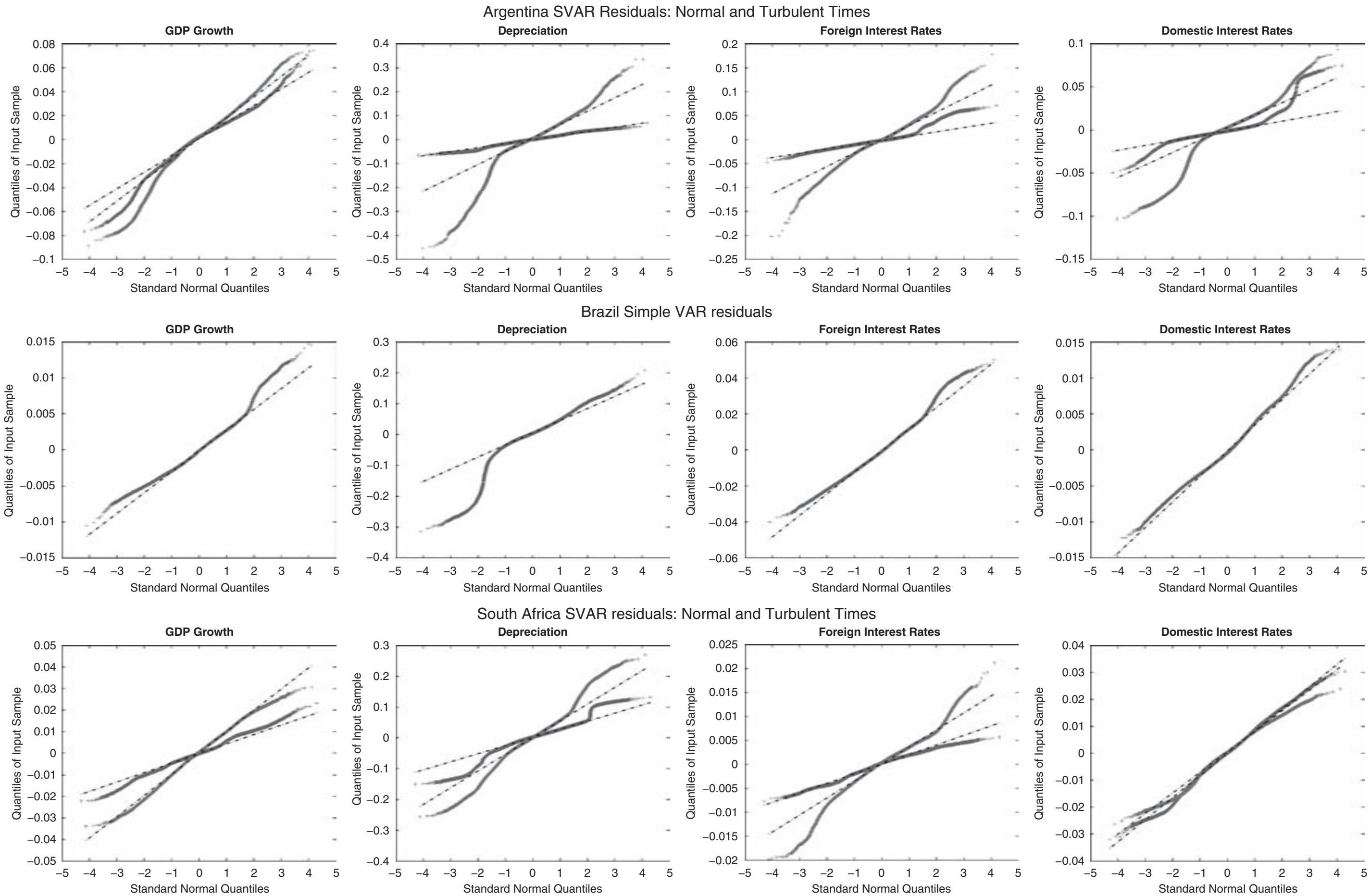


Figure 2. Quantile-Quantile Normal Plots for the SVAR Residuals: Argentina, Brazil, and South Africa



Argentina

As outlined above, estimation for Argentina is conducted separately for the two periods, which correspond to times of tranquility and financial distress, respectively. Figure 3 presents the distribution of the residuals from 1,000 bootstrap replications. As noted above, normality of the residuals is rejected (Figure 2).

Two further points of interest are to be noted. First, during the noncrisis and the crisis periods, shocks are drawn from differing distributions, whereby the latter exhibits a greater variance relative to the former (up to a fivefold factor). This finding highlights the problems associated with estimating a VAR across the entire sample period—as in this case shocks would be drawn from a mixture of these two distributions, thus implying misleading debt projections.

Second, there is evidence of asymmetric residuals. During the normal, tranquil times the interest rate distributions exhibit greater upside risk, whereas during the crisis period, the histogram for the exchange-rate movements is characterized by a long left-hand tail. Clearly, the assumption of normality adopted in the existing literature does not allow for the quantification of this asymmetric tail behavior. The imposition of symmetry directly affects the probability distribution of the debt projections and thus leads to inappropriate conclusions with regard to fiscal sustainability.

Figure 3. Argentina—Histograms of the SVAR Residuals: Normal and Turbulent Times

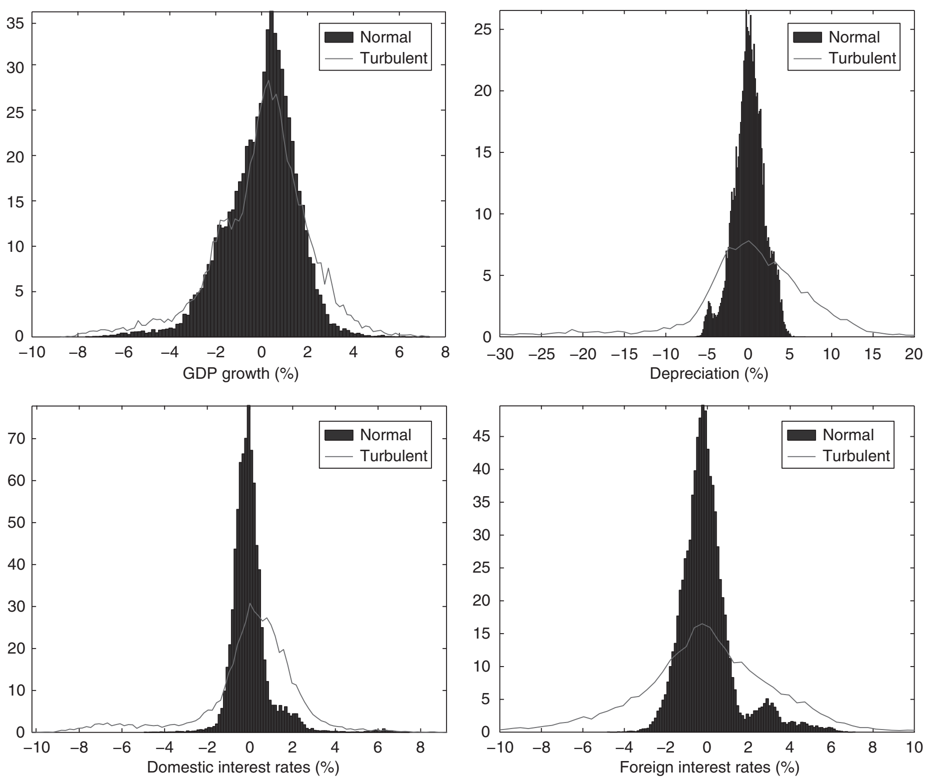


Figure 4. Argentina: Debt-to-GDP Projections, in Normal and Turbulent Times, Contrasted with CDO Projections

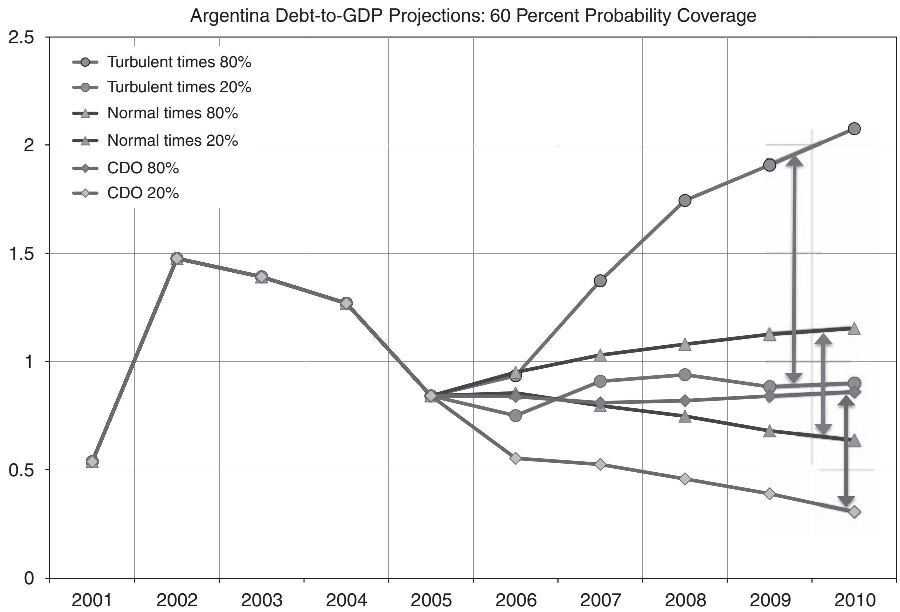
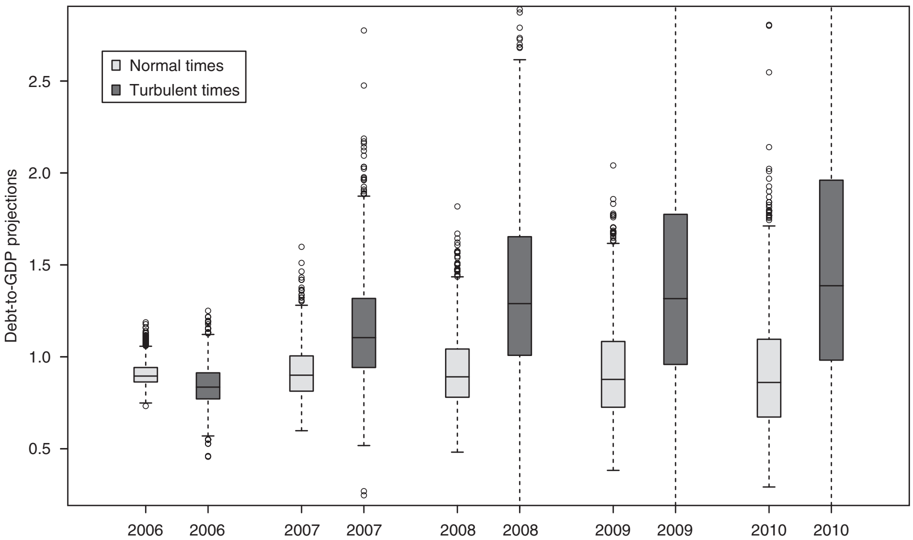


Figure 4 compares our projections with those of CDO. We represent the second and eighth deciles, whose range covers 60 percent of the probability mass. As it can be seen, the CDO projections present a more optimistic picture than even our normal-times projections. This is driven by the CDO's assumed prudent fiscal behavior embedded in their estimated fiscal-reaction function. In our projections, in contrast, we specify the debt-stabilizing balance in each period—conditioned on the current values of the macroeconomic variables, which then get shocked going forward. This is a more agnostic approach, allowing us to get a sense of how things might look under minimal fiscal adjustment. Another significant difference with respect to CDO arises by allowing asymmetries in the shock behavior, which is especially important during turbulent times, as shown in Figure 5.

The observed asymmetries and differing variances of the shock distributions are directly translated into the debt projections shown in Figure 4 and 5. During the noncrisis period, the mean debt level remains approximately constant, but there is evidence of upside risk in the form of the skewed probability distribution of future debt levels. This is driven primarily by the positive interest rate shocks illustrated in Figure 3. During the period of financial instability both the average debt level and the variance of the forecasts increase significantly, whereby the former is due to an upward jump in interest rate spreads raising the costs of borrowing. As a conclusion of this country example, we argue that estimation over the entire sample period will, in addition to the aforementioned parameter inconsistency, incur an additional information loss. If the structural breaks are ignored, the

Figure 5. Argentina: Boxplots of the Debt-to-GDP Projections in Normal and Turbulent Times



resulting debt projections would incorrectly be comprised of a mixture of those represented in Figure 5.

Brazil and South Africa

Figure 6 displays replicated VAR residuals for Brazil. We cannot estimate a SVAR in this case because the normal and turbulent states alternate too often, resulting in subsamples that are too short. In the case of Brazil, the most striking features with regard to asymmetries are the upside risk associated with interest rate changes, in addition to the long left-hand tail in the distribution for exchange rates—indicating greater probability of a currency depreciation. As in the case of Argentina, the assumption of joint normality is rejected. These asymmetries in the data are reflected in the debt projections in Figure 7. Whereas the mean level of future debt remains approximately constant, the box plots illustrate the significant probability mass corresponding to increased debt-to-GDP ratios.

As argued previously, for South Africa there appear to be two subsamples exhibiting constancy in their data generating mechanisms. As a result, estimation is conducted for 1980–87, which is defined as ‘‘normal’’ period, and for 1987–2001, ‘‘turbulent’’ period. This country case differs in terms of interpretation when compared with Argentina, where a distinct episode of financial distress is identified. Figure 8 shows that asymmetries in the VAR residuals are not as pronounced for South Africa, but that the variances differ across the two distributions. This propagates to the debt projections, which exhibit differences in terms of the width of their respective confidence intervals. Also, it is noteworthy, that for both subperiods the debt

Figure 6. Brazil—Histograms of the VAR Residuals

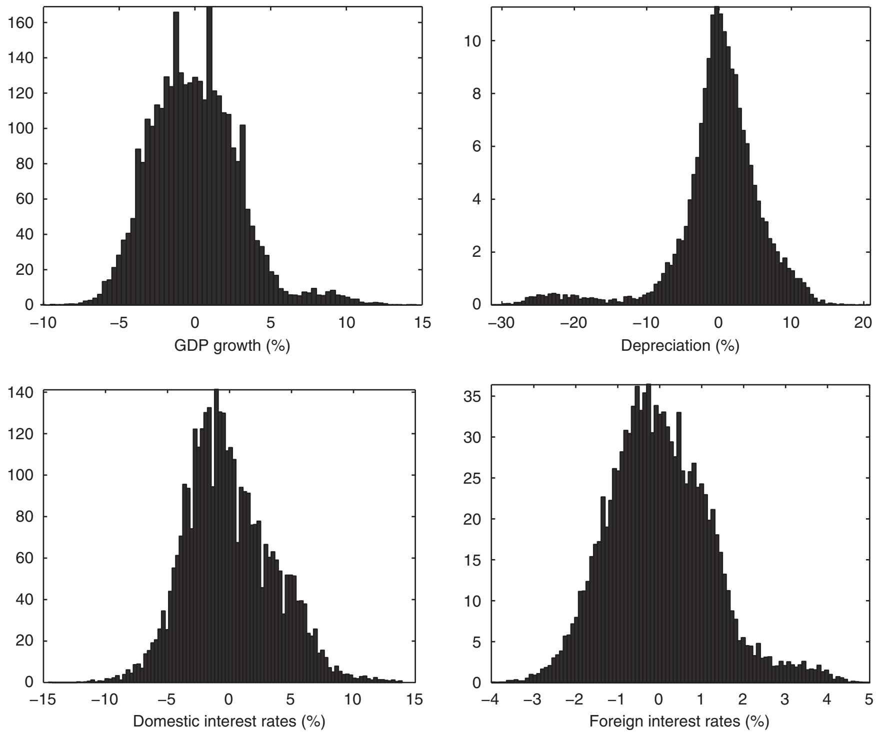


Figure 7. Brazil: Boxplots of the Debt-to-GDP Projections

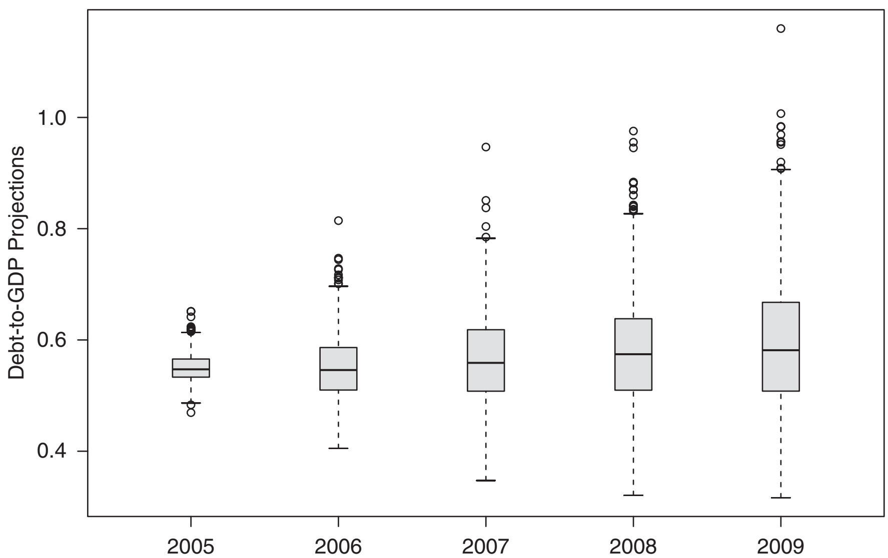
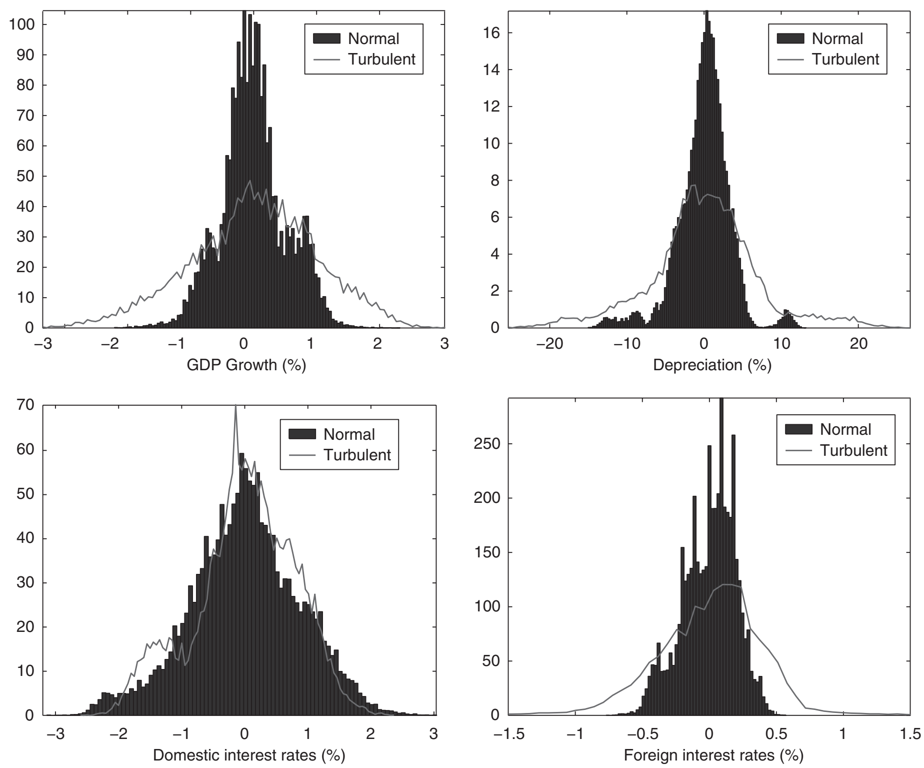


Figure 8. South Africa—Histograms of the SVAR Residuals: Normal and Turbulent Times



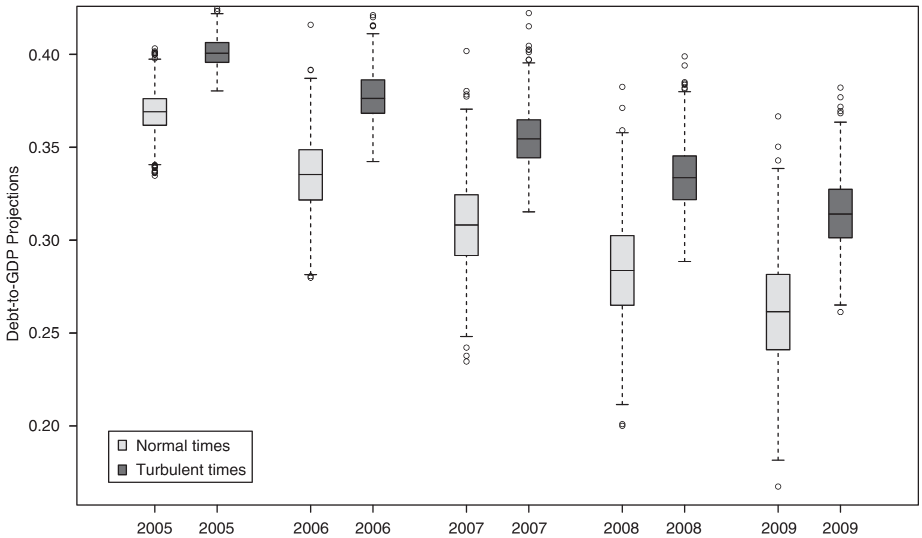
forecasts are downward sloping, a fact that is driven by low interest and high GDP growth rates. These are unlikely to continue in the future, and we do not take these projections as realistic—we just illustrate what the mechanical application of this framework implies (Figure 9).

IV. Debt-Stabilizing Balances

The final departure from the technique proposed by CDO is that we avoid the use of fiscal reaction functions altogether. Their usage introduces estimation and model uncertainty when attempting to make inference with regard to the statistical relationship between primary balances, lagged debt, and output growth. In CDO, shocks are drawn from the joint distribution of the residuals, which in turn are utilized in the forward iteration of the fiscal reaction function and the debt equation.

Here we invert the problem by specifying the government balance which, under the current values of $x_t = (g_t, e_t, i_t^f, i_t^d)$ is consistent debt stabilization and we obtain the implied balances within this stochastic framework. Given the complex form of the debt equation, which accounts for differing maturity structures of debt, a simple inversion is not possible. Thus numerical

Figure 9. South Africa: Boxplots of the Debt-to-GDP Projections in Normal and Turbulent Times



techniques in the form of a grid search are employed. This fiscal balance is a benchmark for minimal adjustment. This technique has the advantage—in addition to overcoming data constraints affecting the fiscal variables—that it allows for evaluation of currently implemented policies, by contrasting them to the debt-stabilizing ones.

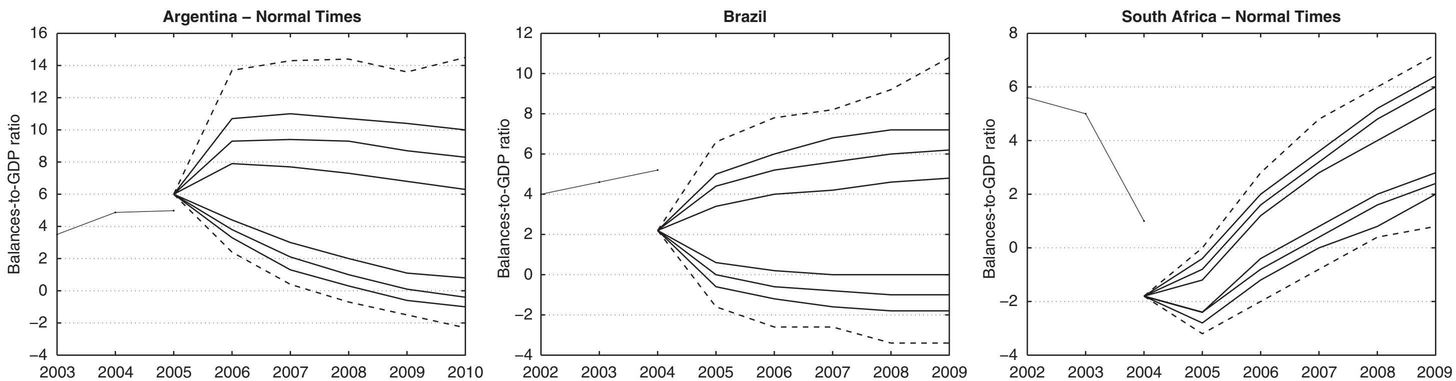
Figure 10 provides both the actual primary balances and those required in order to achieve debt sustainability under the true shock distribution. In Argentina and Brazil governments are currently achieving balances of approximately 5 percentage points of GDP, which is consistent with our notion of fiscal sustainability. As pointed out above, the decreasing debt levels for South Africa are driven by high levels of GDP growth and low interest rates. These factors allow their respective government to incur negative primary balances, whilst not endangering their future fiscal position.

V. Conclusion

This paper proposes several methodological modifications to the probabilistic fiscal sustainability analysis framework developed by Celasun, Debrun, and Ostry (2007).

First, in the VAR specification, we allow for structural breaks in the data-generation mechanism, through the application of Markov-Switching models. This is important when characterizing the statistical behavior of growth, interest and exchange rates for countries that may have undergone crisis or even just turbulent times. Second, in the Monte-Carlo simulations, we replace the assumption of normally distributed shocks by bootstrapping techniques to draw directly from the empirical distributions. This allows for a

Figure 10. Debt-Stabilizing Balances: Argentina, Brazil, and South Africa



better risk assessment as it allows for thicker tails and asymmetries in the debt projections. Third, the estimation of the fiscal reaction function is avoided by using instead the debt-stabilizing balance each period when producing baseline projections.

This specification is *not* intended to predict fiscal behavior, but rather to serve as an agnostic reference when assessing debt sustainability—most of the action in the behavior of the debt-to-GDP ratios will have to come from elsewhere. We look at what would be the predictive density for the debt-to-GDP ratios if the government's balance at each period were precisely the debt-stabilizing balance conditioned on current values of interest rates, GDP growth, and the exchange rate, while then going forward, shocks to these variables exhibited joint behavior similar to past experience—differentiating between normal (tranquil) and turbulent (nontranquil) times, when needed.

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External Linkages and Contagion Risk in Irish Banks

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Increasing financial integration makes the assessment of cross-country linkages crucial for effective financial surveillance. This paper estimates contagion risk between large Irish banks and European and U.S. banks during 1994–2005, using distance-to-default measures and the methodology of extreme value theory. Employing an ordered logit model, and controlling for Ireland-specific and global shocks, we find evidence of significant contagion risk coming from the United Kingdom, the United States, and the Netherlands toward Ireland. We also find that patterns of contagion to Irish banks have shifted over time, coming from the United Kingdom in the pre-euro period and from the United States in the post-2001 period. [JEL C51, G21]
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In the past decade Ireland has increasingly become financially open. Global trends in financial liberalization, innovation, and banking consolidation activity, as well as greater bank reliance on wholesale funding, increasing competition from foreign-owned banks, and the development of the Irish

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Financial Services Center (FSC) have led to increasing international exposure for the banking system. At the same time, the Irish domestic banking system has remained relatively concentrated, with Bank of Ireland (BoI), Allied Irish Banks plc (AIB), and Anglo Irish Bank Corp. plc (Anglo IB) representing about 45 percent of total banking assets but nearly 80 percent of the domestic retail market.

In this paper, following Gropp, Lo Duca, and Vesala (2005)[1] and De Nicolo and Tieman (2005), we explore whether international openness is accompanied by increasing financial risk for the three major Irish banks.[2] We explore trends in international interdependencies between Irish banks and banks in the major European countries and in the United States over the period 1994–2005, and analyze whether international linkages have led to possible contagion risk. We assess whether and to what extent large negative shocks in banks of these countries affected banks in Ireland. In particular, we measure the probability that a higher number of banks in Ireland will experience a large negative shock at the same time as banks in other countries (“coexceedances”). The existence and the magnitude of these effects have implications for the monitoring of financial stability, as contagion risk is widely perceived to be an important element of banking crises and systemic risk. Understanding the direction of such risk is important for policymakers to better focus limited financial supervisory resources.

We use the term “contagion” to mean the transmission of an idiosyncratic shock affecting one bank or a group of banks to other banks or other banking sectors, using distance to default (DD) as a measure of bank risk. There could be a number of possible contagion channels (Box 1) and several ways to measure contagion risk: estimating autocorrelation and survival time tests using historical data on bank failures (Calomiris and Mason, 2000), using the interbank market lending exposure matrix and measuring domino effects of insolvency in one bank spreading to others (Furhine, 2003), or analyzing the correlations of stock returns of banks to measure interdependencies (De Nicolo and Kwast, 2002).

The paper follows the approach employed in the recent literature, which estimates contagion risk by using the DD as a comprehensive measure of default risk (Gropp and Moerman, 2004; De Nicolo and Tieman, 2005; Gropp, Lo Duca, and Vesala, 2005). The DD represents the number of standard deviations away from the default point—the point at which the book value of liabilities of the bank are just equal to the market value of assets. Unlike unadjusted stock returns, the DD combines information about stock returns with leverage and asset volatility information, thus encompassing the most important determinants of default risk. The higher the DD, the greater the distance of the bank from default point, and the

[1]Gropp, Lo Duca, and Vesala (2005) find evidence of significant cross-border contagion in Europe (Ireland not included in the sample) during January 1994 to January 2003.

[2]The choice of the banks was guided by the availability of data on their equity prices.

Box 1. Sources of Contagion: A Literature Survey

Sources of cross-country contagion risk can be grouped into the following:

- *Bank deposit runs*: Traditionally, the literature has focused on the implications of bank deposit runs for the payment system, the money supply, and financial intermediation. Liquidity shocks hitting one bank could cause depositors to also run on other solvent banks, in fear of lacking reserves of liquid assets in the banking system (Freixas, Parigi, and Rochet, 2000). The runs could be triggered by rumors about banking system fragility, reputation, or operational risk in countries where banks may have subsidiaries, branches, and even representative offices.
- *Wholesale funding channels*: With the developments in technological change, deregulation, globalization, and the increased use of financial markets, the focus has shifted to systemic risks arising at the wholesale level (intermediation, investment banking, securities trading, asset management, business banking), and concentrating on the largest and most complex financial institutions.
- *Liquidity and credit risk in interbank markets*: One possible channel of contagion is through the interbank market (Allen and Gale, 2000; Freixas, Parigi, and Rochet, 2000) in the form of liquidity shocks when banks withdraw their deposits at other banks, or in the form of credit risk when deposits at other banks are not being repaid.
- *Unidentified channels*: There could also be contagion in the financial markets in the absence of explicit links, when in the presence of asymmetric information difficulties in one market are perceived as a signal of possible difficulties in others (Morgan, 2002).

less the risk or probability of default of the bank. The benefit of this approach is that co-movements in DD can be analyzed without specifying a particular channel of contagion. Rather, these co-movements reflect interdependencies between domestic and cross-border banks encompassing all potential channels of contagion, including those occurring in the absence of explicit links between banks.

We first use rolling correlations of changes in DD of Irish banks with major European and U.S. banks to analyze trends in cross-country interdependencies. Then, we use an ordered logit model to estimate contagion risk: the probability of Irish banks experiencing a large shock on the same day (coexceedances) as banks in other countries after controlling for Ireland-specific and global factors—common shocks that could affect all banks simultaneously. Large shocks are defined by the bottom 15th percentile of the weekly difference in the daily DD of all banks. The 15th percentile threshold was used, instead of (say) the bottom 5th percentile, because we wanted to include incidents that have a higher probability of occurrence—those that could be associated with large shocks, and not necessarily only with crisis scenarios.

Analyzing trends in rolling correlations in the percentage changes in DD, we find evidence of increasing simultaneous occurrence of shocks across countries, suggesting increasing global interdependencies over the last decade. Further, following the approach of Gropp, Lo Duca, and Vesala (2005), we find evidence in favor of significant cross-border contagion from the United Kingdom, the United States, and the Netherlands to Ireland.

Moreover, it seems that contagion risk has shifted from coming from Europe in the pre-euro period to coming from the United States in the post-euro period. These findings have policy implications for the monitoring of financial stability and could direct scarce supervisory resources in managing foreign exposures.

However, all results come with significant caveats: we are using equity prices of banking *groups*, and yet mainly discussing international links typically associated with banking without explicitly stating links through securities and insurance (the top two Irish banks have links to insurance companies). We are using data available for only a small group of listed banks in each country—so it might not be representative of the system, and the banks in Ireland comprise less than 50 percent of the system (although a much larger share of the retail market). The number of observations experiencing large shocks is low; the results could, therefore, be driven by the large shocks associated with the tech-bubble burst in 2000.

The following section describes various external linkages of Irish banks. Trends in interdependencies using rolling correlations of DDs between the top three Irish banks and banks from other European countries and the United States are analyzed in Section II. An ordered logit model, to estimate the probability of coexceedances in Irish banks, is presented in Section III, while Section IV summarizes and concludes.

I. Some Evidence of External Linkages of the Irish Banking Sector

This section aims to find the channels and the direction of potential contagion risk from other countries to Ireland, looking at external linkages of the Irish banks. The Irish financial system has been more closely integrated with the United Kingdom in the past, but in recent years links with continental Europe and the United States have strengthened. There are many external linkages of the Irish banking sector, stemming from direct equity exposures in cross-border banks; direct exposure through loan books in other countries; deposit and funding sources from other countries or from numerous foreign banks operating in Ireland; stock market participation in other countries—through securities and asset management firms and holding of credit risk transfer (CRT) instruments written on assets located in another country, indirectly exposing Irish banks to international shocks. Each of these is explored in turn.[3]

First, Irish banks have foreign equity exposure through their expansion overseas. Both AIB and BoI have sizable operations in the United Kingdom and have been growing their niche wholesale international businesses. Both

[3]Some empirical evidence of foreign contagion risk in Ireland is provided by Gropp and Moerman (2004), who study the joint occurrence of both positive and negative extreme shocks in banks' distance to default (DD) among large European Union (EU) banks in the period January 1991 to January 2003, and identify AIB and BoI as among the systemically important banks in the European Union. Moreover, they find a contagion effect from the United Kingdom to Ireland.

banks own universal banks in Northern Ireland. AIB has a large stake in a U.S. regional bank and majority-owns a Polish bank—the U.S. and Polish investments contributed a combined 16 percent to AIB's pretax profit in the first half of 2005 (Standard & Poor's, 2006), although the Polish operations also result in a relatively high nonperforming loans.

Second, Irish banks have large loan-book exposures abroad. The two largest banks, AIB and BoI, are geographically diversified—each with almost equal share of domestic and foreign assets. Nearly 28 percent of AIB, 44 percent of BoI, and 41 percent of Anglo IB loan book exposures were in the United Kingdom. Although AIB held over a 20 percent stake in a U.S. bank, U.S. operations were only 2 percent of its loan book. Anglo IB, on the other hand, has about 5 percent of its loan book exposed to the United States; without equity exposures in the United States, it operates through a representative office.

Third, the increased reliance on wholesale funding in recent years, including interbank borrowing and capital market issues, is another potential source of international interdependencies. The average loan-to-deposit ratio for Ireland exceeds 150 percent, one of the highest for industrial countries. All major Irish banks are dependent for funds on the interbank and securities markets—AIB and BoI fund about 40 percent of lending in the wholesale market, but the market funding requirement for Anglo IB is about 35 percent. The overwhelming bulk of both nonresident interbank borrowing (83 percent) and debt securities issued and held by nonresidents (83 percent) in 2004 were vis-à-vis non–euro area residents.

Fourth, Irish banks buy risk protection mainly from banks of other countries. The underlying asset in structured CRT products include mainly loans and bonds issued by financial and nonfinancial firms; mortgages (for asset-backed securities); and financial and nonfinancial firm debt as underlying assets for the more traditional CRT (for example, mortgage indemnity guarantee). The United States and the United Kingdom are the main counterparty locations selling risk protection to Irish banks. Other countries include France, Germany, Canada, Switzerland, the Netherlands, Italy, and Poland. The major currencies of denomination are the euro and the dollar, along with the pound sterling. Irish banks also issue covered bonds—BoI has transferred the bulk of its domestic residential mortgage assets to a designated mortgage credit institution, which has a banking license to issue mortgage-covered securities. These covered bonds are used both for hedging interest rate risk and for generating additional funding. Almost 60 percent of these securities are held by other euro area members, but 25 percent are held in dollars by other countries.

Fifth, Irish banks are directly and indirectly exposed to property markets abroad. All the top three banks have loan-book exposures to the U.K. property market. At least AIB and Anglo IB sell mortgages in the United States—AIB through its U.S. subsidiary, and Anglo IB through its representative office. The latter is more focused on commercial property lending in the United States. BoI has launched a new venture with a leading Spanish bank, La Caixa, to provide extra mortgage options for Irish people buying property in Spain, which includes equity release from existing BoI mortgages. Part of the real estate price risk is mitigated by the Irish banks buying risk protection against these exposures. Irish legislation on covered bonds broadens the scope of risk protection by making loans made from countries such as the United States, Canada, Switzerland, and Japan eligible for the collateral pool.[4] However, Irish banks could be indirectly exposed to property markets by *selling* risk protection (buying of covered bonds, credit default swaps, and mortgage-backed securities) to other banks that are exposed to foreign property markets. From anecdotal evidence, some small FSC banks, exposed to international property markets, are selling credit default swaps to other domestic-oriented banks, indirectly exposing the latter to these property markets even with no loan book exposure.

Sixth, Bank for International Settlements (BIS) data on banks *resident* in Ireland show their net asset positions vis-à-vis banks and nonbanks in various countries (Figure 1). During 2001–05 the Irish resident banks had a negative net asset position vis-à-vis banks and nonbanks in the United Kingdom, an overall positive net asset position vis-à-vis banks and nonbanks in the United States, and large positive net asset position with Italy, Spain, and France. However, much of these positions could belong to foreign-owned banks operating through branches in the FSC[5] in Dublin—some of these banks operate almost exclusively with nonresidents and have some, but limited, links with the domestic economy.

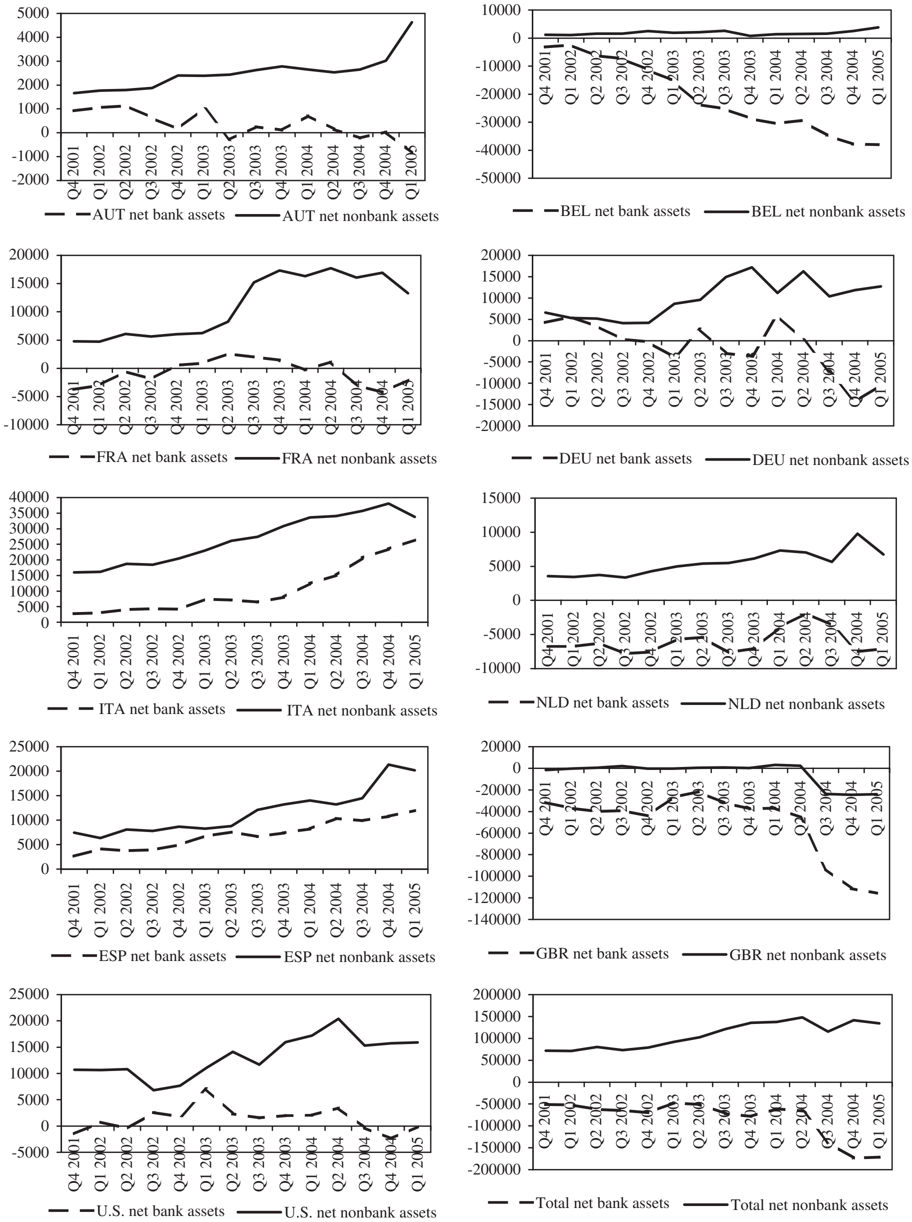
Seventh, BIS consolidated statistics report Irish banks' on-balance sheet financial claims vis-à-vis the rest of the world. These type of data would give an idea of the extent of exposure of the Irish banks during a credit event in these countries. Irish banks are mostly exposed to the United Kingdom and Germany among EU countries—and to the United States (although small in comparison) among non-EU countries.

Given the above evidence, we could expect the following possible channels of contagion risk: between Ireland and the United Kingdom and the United States, on equity exposure, loan exposure, and exposure via the interbank and the securities markets; between Ireland and the Netherlands, based on the presence of large the Netherlands-owned banks in Ireland; and between Ireland and Italy, Spain, France, and Germany based on interbank market exposure. The direction of contagion is more challenging to establish.

[4]With the exception of Luxembourg, most European countries limit the asset pool to European Economic Area assets.

[5]The FSC was established in Dublin in 1987 (known at that time as the International Financial Services Center or the IFSC) to facilitate financial operations with nonresidents, and was endowed with corporate tax benefits. These tax benefits have now been lifted. There are about 450 international institutions that have established offices in the FSC (accounting for about 40 percent of banking assets), including 50 percent of the 50 largest financial institutions in the world—including major banks from the United States, the United Kingdom, the Netherlands, Italy, and Germany. In turn, almost all domestic credit institutions also conduct business from the FSC.

Figure 1. Ireland-Resident Banks, Net Asset Position vis-à-vis Banks and Nonbanks in Various Countries
(In millions of U.S. dollars)



Source: Authors' calculations from Bank for International Settlements banking statistics. http://www.bis.org/statistics/bankstats.htm.

Note: DEU refers to Germany, ITA—Italy, GBR—Great Britain, U.S.—United States, ESP—Spain, AUT—Austria, BEL—Belgium, NLD—Netherlands, and FRA—France.

However, evidence from BIS data (Figure 1) and the discussion in this section suggest the following:

- Countries with which Ireland has a net nonbank asset position would likely expose Ireland to credit risk in loans and other asset markets: Austria, France, Germany, Italy, the Netherlands, Spain, and the United States.
- Ireland's positive net bank exposure in Italy and Spain, and previously the United States, Austria, and France, possibly exposes Ireland to credit risk that could lead to liquidity risk.[6]
- Ireland has a negative net bank-asset position, which exposes it to liquidity risk if some problem were to arise in the country from which Ireland is borrowing: Germany, the United Kingdom, Belgium, the Netherlands, and currently the United States, Austria, and France.

There could also be contagion in the absence of explicit links, when due to asymmetric information, difficulties in one country/market are perceived as a signal of possible difficulties in others. The interdependencies mentioned above may be overstated to the extent they include banks resident in Ireland but operating in the FSC mainly with nonresidents; however, data were not sufficient to exclude these from the BIS data sample. We next look at only the three largest domestic banks—that have nearly 80 percent of the retail market—and their interlinkages and possible contagion risk channels with banks in other European countries and in the United States.

II. Trends in Interdependencies Using DD Indicators

This section describes correlations of banking risks between Ireland and other countries after a discussion of the data. We use the DD indicator to measure bank financial risk; the DD combines information about stock returns with leverage and asset volatility information, thus encompassing the most important determinants of default risk. The cross-country correlation of changes in the DD would indicate interdependencies arising from a broad set of channels, including contagion occurring in the absence of explicit links between banks.

The DD is based on the Black-Scholes option-pricing model, and is estimated using stock price data (Box 2). A bank's equity is viewed as a call option on the bank's assets, with the strike price equal to the current book value of total liabilities. When the value of the banks' assets is less than the strike price, its equity value is zero. The DD represents the number of asset value standard deviations that the bank is away from the default point, where the default point is defined as the point at which the liabilities of the

[6]The Irish banking supervisors have a specific requirement for banks on liquidity maintenance: 25 percent of deposits and short-term liabilities have to be covered by liquid assets.

Box 2. The Distance to Default (DD) Measure and Data Issues

The derivation of distance to default (DD) is described in detail in Gropp, Lo Duca, and Vesala (2005) and in Gropp and Moerman (2004), and in the case of a portfolio of bank assets in De Nicolo and Tieman (2005). The DD measure is based on the structural valuation model of Black and Scholes (1973) and Merton (1974), and is defined as follows:

$$DD_t = \frac{\ln(V_{A,t}/X_t) + (r - (\sigma_A^2/2))T}{\sigma_A\sqrt{T}},$$

where $V_{A,t}$ is the firm's assets value with mean r and volatility σ_A, and X_t is the book value of the debt at time t, that has maturity equal to T. The market value of equity of the firm is viewed as a call option on the firm's assets, V_A, with time to expiration equal to T. The strike price of the call option is the book value of the firm's liabilities, X_t. Default occurs when the value of the firm's assets is less than the strike price—that is, when the ratio of the value of assets to debt is less than one. The DD tells us by how many standard deviations the log of this ratio needs to deviate from its mean in order for default to occur.

An estimation of DD requires knowing both the asset value and asset volatility of the firm. The required values, however, correspond to the forward-looking *economic* values rather than the accounting figures, and it is not appropriate to use balance-sheet data for estimating these two parameters. Instead, the asset value and volatility are estimated using equity data. The DD measures we use are taken from IMF's Monetary and Capital Markets Department Distance-to-Default database, with the methodology described in Vassalou and Xing (2004), except that the value of assets is taken to be equal to the value of equity plus the book value of liabilities. At each date, the value of assets, the return on assets and its volatility is derived using the Black-Scholes option-pricing formula, using one year of daily equity return data preceding the estimation date, and the accounting value of liabilities for the relevant year.

Declines in the $V_{A,t}/X_t$ ratio are equivalent to declines in capitalization. Thus, the DD measure combines information about equity returns with leverage and asset volatility information, hence encompassing the most important determinants of default risk. Empirical studies have shown that the DD is a good predictor of corporate defaults (Moody's KMV), and predicts banks' downgrades in developed and emerging market countries (Chan-Lau, Jobert, and Kong, 2004; Gropp, Vesala, and Vulpes, 2006).

We use a data set of daily DD data for 40 banks in eight countries: France (two banks), Germany (four banks), Ireland (three banks), Italy (six banks), the Netherlands (two banks), Spain (four banks), the United Kingdom (five banks), and the United States (14 banks), for the period January 1994 to November 2005 (Table 1). The data set includes all banks in these countries that are listed at a stock exchange and whose DD are available from the Distance-to-Default database. We dropped four banks for which the DD were not available for the entire period (one bank for France, one bank for the United Kingdom, and two banks for Italy). In general, the banks in the sample are quite large relative to the population of banks in the European Union, and represent a high fraction of total assets of commercial banks in each country. For each bank, the sample contains 3,105 daily observations (except one U.K. bank with 2,522 observations and one bank from the Netherlands with 2,803).

bank are just equal to the market value of assets (alternatively, the point where the stock price is zero). The market value of assets is not observable, but is estimated using equity values and accounting measures of liabilities (Table 1).

The banks in the sample are just about five standard deviations from the default point—the (pooled) average DD is 5.2; the median 5.0. There is some

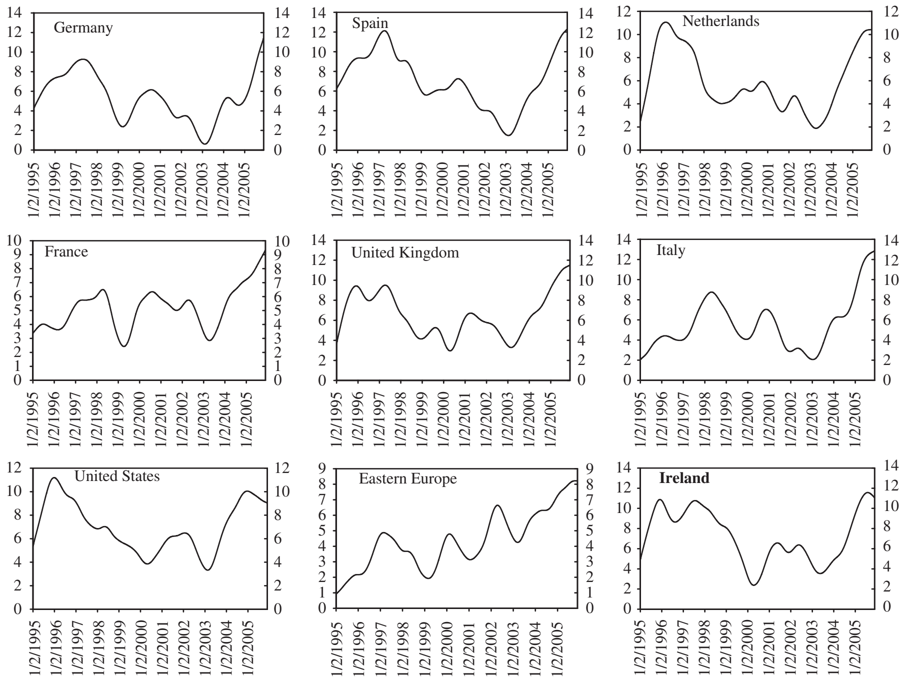
Table 1. Sample Banks

1	ALLIED IRISH BANKS	Ireland
2	ANGLO IRISH BANK CORP.	Ireland
3	BANK OF IRELAND	Ireland
4	CITIGROUP	U.S.A.
5	BANK OF AMERICA	U.S.A.
6	JP MORGAN CHASE & CO.	U.S.A.
7	WELLS FARGO & CO.	U.S.A.
8	WACHOVIA	U.S.A.
9	U.S. BANCORP	U.S.A.
10	SUNTRUST BANKS	U.S.A.
11	NAT.CITY	U.S.A.
12	BANK OF NEW YORK	U.S.A.
13	BB & T	U.S.A.
14	FIFTH THIRD BANCORP	U.S.A.
15	STATE STREET CORP.	U.S.A.
16	KEYCORP	U.S.A.
17	PNC FINL.SVS.GP.	U.S.A.
18	BARCLAYS	U.K.
19	HSBC HDG.	U.K.
20	LLOYDS TSB GP.	U.K.
21	RYL. BK. OF SCTL.	U.K.
22	STD. CHARTERED	U.K.
23	BANKGESELLSCHAFT BERLIN	Germany
24	BAYER.HYPO-UND-VBK.	Germany
25	COMMERZBANK	Germany
26	DEUTSCHE BANK	Germany
27	BNP PARIBAS	France
28	SOCIETE GENERALE	France
29	BANCO ESPANOL DE CREDITO	Spain
30	BANCO POPULAR ESPANOL	Spain
31	BANCO SANTANDER CENTRAL HISPANO	Spain
32	BBV ARGENTARIA	Spain
33	ABN AMRO HOLDING	Netherlands
34	FORTIS (AMS)	Netherlands
35	UNICREDITO ITALIANO	Italy
36	SAN PAOLO IMI	Italy
37	CAPITALIA	Italy
38	BANCA INTESA	Italy
39	BANCA INTESA RNC	Italy
40	UNICREDITO ITALIANO RNC	Italy

variation among banks—the mean DD by bank ranges from 3.3 (Italy) to 7.7 (Spain). The three Irish banks have mean DD of about 6 (median about 5.7). Figure 2 shows the trends in the system-wide DDs by country, including an Eastern Europe average comprising Poland and Hungary.[7] There seems to be a general increase in the DD starting in 2003 for all countries, indicating a

[7]Weighted average DD of each listed bank, weighted by assets.

Figure 2. Hodrick-Prescott (HP) Trend of Banks' Aggregate Distance to Default (DD) by Country



Note: Each panel shows the aggregated DD of sample banks—weighted by assets—in a country. See Table 1 for the sample of banks from each country. The aggregates are HP-filtered. An increase shows a decline in systemic risks.

global improvement in bank health in the past two years. There also seems to be an upward trend over the past decade for France, Italy, and Eastern Europe. Irish banks suffered a trend decline in the DD around 1997 that continued until end-1999—this partly coincides with the experiences of the United States, the United Kingdom, and (partially) the Netherlands, Germany, and Spain, as does the continued recovery after 2003. They had suffered another negative shock during 2002—likely the after-effects of September 11, 2001 and possibly connected with the AIB's U.S. subsidiary scandal in 2002. Irish banks have recovered starting in 2003 and currently have DD levels comparable to the levels of the late 1990s.

Table 2 shows the correlation of the DD between Ireland and the other countries. We find that on average the correlations are positive, and quite high for the Netherlands, the United Kingdom, the United States, and Spain. However, if we analyze the past decade in three separate periods—pre-euro period 1/3/1994–12/31/1998, post-euro period 1/1/1999–9/11/2001, and post–September 11th period of 9/12/2001–11/25/2005, we find that the correlations become generally much smaller or negative in the post-euro period compared

Table 2. Correlations in Distance to Default (DD)

Ireland and ...	1994–2005	Pre-Euro 1994–98	Post-Euro 1999–2001	Post–Sept. 11th 2001–05
France	0.31	0.34	−0.53	0.89
Germany	0.64	0.61	−0.69	0.81
Italy	0.5	0.37	0.18	0.88
Netherlands	0.71	0.61	−0.39	0.94
United Kingdom	0.81	0.68	0.63	0.92
United States	0.73	0.45	0.83	0.76
Spain	0.76	0.74	−0.37	0.91
Eastern Europe	0.03	−0.12	−0.71	0.81

Note: Cross-country correlations in banking systems' DD. The sample covers January 3, 1994–November 25, 2005.

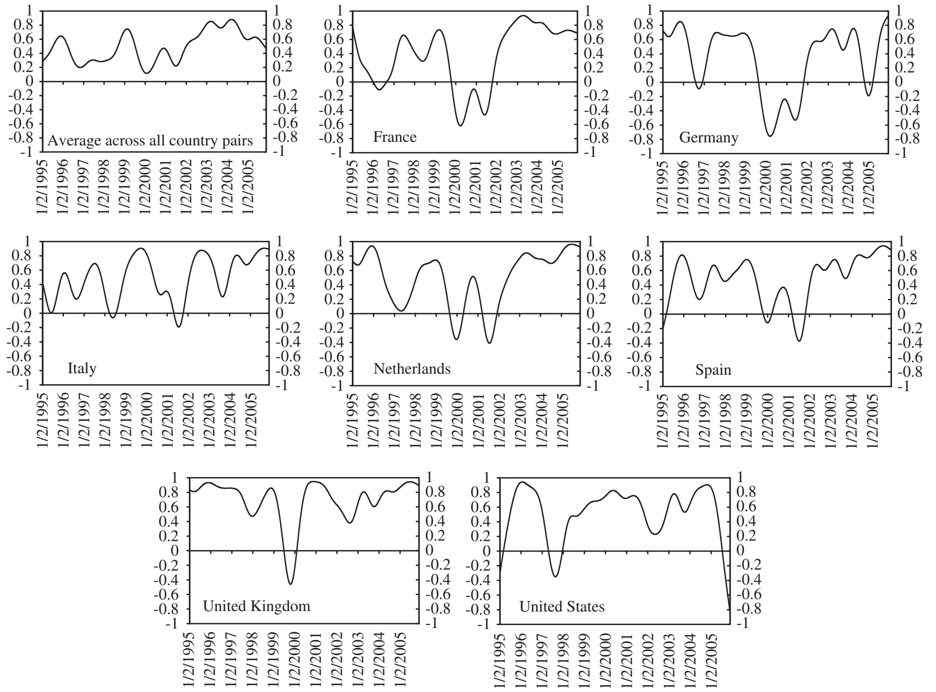
with the pre-euro period, and then increase to very high levels in the post–September 11th period.

Figure 3 shows the trends in one-year rolling correlations in the DD for all countries, including an average across all country pairs. The latter, which should be more indicative of global trends, seems to be increasing over time, being stronger in the 2003–04 period. It seems that the average pair-wise correlation has decreased in 1999 with the introduction of the euro, and has increased after 2001. It also seems that over the last couple of years of improving bank health the correlations seem to have diminished. The rolling correlations of Ireland and the United Kingdom seem to have been high during the whole period with the exception of 1999. The correlations of Ireland with France, Italy, and Spain seem to have increased over time. In general, the correlations of Ireland with continental Europe seem to have been negative during 1999–2001.

Occurrence of shocks, especially large shocks, in banks is captured by the weekly percentage change in the DD for each bank, $(\Delta dd_{it}/|dd_{it}|)$. The mean of the percentage change in the DDs is zero as expected, and the largest negative change is 237 percent, which represents a sizable shock. Figure 4 shows the trends in the percentage change in DDs for all countries. The most volatile period for Ireland was 1999–2000, which was also volatile for the United Kingdom and France as well. On the other hand, Italy, the Netherlands, Spain, and Germany experienced large shocks in 2001–03. Overall, for most countries there were fewer shocks in 2003–05 than in the decade before. This could reflect a general drop in credit events across the world mainly due to benign macroeconomic conditions and credit cycle for the last couple of years. For instance, FitchRatings (2005) reported a sharp drop in credit events to 37 in 2005 from 94 in 2003.

In order to describe how interdependencies across countries have evolved during the past decade, we try to capture the simultaneous occurrence of shocks across countries through simple correlations (Table 3) and the

Figure 3. Hodrick-Prescott (HP) Trend of One-Year Rolling Correlations of Distance to Default (DD)—Average of All Country Pairs, and between Ireland and Other Countries



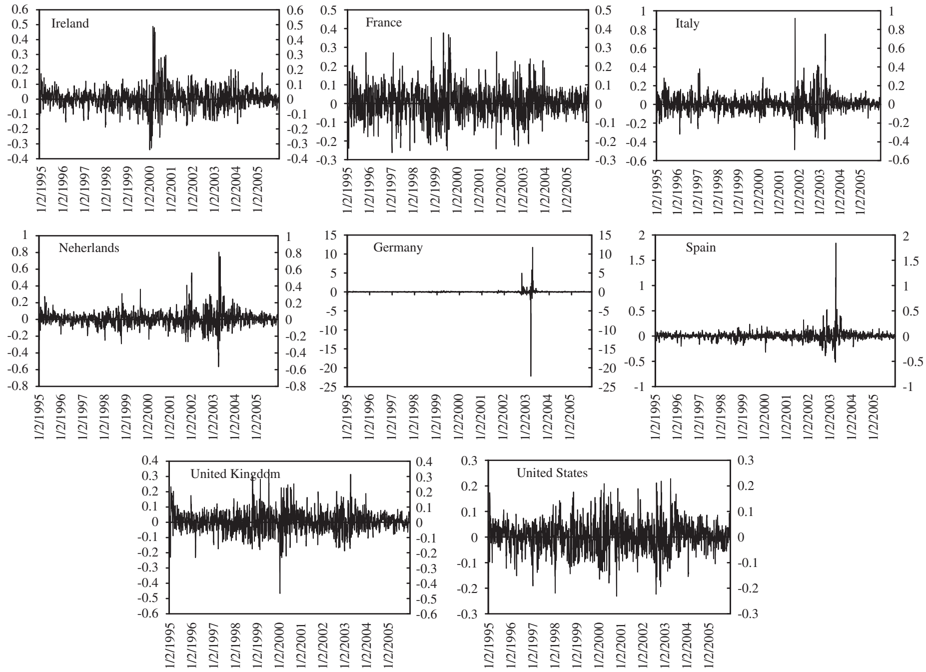
Note: The top left corner chart shows the HP-trend of 249-day rolling average correlations of DDs across all country pairs. The rest of the charts show the same for correlations between Ireland and other countries. The HP-filter uses a lambda of 6,812,100.

one-year rolling correlations (Figure 5) in the percentage change in DD, between Ireland and the other countries. On average, the correlations between the ($\Delta dd/|dd|$) between Ireland and the other countries are positive. They seem to have a stronger magnitude for the United Kingdom and the United States than for the other European countries. Unlike the correlations in the DD, the correlations in the ($\Delta dd/|dd|$) seem to be strongest in the pre-euro period, and weakest in the post-euro period (but still positive), suggesting that the joint improvement in the DDs in the past couple of years was accompanied by fewer simultaneous shocks. Over the past decade, the rolling correlations seem to be decreasing over time for Germany, the United Kingdom, and the United States, with no clear trend for the other countries. However, the average of the rolling correlations across all country pairs has a positive trend over the past decade, suggesting increasing global interdependencies.

III. Contagion Determinants: Multivariate Analysis Using Coexceedances

The number of banks in Ireland that experience a large shock on the same day as banks in other countries, after controlling for Ireland-specific and

Figure 4. Weekly Change in the Distance to Default (DD) by Country



Note: The charts show the five-day weekly percent change in the banks' aggregated (weighted by assets) DDs for each country. See Table 1 for the sample of banks from each country. The percent changes are expressed as fractions.

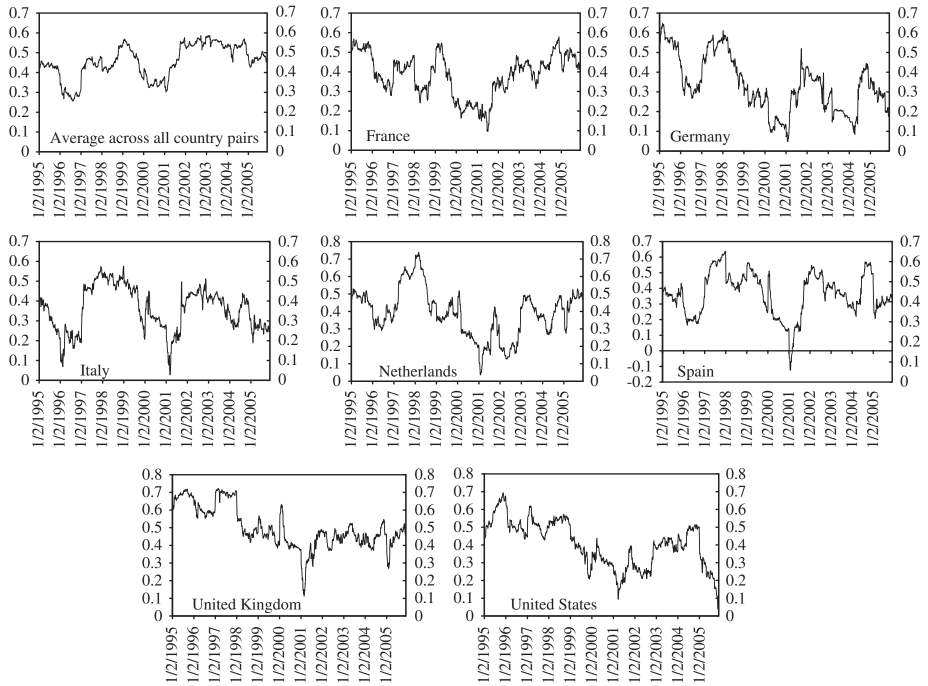
Table 3. Correlations in Changes in Distance to Default (DD)

Ireland and ...	1994–2005	Pre-Euro 1994–98	Post-Euro 1999–2001	Post–Sept. 11th 2001–05
France	0.32	0.43	0.17	0.42
Germany	0.08	0.45	0.17	0.12
Italy	0.31	0.36	0.26	0.37
Netherlands	0.31	0.47	0.23	0.32
United Kingdom	0.45	0.6	0.37	0.44
United States	0.36	0.5	0.28	0.34
Spain	0.27	0.41	0.19	0.33
Eastern Europe	0.02	0.03	0	0.26

Note: Cross-country correlations in banking systems' weekly changes in the DD. The sample covers January 3, 1994–November 25, 2005.

global shocks, is labeled as "coexceedances." We use an ordered logit model to estimate such coexceedances. Such methodology of extreme value theory to assess contagion risk was first proposed by Bae, Karolyi, and Stulz (2003) in the

Figure 5. Rolling Correlations of Weekly Percent Changes in Distance to Default (DD)—Between Ireland and Other Countries



Note: Each chart shows the one-year rolling correlations between Ireland and another country's weekly change in the banking systems' DD; the last chart shows the average of these correlations. The percent changes are expressed as fractions.

context of stock market returns in emerging markets. Bae, Karolyi, and Stulz (2003) and Gropp and Moerman (2004) show that it is useful to examine only the tails of the distributions of returns and of the DDs, as the distributions exhibit fat tails, and the correlation among the observations is substantially higher for larger shocks. We use weekly changes in the DD to examine whether shocks in one bank/banking system appear to influence the DD of other banks, controlling for common shocks affecting all banks simultaneously.

Methodology and Data

The dependent variable is the number of Irish banks simultaneously experiencing large shocks or tail events. We follow Gropp, Lo Duca, and Vesala (2005) in arguing that contagion is associated with extreme negative movements in bank's default risk. These events can be identified from the negative tail of the distribution of the changes in the DD. We define large shocks as the negative 15th percentile of the common distribution of the percentage change in the DD across all banks. We compute the "coexceedances" of banks in a given country as the number of banks in a given country that were simultaneously in the tail on the same day (Table 4).

Table 4. Data Description and Summary Statistics, January 3, 1994–November 25, 2005

Variable	Definition	Number of Observations	Mean	Median	Std. Dev.	Min	Max
Bank-specific							
dd_{it}	Distance to default (DD) of bank i at time t	123,304	5.19	5.00	2.33	−2.16	27.17
$\Delta dd_{it}/\lvert dd_{it-1}\rvert$	Percentage change in the DD	123,104	0.02	0.00	3.15	−236.53	895.32
Country-specific							
$\Delta dd_{ct}/\lvert dd_{ct-1}\rvert$	Percentage change in the DD of banking system c at time t						
	France	2,845	0.00	0.01	0.08	−0.26	0.38
	Germany	2,845	0.01	0.00	0.60	−22.25	11.80
	Ireland	2,845	0.00	0.00	0.07	−0.34	0.49
	Italy	2,845	0.01	0.01	0.09	−0.49	0.92
	Netherlands	2,845	0.01	0.01	0.08	−0.57	0.81
	Spain	2,845	0.00	0.00	0.09	−0.52	1.84
	United Kingdom	2,845	0.00	0.01	0.07	−0.47	0.35
	United States	2,845	0.00	0.01	0.06	−0.23	0.23
	Eastern Europe average	2,845	0.01	0.01	0.10	−0.68	0.95
Coexceedances IRL	Number of banks in the 15th percentile negative tail of $\Delta dd_{it}/\lvert dd_{it-1}\rvert$ in IRL	2,845	0.36	0.00	0.71	0.00	3.00
Coexceedances U.S.	Number of banks in the 15th percentile negative tail of $\Delta dd_{it}/\lvert dd_{it-1}\rvert$ in U.S.	2,845	1.76	0.00	2.88	0.00	14.00

Table 4 *(concluded)*

Coexceedances U.K.	Number of banks in the 15th percentile negative tail of $\Delta dd_{it}/\lvert dd_{it-1}\rvert$ in U.K.	2,845	0.78	0.00	1.24	0.00	5.00
Coexceedances GER	Number of banks in the 15th percentile negative tail of $\Delta dd_{it}/\lvert dd_{it-1}\rvert$ in GER	2,845	0.79	0.00	1.08	0.00	4.00
Coexceedances FR	Number of banks in the 15th percentile negative tail of $\Delta dd_{it}/\lvert dd_{it-1}\rvert$ in FR	2,845	0.33	0.00	0.64	0.00	2.00
Coexceedances SPA	Number of banks in the 15th percentile negative tail of $\Delta dd_{it}/\lvert dd_{it-1}\rvert$ in SPA	2,845	0.46	0.00	0.86	0.00	4.00
Coexceedances NL	Number of banks in the 15th percentile negative tail of $\Delta dd_{it}/\lvert dd_{it-1}\rvert$ in NL	2,845	0.32	0.00	0.59	0.00	2.00
Coexceedances IT	Number of banks in the 15th percentile negative tail of $\Delta dd_{it}/\lvert dd_{it-1}\rvert$ in IT	2,845	1.13	0.00	1.56	0.00	6.00

Memo items	**Cut-off point**
Negative 15th percentile of $\Delta dd_{it}/\lvert dd_{it-1}\rvert$	−0.070

Note: IRL refers to Ireland, U.S. to United States, U.K. to United Kingdom, GER to Germany, FR to France, SPA to Spain, NL to Netherlands, and IT to Italy. The banking system changes in DD for each country is calculated as follows: the DDs for all the banks are aggregated (weighted by total assets) to get the system-wide DD; the summary statistics of the change in the system-wide DD is reported above.

Table 5. Description of the Sample by Country

Country	Number of Banks	Maximum Number of Coexceedances	Coexceedances $=0$ Count	Coexceedances $=0$ Percent	Coexceedances $=1$ Count	Coexceedances $=1$ Percent	Coexceedances ≥ 2 Count	Coexceedances ≥ 2 Percent
Ireland	3	3	2,329	75.01	497	16.01	279	8.99
United States	14	14	1,616	52.05	498	16.04	991	31.92
United Kingdom	5	5	1,909	61.48	556	17.91	640	20.61
Germany	4	4	1,718	55.33	755	24.32	632	20.35
France	2	2	2,355	75.85	465	14.98	285	9.18
Spain	4	4	2,238	72.08	460	14.81	407	13.11
Netherlands	2	2	2,330	75.04	563	18.13	212	6.83
Italy	6	6	1,594	51.34	611	19.68	900	28.99
Total	40	40						

Note: A "coexceedance" refers to the simultaneous occurrence of an extreme shock in a number of banks. An extreme shock is defined by one when bank risk (weekly change in the DD) increases beyond a 15 percent threshold. This threshold is given by the 15th percentile left-tail cutoff of weekly changes in DD of the pooled sample of banks—when a bank's weekly change in the DD ≤ -0.07. Also refer to Table 4.

Given the discrete ordinal nature of the dependent variable, we use an ordered logit model to estimate the probability of a number of Irish banks simultaneously in the tail as a function of the number of banks in the tail in the other countries, controlling for Ireland-specific and global shocks:

$$Pr_c[Y=j] = \frac{e^{[\alpha_j F_c + \beta_j C_{c,t-1} + \gamma_j C_{d,t}]}}{\sum_k^J e^{[\alpha_k F_c + \beta_k C_{c,t-1} + \sum_{d \neq c} \gamma_{d,k} C_{d,t}]}},$$

where $j = 0, 1, 2$ represents the number of banks in the tail simultaneously ("coexceedances") in country c—0 if no banks are in the tail, 1 if one bank is in the tail, and 2 if two or more banks are in the tail. The vector F_c comprises Ireland-specific and global shocks affecting Ireland. $C_{c,\,t-1}$ is the lagged number of coexceedances in country c, and $C_{d,\,t}$ represents the coexceedances in period t in country d. Table 5 summarizes such coexceedances by country.

We perform three sets of estimations—the base model, the extended model, and the extended model for each of the three large Irish banks. The base model consists of a set of Ireland-specific and common global shocks, F_c, and persistence in Irish coexceedances, $C_{c,\,t-1}$. We use four variables to represent F_c:

(1) Systemic risk—we use a systemic risk indicator measuring the number of stock markets that are experiencing a large shock at time t. We

construct this control similarly to the modeling of large shocks in banks: we use an indicator variable that we set equal to one if a stock market of a given country experienced a shock large enough to be in the bottom 15th percentile of the distribution of the weekly change in returns, and zero otherwise. We calculate such an indicator variable for the euro area stock market index (N100 or Euronext Top 100 Index), the U.S. NASDAQ[8], and the Irish stock market (ISEQ or Irish Overall Index) indices, from Bloomberg. The systemic risk control is the sum of the indicator variables measuring whether or not the European stock market, the Irish stock market, and the U.S. stock market were in the tail on a given day, and it ranges from 0 to 3. This control should be positively related to the number of coexceedances in Ireland.

(2) U.S. stock market volatility—we use the weekly change in the volatility of the U.S. stock market to control for volatility spillovers from the United States. There is a large presence of U.S. software firms located in Ireland—nearly all of the largest U.S. software companies use Ireland as their primary base of operations for servicing markets in Europe, Africa, and the Middle East. Also, Irish banks are directly and indirectly exposed to the U.S. stock market through their securities and asset management businesses, and through their equity ownership, trade of CRT instruments, and wholesale funding activities. We estimated the volatility in NASDAQ using a GARCH(1, 1) model of the form $\sigma_t^2 = \alpha + \beta\sigma_{t-1}^2 + \gamma\varepsilon_{t-1}^2$, using maximum likelihood. The regression results are reported in Table 6.

(3) Irish stock market volatility—similar to above, the weekly change in the volatility of the ISEQ was calculated to account for Ireland-specific shocks.

(4) Interest rate shock—the fourth control includes the weekly change in the yield of the Irish 10-year government bond to reflect interest rate shocks.[9] We would expect the interest rate control to be positively related to the number of coexceedances.

The extended model has, in addition, the number of coexceedances in other countries in the sample—the United States, the United Kingdom, Germany, France, Spain, the Netherlands, and Italy. The model for each of the three main banks includes, in addition to the extended model regressors, coexceedances from each of the other two banks.

The summary statistics and the descriptive statistics for the number of coexceedances per country, that is, the number of banks simultaneously in the tail on a given day, are presented in Tables 1, 4, and 5. Even though the number of banks per country differs somewhat, there is at least one day on

[8]The Dow Jones Industrial Average gives similar results.

[9]This assumes that foreign interest rate shocks are reflected in the movement of Irish interest rates and in stock market movements.

Table 6. Results from GARCH (1, 1) Model:

$$\sigma_t^2 = \alpha + \beta\sigma_{t-1}^2 + \gamma\varepsilon_{t-1}^2$$

	Coefficient	Standard Error	Z-stat	p-value
U.S. stock market volatility—NASDAQ				
Const.	7.52E-05	6.54E-06	11.5	0
ε_{t-1}^2	0.65	0.03	16.97	0
σ_{t-1}^2	0.38	0.01	24.02	0
Ireland stock market volatility—ISEQ				
Const.	7.79E-05	4.46E-06	17.46	0
ε_{t-1}^2	0.67	0.04	15.34	0
σ_{t-1}^2	0.22	0.02	12.52	0

Note: "U.S. stock market volatility—Nasdaq" ("Ireland stock market volatility—ISEQ") implies the weekly change in the volatility of the U.S. stock market or NASDAQ (Irish stock market, ISEQ), estimated using a GARCH(1, 1) model using maximum likelihood, where σ_t^2 refers to volatility of weekly stock returns.

which all banks experienced a large adverse shock simultaneously. We limit the number of outcomes to 0, 1, and 2 or more coexceedances.[10] The time pattern of coexceedances in Ireland is presented in Figure 6. The most turbulent period for Irish banks has been the period 1999–2000, and the calmest being the mid-1990s and the past couple of years.

Econometric Model Results

The results for the basic and the extended models are given in Table 7. The dependent variable is the number of banks whose weekly percentage change in the DD was in the 15th percentile negative tail in a given day. In all countries with more than two banks, we limit the model to estimating three outcomes: 0, 1, and 2 or more banks simultaneously in the tail.

The basic regression results suggest that the probability of Irish banks being simultaneously in the bottom tail varies positively with systemic risk—real shocks experienced through large stock market movements in Ireland and elsewhere increase the probability of a higher number of Irish banks being simultaneously in the tail. However, controlling for global shocks and including that of the ISEQ, Irish coexceedances still respond positively and significantly to changes in a long-term Irish interest rate. Long-term lending

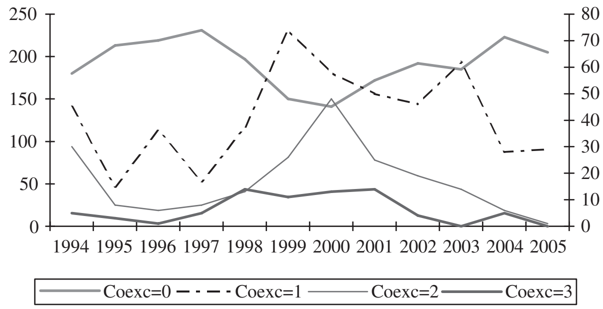
[10]The number of coexceedances depends on the number of banks included in the sample and may not necessarily reflect the strength of the banking system per se. Still, comparing countries with equal number of banks in the sample suggests that Spanish banks tend to experience fewer shocks compared to German banks and that Dutch banks tend to be somewhat less frequently subject to shocks compared to French banks.

Figure 6. Time Pattern of Coexceedances in Ireland, January 3, 1994—November 25, 2005
(In number of days each year)

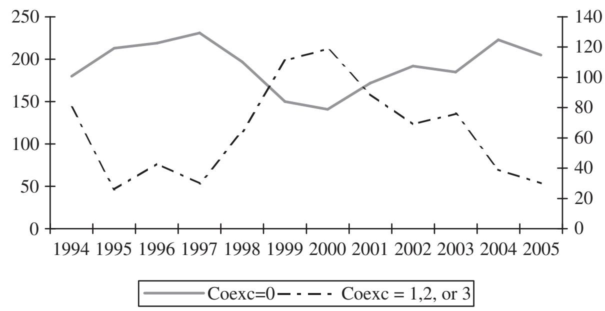
Ireland

Period	Number of coexceedances 0	1	2	3	1+2+3
1994	180	45	30	5	80
1995	213	15	8	3	26
1996	219	36	6	1	43
1997	231	17	8	5	30
1998	197	37	13	14	64
1999	150	74	26	11	111
2000	141	58	48	13	119
2001	172	50	25	14	89
2002	192	46	19	4	69
2003	185	62	14	0	76
2004	223	28	6	5	39
2005	205	29	1	0	30

Ireland: Number of Coexceedances per Year



Ireland: Number of Coexceedances per Year



Note: The figure shows the number of days in each year, "n" ($n = 0, 1 \ldots 3$) banks simultaneously experienced an extreme shock ("coexceedance"). The extreme event is defined by one when bank risk (weekly change in the DD) increased beyond a 15 percent threshold. This threshold is given by the 15 percent left-tail cutoff of weekly changes in DD of the pooled sample of banks—when a bank's weekly change in the DD ≤ -0.07.

rates (especially mortgage rates) would likely follow movements in long-term government bond yields. Given the large exposure of Irish banks to the real estate market and the prevalence of variable loan rates, sudden increases in

Table 7. Ordered Logit Model: Contagion in Daily Coexceedances of the Weekly Change in Distances to Default (DD), January 3, 1994–November 25, 2005

(Dependent variable: number of Irish banks simultaneously in the tail "coexceedances")

	Ireland				AIB				Anglo IB				BoI			
	Basic model		Extended model		Basic model		Extended model		Basic model		Extended model		Basic model		Extended model	
	Coeff.	St. error	Coeff.	St. error	Coeff.	St. error	Coeff.	St. error	Coeff.	St. error	Coeff.	St. error	Coeff.	St. error	Coeff.	St. error
Coexceedances own lagged	2.16**	0.07	2.08**	0.08	3.36**	0.14	3.11**	0.16	3.51**	0.14	3.36**	0.15	3.07**	0.14	2.77**	0.15
Systemic risk	0.81**	0.06	0.73**	0.06	0.87**	0.08	0.79**	0.09	0.62**	0.08	0.52**	0.09	0.81**	0.08	0.67**	0.08
Volatility ISEQ	−33.96	34.04	−34.41	33.97	37.56	53.89	11.23	55.89	5.9	47.39	−20.08	47.99	-19.56	41.21	−44.7	40.75
Volatility U.S.	−18.66	11.74	−19.42*	11.52	−1.92	17.4	−0.96	16.96	−10	16.1	−10.29	16.03	−28.82*	16.67	−35.07**	15.91
Interest rate IRL	2.18**	0.41	1.92**	0.41	3.24**	0.58	3.03**	0.59	1.53**	0.59	1.15**	0.59	2.45**	0.54	2.03**	0.55
Contagion FR			0.11	0.08			0.12	0.11			−0.02	0.11			0.13	0.11
Contagion GER			0.01	0.07			−0.04	0.1			−0.02	0.1			0.006	0.09
Contagion IT			0.002	0.06			−0.13	0.09			0.05*	0.09			−0.05	0.09
Contagion NL			−0.21**	0.09			−0.36**	0.13			−0.17	0.13			0.019	0.12
Contagion SP			−0.04	0.07			0.21**	0.11			−0.07	0.11			−0.14	0.11
Contagion U.K.			0.13**	0.069			0.17*	0.09			0.15	0.1			0.06	0.09
Contagion U.S.			0.27**	0.06			0.20**	0.08			0.23**	0.08			0.38**	0.08

Table 7 *(concluded)*

AIB											0.03	0.2			0.51**	0.17
Anglo IB							0.19	0.19							0.34**	0.17
BoI							0.41**	0.18			0.39**	0.19				
Pseudo-R^2		0.31		0.32		0.393		0.408		0.351		0.362		0.327		0.347

Note: **, * indicate statistical significance at the 5 and 10 percent level respectively. "Coexceedances own lagged" refers to one lag of the dependent variable. "Systemic risk" measures the number of stock markets that are experiencing a large shock at time t. It is constructed from indicators of three stock markets: the Euro Area stock market index (N100 or Euronext Top 100 Index), the U.S. NASDAQ, and the Irish stock market (ISEQ or Irish Overall Index) indices. Each indicator is set equal to one if a stock market of a given country experiences a shock large enough to be in the bottom 15th percentile of the distribution of the weekly change in returns, and zero otherwise. Systemic risk is the sum of the indicator variables and it ranges from 0 to 3. "Volatility U.S." ("Volatility ISEQ") implies the weekly change in the volatility of the U.S. stock market or NASDAQ (Irish stock market, ISEQ), estimated using a GARCH(1, 1) model of the form: $\sigma_t^2 = \alpha + \beta\sigma_{t-1}^2 + \gamma\varepsilon_{t-1}^2$, using maximum likelihood, where σ_t^2 refers to volatility of weekly stock returns (also see Table 8). "Interest rate IRL" is the weekly change in the yield of the Irish 10-year government bond. "Contagion …" refers to one lag of coexceedances in France (FR), Germany (GER), Italy (IT), Netherlands (NL), Spain (SP), United Kingdom (U.K.), United States (U.S.). The extreme movements in risk of the three Irish banks are referred to as AIB (Allied Irish Bank), Anglo IB (Anglo Irish Bank), and BOI (Bank of Ireland). The goodness of fit in LOGIT (and other binary) models is given by the "Pseudo R_2". This statistic is the likelihood ratio index, computed as

$$R^2 = 1 - l(\tilde{\beta})/l(\bar{\beta}),$$

where $l(\beta)$ is the restricted log likelihood or the maximized log likelihood value when all slope coefficients are restricted to zero, which is equivalent to estimating the unconditional mean probability of an observation being in the tail.

interest rates could be associated with credit events that might have a negative impact on banks. In addition, the notion that the number of coexceedances could be sticky is supported: the lagged (by one day) number of coexceedances is positive and significant. The stock market volatility controls are not significant, suggesting that the systemic risk variable might be sufficiently capturing stock market spillover effects.

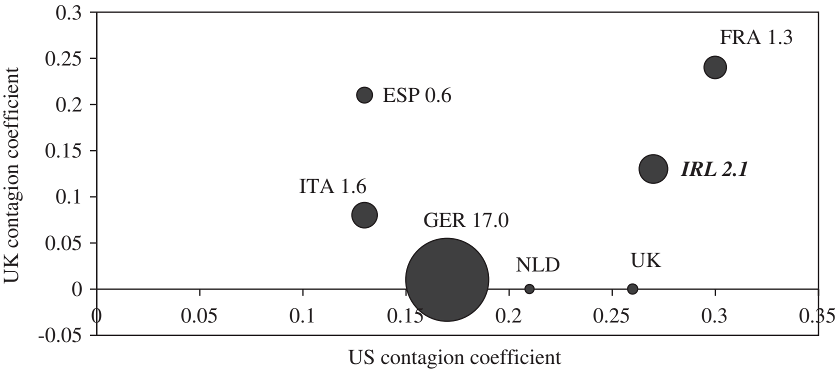
Next, we extend the model to include contagion or coexceedances from other countries (Table 7). We measure contagion by including the first lag of the coexceedances in the other seven countries. If, after controlling for common shocks, any of these variables turn out to be significant, we interpret this as contagion from that country. We find evidence of contagion from the United Kingdom, the United States, and the Netherlands (with a negative sign) toward Ireland. Adding foreign coexceedances adds information to the specification, which is reflected in the fact that the significance of the controls remains largely unchanged.

How does the size of the contagion from the United States to the Irish banks compare with that from the United States to other countries? Although the coefficient estimates should not be interpreted as marginal effects, the relative effect of one variable compared with another can be gauged by the relative sizes of the coefficient estimates. To get the relative U.S. effect, we benchmark the U.S. influence with that of the United Kingdom. That is, the relative size of the U.S. influence is the coefficient for the United States divided by the coefficient for the United Kingdom. For the Irish banks, the relative U.S. influence on the three Irish banks is (0.27/|0.13|) 2.1 (from Table 7). Next we run similar regressions for each of the other countries (not reported here) and calculate the relative U.S. influence where both the U.S. and the U.K. influences are significant. Figure 7 shows a bubble chart for the four euro area countries with significant coefficients for both the United States and the United Kingdom. The relative contagion from the United States to the Irish banks is second only to Germany.

The final set of estimations involves estimating the extended model for each of the three individual Irish banks (Table 7). There is evidence of contagion from the United Kingdom (at the 10 percent significance level), the United States, Spain, and the Netherlands (negative coefficient) toward AIB; contagion from Italy (at the 10 percent significance level) and the United States toward Anglo IB; and from the United States toward BoI. We find evidence of a two-way contagion between AIB and BoI, and between Anglo IB and BoI—these could reflect both interbank linkages and off-balance sheet or derivative positions where one bank is the buyer and the other the seller of various risk protections.

The results for the extended model qualitatively survive several robustness checks. First, when the number of U.S. banks was reduced from 14 to 5 (the top 5), the U.S. influence on Irish banks remained the same. Second, taking monthly changes in DD, rather then weekly changes, does not change the U.S. and the U.K. influence; the contagion from the Netherlands disappears. Third, changing the threshold that defines the negative tail of the

Figure 7. Relative Size of Contagion from the United States to Selected European Countries



Note: The black patches show the relative size of contagion from the United States (U.S.), benchmarked against contagion from the United Kingdom (U.K.). For example, for Ireland (IRL) it is $0.27/|0.13| = 2.1$, derived from Table 7 ("extended model for Ireland" column). The patches for the other countries—France (FRA), Germany (GER), Italy (ITA), the Netherlands (NLD), Spain (ESP)—are derived from similar regressions that are not reported in the paper. For Italy, the Netherlands, and the United Kingdom, only the U.S. coefficient is shown, because the U.K. coefficient was not significant for Italy and the Netherlands, and not applicable for the United Kingdom.

distribution of large shocks does not qualitatively change the results for the United States and the United Kingdom. Finally, looking at the pattern of contagion risk over time, we find different linkages in the pre-euro, post-euro, and the post–September 11th periods (Table 8). We find evidence of contagion from the United Kingdom and Germany (at the 10 percent significance level) in the pre-euro period; from the United States and the United Kingdom (at the 10 percent significance level) in the post-euro period; and only from the United States (at the 10 percent significance level) to Ireland in the post–September 11th period. This suggests a changing pattern of contagion risk, from stronger linkages with Europe in the earlier periods to stronger linkages with the United States later on, consistent with the evidence presented above.

IV. Summary and Conclusions

This paper examines the external linkages of the Irish banking sector and estimates an indicator of potential contagion risk—arising from idiosyncratic shocks in other countries—that the three major Irish banks may be exposed to. Aggregate balance sheet data of Irish-resident banks suggest several channels of external interdependencies, including foreign equity exposures, loan-book exposures abroad, and wholesale funding through interbank and capital market issues. However, apart from these links, there could also be foreign exposures through CRTs (for example, Irish banks selling risk protection to cross-border banks that could be subject to credit events in

Table 8. Ordered Logit Model: Contagion in Daily Coexceedances of the Weekly Change in Distances to Default (DD) by Subsample

(Dependent variable: number of Irish banks simultaneously in the tail "coexceedances")

	Pre-Euro Period (Jan. 1994–Dec. 1998)		Post-Euro Period (Jan. 1999–11 Sept. 2001)		Post–September 11 Period (12 Sept. 2001–Nov. 2005)	
Ireland	Coeff.	St. error	Coeff.	St. error	Coeff.	St. error
Coexceedances IRL lagged	1.73**	0.14	2.20**	0.14	2.05**	0.14
Systemic risk	1.17**	0.17	0.46**	0.12	0.76**	0.1
Volatility ISEQ	–120.12*	65.56	91.05	99.59	–27.86	60.73
Volatility U.S.	33.86	48.71	–20.65*	12.81	–12.98	39.03
Interest rate IRL	2.52**	0.61	1.50*	0.91	–0.15	0.94
Contagion FR	0.12	0.14	–0.01	0.15	0.18	0.18
Contagion GER	0.22*	0.12	0.08	0.12	–0.14	0.13
Contagion IT	0.06	0.11	0.005	0.11	–0.002	0.12
Contagion NL	–0.24	0.16	–0.14	0.17	–0.03	0.17
Contagion SP	–0.01	0.13	–0.06	0.14	0.01	0.15
Contagion U.K.	0.29**	0.12	0.20*	0.12	–0.19	0.13
Contagion U.S.	0.14	0.1	0.28**	0.11	0.18*	0.11
Pseudo-R^2		0.339		0.322		0.286

Note: **, * indicate statistical significance at the 5 and 10 percent level, respectively. "Coexceedances IRL lagged" refers to one lag of the dependent variable. "Systemic risk" measures the number of stock markets that are experiencing a large shock at time t. It is constructed from indicators of three stock markets: the Euro Area stock market index (N100 or Euronext Top 100 Index), the U.S. NASDAQ, and the Irish stock market (ISEQ or Irish Overall Index) indices. Each indicator is set equal to one if a stock market of a given country experiences a shock large enough to be in the bottom 15th percentile of the distribution of the weekly change in returns, and zero otherwise. Systemic risk is the sum of the indicator variables and it ranges from 0 to 3. "Volatility U.S." ("Volatility ISEQ") implies the weekly change in the volatility of the U.S. stock market or NASDAQ (Irish stock market, ISEQ), estimated using a GARCH(1, 1) model of the form: $\sigma_t^2 = \alpha + \beta\sigma_{t-1}^2 + \gamma\varepsilon_{t-1}^2$, using maximum likelihood, where σ_t^2 refers to volatility of weekly stock returns. "Interest rate IRL" is the weekly change in the yield of the Irish 10-year government bond. "Contagion …" refers to one-lag of coexceedances in France (FR), Germany (GER), Italy (IT), Netherlands (NL), Spain (SP), United Kingdom (U.K.), United States (U.S.). The extreme movements in risk of the three Irish banks are referred to as AIB (Allied Irish Bank), Anglo IB (Anglo Irish Bank), and BOI (Bank of Ireland). The goodness of fit in LOGIT (and other binary) models is given by the "Pseudo R^2". This statistic is the likelihood ratio index, computed as

$$R^2 = 1 - l(\tilde{\beta})/l(\bar{\beta}),$$

where $l(\beta)$ is the restricted log likelihood or the maximized log likelihood value when all slope coefficients are restricted to zero, which is equivalent to estimating the unconditional mean probability of an observation being in the tail.

other countries), and through operational risks that are difficult to measure but can quickly lead to large fluctuations in bank stock prices.

Because Irish-resident banks include a large number of foreign banks operating in the FSC (that have limited Irish linkages in the retail market), we focus on the three major listed banks—BoI, AIB, and Anglo IB—who have nearly 80 percent of the domestic retail market. We proxy banking risk by DD measures constructed from bank equity prices and look at correlations of changes in the DD of Irish banks with banks in other countries. Following Gropp, Lo Duca, and Vesala (2005), we then define large changes in DD of each bank—coexceedances—if the changes fall below the 15 percentile of the negative tail of the joint distribution of the DD changes across all banks of the sample. We use an ordered logit model to estimate the probability of the number of Irish banks being in the tail at the same time as banks in other countries, controlling for Irish-specific and global factors.

There is evidence suggesting contagion risk from the United States and the United Kingdom to Ireland, although the size of the U.S. influence dominates that of the United Kingdom in almost all regressions and across many robustness checks.[11] There are obvious balance sheet linkages of Irish banks with the United Kingdom through subsidiaries. The United States and the United Kingdom also remain the major countries selling risk protection to Irish banks. Aggregate data show that Irish-resident banks (but not necessarily the three banks in the empirical study) have positive net-asset exposures in the United States, that could give rise to credit risk from exposure to nonbank assets and liquidity risk from Irish banks' (net) interbank borrowing from the United States, from mid-2004. On the other hand, Irish resident banks have very large and *negative* net-asset exposure to the United Kingdom, suggesting that risks in Irish banks' on-balance-sheet exposures to the United Kingdom might have been mitigated by off-balance-sheet-risk protection bought from U.K. banks. Still, both the United States and the United Kingdom have had booming property markets—Irish banks are exposed to them and could be affected in the event of a substantial downturn in any of these markets.

Some tentative policy lessons could be drawn from the results of this exercise. The Central Bank and Financial Services Authority of Ireland may want to stress test specific categories of exposures of Irish banks to both the United States and the United Kingdom. Even though linkages with the United States do not come out strongly from *aggregate consolidated* balance sheet exposures, there might be derivatives or other off-balance-sheet exposures that the bank supervisors may need to be vigilant of. The Irish

[11]Econometric estimates (not reported) suggest that while there is no evidence of contagion from Europe to the United States, the United States is a source of contagion not only for Ireland but also for all the other European countries. Germany seems to be mainly a receiver of contagion risk rather than a source, which is consistent with the evidence from BIS that Germany is mainly "a buyer" of risk.

authorities may need to collect more information about types and counterparties of derivative positions and risk transfers through structured products of Irish banks, as the use of these is likely to grow rapidly in the future.[12] This would especially be necessary if Irish banks are buying CRT products from foreign banks (that is, selling risk protection) that are in turn exposed to property markets or other loan products in the United States or the United Kingdom—thus exposing the Irish banks to these markets even though there is no direct loan exposure.

Finally, some caveats apply to the econometric results. We are using equity prices of banking *groups*, and yet mainly discussing international links typically associated with banking without explicitly stating links through securities and insurance (the top two banks are also involved in insurance and securities business). We are using data available for only a small group of listed banks in each country—so it may not be representative of the system, and the banks in Ireland comprise less than 50 percent of the system in terms of total assets (although nearly 80 percent of the retail market). The number of observations experiencing large shocks is low; the results could therefore be driven by the large shocks associated with the tech-bubble burst of 2000.

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Foreign Aid with Voracious Politics

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Fractious domestic politics are at the root of continued poverty in some developing countries and pose a dilemma for donors and international financial institutions. This paper examines the effects of foreign assistance in countries with plentiful investment opportunities when interest groups compete for unproductive government transfers. We assess conditional and unconditional assistance (project and program aid, loans, and grants). We find that project conditionality alone may fail to spur growth. Official development loans channeled to investment may not increase the recipient's growth and welfare even if interest groups are unable to appropriate aid funds directly. Conditions must tackle the domestic drivers of inefficient fiscal policies. To improve the composition of government expenditure, increase growth, and improve welfare, tax rates must be kept constant and loan repayment be financed by cuts in unproductive transfers. Official development grants are superior to loans of the same net present value if donors cannot enforce conditions on assistance.
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The economic development of some countries is greatly hampered by international credit constraints and by fractious domestic politics. Lack of access to international capital causes under-provision of productive public investments in health, education, and infrastructure even in countries with perfect political systems. Flawed political institutions compound the problem.

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When central governments are weak and public spending decisions reflect the political influence of organized constituencies rather than the general public's interests, patterns of government spending are distorted and economic performance suffers. If excessive appetite for rents dominates public spending decisions, government resources are diverted from productive activity toward rent-seeking. In such circumstances, foreign transfers and other potentially favorable developments may actually cause recipient countries to be worse off—a phenomenon Tornell and Lane (1999) termed "the voracity effect."

In this paper, we examine the economic effects of international assistance (loans and grants) to a credit-constrained country in which interest groups compete for government transfers as in Tornell and Lane.[1] We find that loans conditioned on public investment alone—that is, loans directed exclusively at investment—will not necessarily increase growth or welfare in the recipient country, even when conditions can be perfectly enforced. Thus project conditionality may fail to produce the desired results in countries ravaged by voracious politics. However, extending conditionality to take into account the underlying drivers of a country's fiscal policy can help reduce distortions in the composition of government spending and increase growth and welfare. This prediction is roughly consistent with the aid patterns historically. Broadly, speaking, international development loans were initially conditioned on the implementation of individual projects. Over time, as the limits of project conditionality became apparent, multilateral and bilateral donors expanded the scope of their conditions. The budget support loans used by donors focus on reforms in the recipients' domestic fiscal policies, including of the level and composition of spending, public expenditure management systems, and reform of tax policies and administration.

In our application of Tornell and Lane's idea, organized interest groups devote time to "producing" transfers from a central government. These groups allocate their resources (time, effort, and money) to what Bhagwati (1980, 1982) and Bhagwati and Srinivasan (1982) have dubbed directly unproductive activities: protests, strikes and lobbying aimed at changing laws and regulations, generating unproductive government employment, and obtaining national funds for unproductive local projects.[2] These unproductive activities

[1]Tornell and Velasco (1992) and Amador (2008) also set up economic environments in which domestic interest groups demand too much spending from the government because the costs of spending are spread over the entire economy. Interestingly, Amador shows that it is in the interest of an "over-spending" country to repay its foreign debt. In our paper, we simply assume that foreign debts are repaid. We then examine the consequences for debt repayment on rent-seeking and the productivity of investment project funded by the loans.

[2]In their analysis of how governance problems reduce economic growth, Kaufman, Mastruzzi, and Zaralita (2003) discuss the key sources of patronage politics, several of which relate to fractious political environments. In such environments, politicians face heightened incentives to maintain and expand support from opposing groups through offers of public jobs and other perks. Often coalition governments are formed based on explicit agreements for sharing state patronage. Patronage politics motivates the need for tax and expenditure and civil service reforms in many developing countries. See Agénor (2004, pp. 592–94).

lower current output and the future returns of public investments. We consider a country that is poor for these reasons and, additionally, because it is confronted with international borrowing constraints.

The presence of directly unproductive activities complicates the ability of international assistance to mitigate the effects of borrowing constraints. We show that development loans that firmly condition, dollar-for-dollar, on productive public spending on investment can fail to increase growth and welfare. A transfer paradox of sorts emerges: the weak central government faces pressure to maintain transfers to interest groups and to repay loans by raising future tax rates. In political equilibrium, interest groups have incentives to further divert time away from productive activity and toward the "production" of government transfers. The result is lower overall welfare in the economy, because the return on investment is lowered and may not even exceed loan repayment. A favorable exogenous policy shock, loans from international donors, results in lower welfare and growth!

One way to prevent the voracity effect is to extend the conditions of international loans to tackle the underlying distortions in the country's public expenditure policies. The guiding principle is that the repayment of debt should not favor rent-seeking over productive activity. Toward this end, loan conditions should be extended to require that (1) *future* tax rates be maintained and (2) loan repayments be financed by cuts in those government spending categories that are particularly susceptible to corruption and rent-seeking. Reducing the incentive to seek rents will increase productive activity and guarantee that improvements in growth and welfare materialize. While this principle is clear in theory, identifying the appropriate cuts in government spending may not be obvious to the donors in reality. For example, corrupt public "investment" projects can be used as a vehicle for making transfers to favored groups (Keefer and Knack, 2007). Knowing which spending to cut requires detailed knowledge of the country's politics. For this reason, we stress that the cooperation of the recipient country's finance minister, backed by sound technical, cost-benefit analysis of projects, and with input from other stakeholders, is crucial in establishing the appropriate fiscal conditions.

In practice, detailed fiscal conditions of the sort advocated here have long formed an important part of the conditionality featured in programs of adjustment and reform supported by the International Monetary Fund (IMF) and the World Bank. Initially, the conditions imposed by the international financial institutions (IFIs) focused on broad macroeconomic, price and trade-related reforms, such as ceilings on the rate of expansion of domestic credit and the budget deficit, and liberalization of prices and trade. Detailed fiscal conditions reflect the reality that in many low-income countries, the composition of government spending is distorted, reflecting the political influence of powerful interest groups. In the absence of detailed fiscal conditionality, the composition of government spending in many low-income countries would be distorted away from productive spending, including much-needed investments in human and physical capital. A cap on the overall budget deficit that did not take aim at the bias in favor of unproductive spending and transfers to politically connected interest groups might succeed in meeting short-run stabilization objectives but would fail to raise growth in the recipient. Empirical work finds that fiscal adjustments that require changes in the composition of spending in low income countries, toward investment and away from consumption, are important for growth (Baldacci, Clements, and Gupta, 2003, survey research on this topic).

In response to lackluster growth in some countries receiving IFI assistance, the focus of conditionality gradually shifted away from short-term macroeconomic, trade and price adjustment, to encompass more complex structural fiscal, financial, and corporate sector issues. In particular, starting in the mid-1980s, IFIs began to routinely sponsor comprehensive public expenditure reviews (PERs) that scrutinized the efficiency of various components of public spending against the marginal cost of raising funds, including through borrowing. These PERs typically advocate reallocation of spending away from categories favoring powerful interest groups and toward deserving but under-represented constituencies. The most important recommendations of these reviews eventually become conditions of IFI loans.[3] In recent years, governments and the international community have redoubled their efforts to improve the mix of public spending in order to meet the United Nations Millennium Development Goals (MDGs) by 2015. With aid flows projected to increase markedly, and with many aid recipients still falling short of achieving several of the MDGs, there has been increased analytical scrutiny of the level and efficiency of different categories of public spending. At issue is the appropriate level, composition, and phasing of aid-financed spending between investment in basic infrastructure (such as roads, energy, and irrigation) and spending on social sectors. Quantitative models, such as the World Bank's Maquette for MDG Simulations (MAMS) focus on this distinction, as do econometric examinations of the impact of different types of aid.[4]

Our paper formalizes the mechanism through which public spending gets distorted and provides a welfare justification of the need for detailed fiscal conditions. In our model, as well as in practice, detailed conditions aim at improving the composition of government spending and reducing the resources devoted to directly unproductive activities. Such detailed conditions may have a powerful effect on policies so long as the domestic political economy is not too adverse. In some situations, however, the

[3]Examples are the public sector reform loans of the World Bank and the loans granted under the IMF's Poverty Reduction and Growth Facility. See, for example, World Bank (2006).

[4]On the evolving focus of World Bank conditions, see Koeberle and Malesa (2005, p. 52). On the trends in structural conditionality in IMF-supported programs, see IMF (2001, pp. 8–13) On the appropriate level and mix of different types of aid-financed spending to meet the MDGs, see Sundberg and Lofgren (2006, pp. 147–50). See also Clemens, Radelet, and Bhavnani (2004). Again, the sensible strategy of encouraging public infrastructure investment is subject to the proviso that corrupt "investment" projects can also be the sources of pure transfers to interest groups and of unproductive government consumption.

domestic political system is so configured that external carrots and sticks cannot have decisive impact. In such environments, lending and aid selectivity is the appropriate response of altruistic donors.

I. The Environment

There are m groups of households. As in Tornell and Lane (1999), the groups may be provincial governments, unions, industry advocates, or communities that have political connections with government officials. Each group is represented by a single household-type. For notational simplicity, we assume that there is just one household per group. While the households representing the different groups differ politically and compete with each other over government transfers, they are identical in terms of preferences, productivity, and the ability to generate transfers.[5]

Each household lives for two periods and is endowed with one unit of time each period. The productivity of household time (h) is determined by public investments carried out by the central government (roads, communications infrastructure, schools, and public health provisions). Productivity in the current period (h_1) is given (based on past investments), but future productivity (h_2) is determined by current period investment decisions made by the central government (x). We assume $h_2 = x^{\theta}$, with $\theta < 1$. In each period, households choose how much time to devote toward productive activity (n) and how much to devote toward procuring government transfers ($1-n$), which for convenience we shall call "rent-seeking."[6] Time devoted to work generates net income equal to $(1-\tau)whn$, where τ is the income tax rate and w is the rental rate for human capital.[7]

[5]Our setup is similar to Tornell and Lane (1999) in that there are m different interest groups that compete in a Nash game for public funds. Beyond this basic assumption the details of the two approaches are different. Tornell and Lane focus on taxing physical capital to finance transfers. Agents have the option to avoid the tax by investing in less productive, but untaxed, nonmarket firms. In equilibrium the net-of-tax return to market and nonmarket investments must be equated. The voracity effect results because any increase in the productivity of market investments must be met by a perfectly offsetting rise in taxes and government transfers, so as to maintain equality in rates of return across market and nonmarket sectors.

[6]One could generate similar results with a model of labor-leisure choice, rather than focusing on the choice between productive work and rent-seeking. However, rent-seeking activity seems to be a more important source of poverty in low-income countries, where output *per worker* is very low, than is leisure demand. For a given level of recorded employment, the greater the fraction of the employment that is allocated to unproductive activities, the lower would be average worker productivity. The rent-seeking model also produces sharper predictions because it is a pure time allocation model (between two income generating activities) and therefore includes no wealth effects that make predictions about the labor-leisure choice theoretically ambiguous.

[7]Many poor countries have difficulty collecting taxes on wages and instead rely on taxing physical capital. However, introducing capital would not add any new insights to the analysis. In a small open economy, all capital taxes are passed on to labor in the form of a lower pre-tax wage rate. For our purposes, this is equivalent to taxing wages directly.

We assume that the economy is small and open to private physical capital flows, but that international private loans for human capital investment are unavailable. Goods are produced in a single sector with a Cobb-Douglas technology and factor markets are competitive. The economy's rental rates on human and physical capital are $w=(1-\gamma)k^{\gamma}$ and $r=\alpha k^{\alpha-1}$, where γ is the capital share, k is the ratio of physical to human capital in production, and r is the international rental rate on physical capital.

Activities that generate government transfers are lobbying, legal actions, unproductive government employment, and efforts to obtain national funds for unproductive local projects. We assume the technology for generating transfers (T) is $T=\phi(1-n)^{\phi}h$, where the parameters satisfy $0<\phi_0$ and $0<\phi<1$. There is diminishing marginal productivity associated with devoting time to rent-seeking, but human capital raises productivity in work and in rent-seeking proportionally. We make this assumption because we know of no evidence, casual or otherwise, suggesting that education affects the productivity of work differently from the productivity of rent-seeking.[8,9] Even when interpreting rent-seeking as unproductive public employment, education "credentials" could increase the size of the transfer ("wage") or more educated local official could better secure national government funding for local public works employment.

The period budget constraint of the household is given by

$$c = (1-\tau)whn + T \tag{1}$$

Household preferences are given by

$$\ln c_1 + \beta \ln c_2 \tag{2}$$

where $\beta > 0$ is the household's constant time discount factor. Households also take account of the government budget constraint in period $i = 1, 2$,

$$\sum_{j=1}^{m} (T_{ij} + e_{ij}) = \tau w h_i \sum_{j=1}^{m} n_{ij}, \tag{3}$$

where e_{ij} represents public expenditures per group member, T_{ij} is the transfers to group j, and n_{ij} is the productive work effort of group j. We

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[8]A more general specification would allow rent-seeking to be a function of the size of the government's budget. Here we focus on the case where international assistance is allocated, dollar for dollar, to raise public investment, which implies that the discretionary budget is unaffected by outside loans. We show that even in this idealized setting, development loans may fail to improve welfare.

[9]Easterly (2001, p. 82) argues that increased education will not lead to increased production when the incentives are not right. "One clue as to why education is worth little more than hula hoops to a society that wants to grow comes from what educated people are doing with their skills. In an economy with extensive government intervention, the activity with the highest returns to skills might be lobbying the government for favors. In an economy with many government interventions, skilled people opt for activities that redistribute income rather than activities that create growth."

assume that the central government has the authority to choose e (or, equivalently, that all groups can agree and coordinate on some basic public expenditures).

II. Autarky

We assume throughout that the central government (1) chooses first period public investment to maximize the welfare of each household and (2) *might* like to borrow funds at the international interest rate (r), but cannot do so without going through official creditors or IFIs that have the power, via penalties at their disposal, to enforce repayment. In this section we assume official creditors extend no loans to the country, so that investment and consumption in the first period are constrained by current resources. In this setting, first period public investment is $e_1 = x$ and, with no international borrowing and no second period loan repayment, $e_2 = 0$.

Cooperative Solution

If groups coordinate, perhaps through the facilitation of a strong central government, then each group understands that there is no way to obtain transfers at the expense of the other groups. In other words, it is understood that transfers per group will equal taxes per group. This recognition removes all incentives to divert resources to competitive rent-seeking, making it optimal to set $n = 1$.

The central government then chooses x to maximize the representative household's utility. The solution is $x^* = \theta\beta wh_1/(1+\theta\beta)$. This assumes that the country is sufficiently poor that x^* is less than the productively efficient choice of x, which is found by equating the marginal product of x to $1+r$, or $[\theta w/(1+r)]^{1/(1-\theta)}$.

Noncooperative Solution

If the groups do not coordinate their decisions, then each chooses its level of rent-seeking effort taking the other groups' behavior as given. Households act under the belief that some of the tax burden of raising their transfers can be passed off to other groups. The central government and the different interest groups play a noncooperative Nash game, where all actions are taken simultaneously.[10]

We treat each group symmetrically, so the first-order conditions for the common choice of n in each period i is

$$\left(1 - \tau_i(1 - \frac{1}{m})\right)w = \phi_0\phi(1 - n_i)^{\phi-1}\left(1 - \frac{1}{m}\right). \tag{4}$$

[10]See Appendix I for derivations of group behaviors in the noncooperative Nash equilibrium.

The left-hand side is the marginal benefit of allocating human capital to production, the after-tax rental rate on human capital. This expression is adjusted for the fact that when a group increases its productive work, the tax base increases and tax rates can be lowered. However, the lower taxes are spread across the entire economy so that the individual group only enjoys $1/m$ of the tax saving. The right-hand side is the opportunity cost of allocating time to productive activity, the forgone net transfers that would result from further rent-seeking. Marginal increments in rent-seeking yield positive net transfers because each group views the tax price of a dollar of transfers as $1/m$. Again, this is because the tax increase needed to raise transfers to just *one* group will be spread over m groups via a higher income tax rate.

In general equilibrium, one must account for the effect of all household decisions, and public investments, on the economy's income tax rate. Since all households are identical, the government budget constraint can be written as

$$\tau_i w h_i = \frac{e_i + \phi_0(1-n_i)^{\phi} h_i}{n_i}. \tag{5}$$

Note that, in the end, taxes must cover transfers for each group so that no group actually gains from rent-seeking.

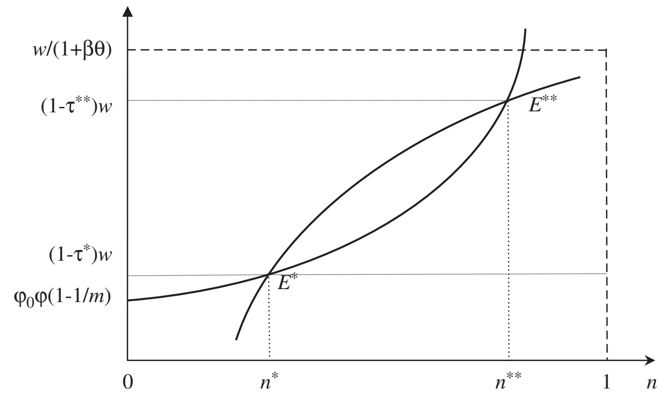
In period 1, the central government's choice of x is $e_1 = \bar{x} = \beta\theta w h_1 n_1/(1+\theta\beta)$. Note that rent-seeking further reduces investment by lowering first period income (since $n_1 < 1$). In equilibrium, the economy's level of investment is lowered by borrowing constraints and by rent-seeking. Substituting the expression for public investment into equation (5) and then substituting equation (5) into equation (4), gives

$$\begin{aligned} & w\left[1 - \frac{\beta}{1+\beta\theta}\left(1-\frac{1}{m}\right)\right] - \frac{\phi_0(1-n_1)^{\phi}}{n_1}\left(1-\frac{1}{m}\right) \\ & = \phi_0\phi(1-n_1)^{\phi-1}\left(1-\frac{1}{m}\right). \end{aligned} \tag{6a}$$

The left-hand side now accounts for the fact that the tax rate is decreasing in the common value of n chosen by all groups (both because transfers fall and the tax base rises with n), causing net-of-tax wages to rise with n. The fact that net wages increase with n at a decreasing rate creates two possible equilibrium outcomes (see Figure 1). Because rent-seeking unambiguously lowers income and welfare, the **equilibrium with higher *n Pareto dominates the *equilibrium with lower* n.

In period 2, we have $e_2 = 0$, but the configuration from Figure 1 remains the same, so we once again have two equilibria. Appendix II gives a numerical example of the equilibria under autarky. While there are other possibilities, we focus on the situation where if a country starts in the *equilibrium in period 1, it remains there in period 2, and likewise for an economy starting in the **equilibrium.

Figure 1. General Equilibrium of Rent-Seeking Game



An important comparative static question concerns the effect on productive work of an increase in the number of interest groups. To address this question, multiply both sides of equation (6a) by $\mu \equiv (1-1/m)^{-1}$, leading to

$$w\left[\mu - \frac{\beta}{1+\beta\theta}\right] - \frac{\phi_0(1-n_1)^{\phi}}{n_1} = \phi_0\phi(1-n_1)^{\phi-1} \tag{6b}$$

Now the graphic configuration of equation (6b) takes the same form as equation (6a) in Figure 1. An increase in the number of interest groups (m) will lower μ, decrease the left-hand side and cause a downward shift in the analogue to the concave net wage function in Figure 1. Thus, an increase in m will have different effects on productive work effort and rent-seeking across the two equilibria. In the *equilibrium productive work effort will increase with m and in the **equilibrium productive work effort will decrease.

The result in the **equilibrium is more intuitive. The cost to any one interest group of demanding additional transfers declines as the number of interest groups increases (because their tax share is smaller). The lower relative cost of rent-seeking results in less productive work, more rent-seeking, more transfers, and higher tax rates. However, the *economy is in an initial position with low levels of n, and high tax rates. In fact, the tax rates are so high that the economy is on the wrong side of the Laffer curve—an increase in tax rates will reduce tax revenue. From this position, a drop in productive work effort, greater rent-seeking, and greater spending cannot be financed by higher tax rates. Instead, equilibrium can only be restored if tax rates are cut, and cut enough to encourage more work effort.

In sum, an increase in the number of groups in our voracity model of special interest politics leads to a reduction in productive work and an intensification of rent-seeking in the equilibrium where the country is on the right side of its Laffer curve. This is in contrast to the Grossman-Helpman model of special interest politics. In that model, as the number of interest groups increases and the entire population is represented in some group, competition among interest groups leads to positive amounts of political contributions but has no effect on the government's choice of policies. In the limit, polices revert to the first-best, distortion-free equilibrium choices of either tariffs or subsidies (see Grossman and Helpman, 1994; Dixit, Grossman, and Helpman, 1997).

III. Foreign Assistance

We now consider the impact of loans extended to the developing country by its international donors. We assume that the government uses each dollar borrowed for productive public spending, which is here identified with government investments in human capital. Donor enforcement of these conditions is perfect and costless. The investment condition keeps the proceeds of international assistance out of the ''common pool'' of resources that interest groups compete over for transfers. However, as we shall see below, this is not enough to guarantee that international assistance generates good outcomes.

Project Conditionality

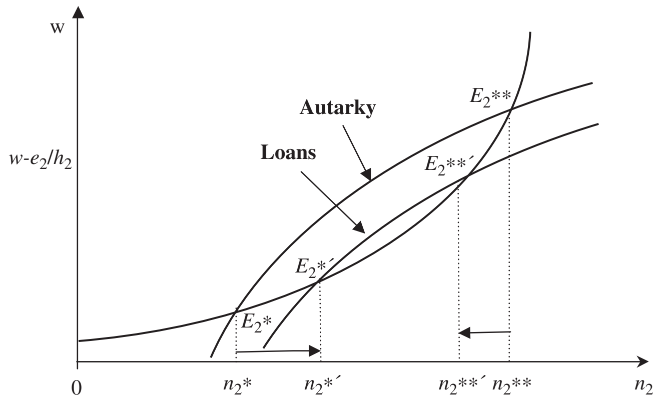
Suppose the donors extend loans equal to ml under the condition that the proceeds are used to augment existing investment, $ml = m(x - x^{*})$. This gives us $e_1 = x - l = x^{*}$ and $e_2 = (x - x^{*})(1 + r)$. Since e_1 and taxes are unchanged in period 1, the solution for n_1 remains unchanged. However, the increase in taxes needed to pay back the debt in period 2 results in a shift of the net wage rate curve downward (see Figure 2), creating mixed results.

Countries starting in the *equilibrium will experience an increase in productive work. The obligation to pay off debt creates a need for additional tax revenue. Recall that the *economy is on the wrong side of the Laffer curve. From this position, tax revenue can be raised only by *lowering* tax *rates* and thereby encouraging additional work.

The opposite is true if the economy starts in the **equilibrium. Here, the need for additional revenue requires that tax rates be increased. This discourages work and gives rise to additional rent-seeking.[11] Even if the period 2 work level is held constant, public investment may not generate enough additional earnings to both pay the debt and increase consumption (since the initial level of rent-seeking may lower the return to investment below $1 + r$). Moreover, since productive work declines further when taxes rise to repay debt, there is a greater likelihood that income will not increase

[11]Higher taxes would also hit the wages paid to those in unproductive government employment. However, interest groups would work to protect their after-tax wages by lobbying for higher before-tax wages, so that their net transfer from the government remains the same. Thus, taxes will primarily lower the reward to productive work.

Figure 2. A Transfer Paradox: Repayment Equilibrium with International Loans



by enough to cover the debt obligations and the recipient country may end up being worse off. Appendix II provides a numerical example of this possibility.

This result amounts to a transfer paradox of sorts: an apparently favorable event such as providing investment loans to a credit-constrained country can make it worse off. Excessive rent-seeking lowers the return to investment other things constant. While wh'_2 might be relatively high, a lack of productive work can make $wh'_2 n_2$ quite low. Thus, the central government of the recipient country may actually be reluctant to consider international borrowing. The reluctance is stronger if the government realizes that the higher taxes needed to repay the loans will generate further attempts to avoid taxes by allocating even more labor away from productive work and toward rent-seeking. Putting pressure on the country to accept loans may lead to adverse economic outcomes in this situation.

Program Conditionality: Extending Conditions to Fiscal Policy

We now seek conditions that could reasonably be imposed by donors that would lead to an increase in second-period work, and thereby guarantee welfare improvements, from *either* initial equilibrium position. One set of conditions that works is the following:

(a) every dollar loaned must be invested (as above);
(b) second period tax rates must remain fixed at pre-loan levels ($\bar{\tau}_2$);
(c) loan repayment must be financed by cuts in transfer spending.

Let α denote the uniform cut in transfers needed to balance the budget (a policy parameter). Conditions (a)–(c) have the following effect on the recipient government's budget constraint:

$$\bar{\tau}_2 w h_2 \sum_{j=1}^{m} n_{2j} = m e_2 + (1-\alpha) \sum_{j=1}^{m} T_{2j}. \qquad (7)$$

The first-order condition for n_2 becomes

$$\left(1 - \bar{\tau}_2(1 - \frac{1}{m})\right)w = (1 - \alpha)\phi_0\phi(1 - n_i)^{\phi-1}\left(1 - \frac{1}{m}\right). \tag{8}$$

Note that the marginal cost of productive work is lower than under autarky, given in equation (4), by the factor $1-\alpha$ (compare equations (8) and (4)). The loan conditions maintain the marginal benefit of productive work independent of n_2 (as the tax rate is fixed at $\bar{\tau}_2$) but reduce the marginal cost of productive work (by implicitly ''taxing'' transfers to pay back the debt). This is sufficient to increase n_2 from either equilibrium position (see Figure 3). Appendix II gives a numerical example of this outcome.

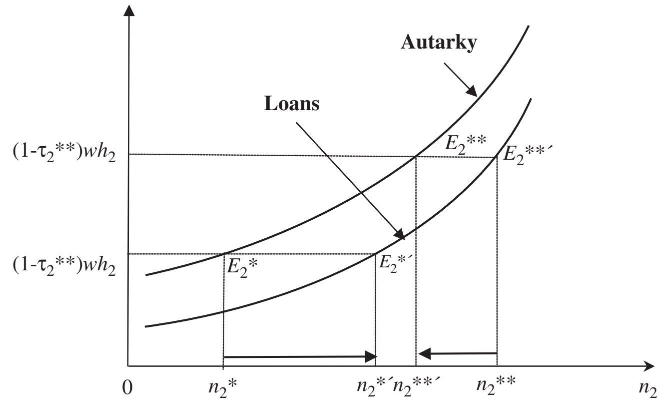
The Key Role of the Finance Minister

Program conditionality can be implemented only by the central government of the recipient country. The combination of foreign financing and conditionality enable the central government to enact spending cuts that are in the national interest but which the government would not be able to enforce in the absence of foreign financing. In effect, the international community's provision of conditional finance helps to ''steel'' the back of the finance minister and his technical staff to implement needed expenditure reforms. Krueger (1990, pp. 18–20) argues that, among government officials, finance ministers in developing countries are the most likely to be focused on the national interest. As she writes, ''Spending ministers will tend to become advocates of programs and policies falling within their domain. By contrast, finance ministries tend to be public interest agencies to a greater degree…Typically, each spending ministry will want to increase spending, believing it in the social interest that those activities within its particular domain are the most important. The finance ministry, by contrast, will be more concerned about raising revenue, and is therefore less likely to represent special interest.''

In practice, rather than imposing across-the-board cuts in unproductive government consumption and transfers, the finance minister is likely ''negotiate'' cuts with the donors in the areas where spending is believed to be particularly ineffective. Behind the scenes, it is very much the case that the finance minister ''owns'' the policy agreements with the donor community.[12] A clear example of how conditional finance strengthens the authorities' ownership of reforms that promote the general interest is documented by Mallaby (2004). He describes the difficult job faced by Uganda's Finance Minister, Emmanuel Tumusiime-Mutebile, in controlling the country's

[12]In stressing the importance of cutting government consumption to repay loans, we do not deny that in many developing countries productive public officials are paid too little, and unproductive ones are paid too much. Although difficult to implement, the best policy would be to cut government consumption overall and reallocate spending to productive government employees.

Figure 3. Equilibrium: International Loans with Fiscal Policy Conditions



fiscal spending. Tumusiime-Mutebile saw the political advantage of having the World Bank and the IMF place conditions designed to limit spending.

> In the same way, controlling government spending has involved constant battles with the cabinet, so Tumusiime was delighted to have the Bank and the IMF mandate spending discipline. In sum, the lesson from Uganda was quite subtle. It might be true that conditionality cannot force reform upon on unwilling country. But that did not make conditionality useless. Where you have strong reformers, conditions can make them even stronger. (Mallaby, 2004, p. 229)

IV. Subnational Governments

The interest groups may be interpreted as subnational (state and local) governments that compete for national funds to finance transfers and unproductive government employment. Under this interpretation, one may ask if it would be better to make conditional loans directly to subnational governments rather than working through the central government. Since the late 1990s, the World Bank has been making conditional loans to state governments in Brazil, Argentina, India, Mexico and other countries. These loans have achieved some success but are generally viewed as riskier than loans to central governments. While it may be difficult to enforce repayment of loans extended directly to the states, it is worthwhile to consider the possibility in order to think through other advantages and disadvantages of the strategy.

One advantage of making loans to state governments is that higher state taxes may not be distortionary if they are clearly earmarked for the repayment of state loans. Taxes viewed as "loan repayment" avoid the negative distortionary effect on productive work associated with higher national taxes. This argument is similar to the claim that social security taxes would not be distortionary under a fully funded system, since workers would view them as contributions to their individual accounts.

Assuming that state taxes are not distortionary implies that loans made directly to the states would not tilt the allocation of labor toward rent-seeking, an improvement over the policy of making loans with project conditions through the central government.[13]

However, by not extending loans through the central government, one *forgoes the opportunity to reduce rent-seeking* by imposing broader conditions on unproductive national fiscal policy. As we demonstrated in Section III, if it is possible to restrict the central government to repay loans by cutting transfer spending or government consumption, then rent-seeking can be reduced.[14] Thus, loans to the central government are preferred if broader conditions can be successfully imposed, otherwise loans to state and local governments would lead to better outcomes.

V. Grants

Our analysis has some implications for the debate whether development grants produce better outcomes than development loans. Pure grants are clearly a more expensive way to assist developing countries in the short run (since loans are, at most, only partially subsidized). However, if grants induce better economic performance in developing countries than do loans, then the long-run cost of supporting these countries may be reduced.

One advantage of grants is that they avoid the accumulation of large debt burdens from being passed from one political regime to another (a particularly difficult case is odious debt contracted by unrepresentative governments). However, our analysis suggests that loan repayment is not necessarily a burden. In a country with excessive rent-seeking activity, loan repayment can be a blessing. Whether repayment is a burden or a blessing depends on how the funds for repayments are obtained.

If the funds needed for repayment are obtained by raising tax rates, then the burden of the debt will exceed the direct loss of funds. Higher tax rates will punish productive activity in favor of rent-seeking and national output will fall. The recipient country's losses exceed the loan repayment in this case.

[13]Note that central government taxes would not change under the loans to subnational governments. There is not even an indirect effect on tax rates from the increased human capital in the second period because higher human capital would increase the tax base and the transfer expenditures proportionally.

[14]Such a policy has the potential to strengthen the hand of weak central governments in their attempt to resist spending initiatives by local governments. For example, in Argentina the initiative for public spending comes from the provinces, while responsibility for raising revenue is often passed off to the central government (Mussa, 2002, p. 14). A decentralized fiscal policy with soft subnational budget constraints is considered to be a general problem in Latin America (Stein, 1999).

Grants could potentially avoid the excess burden of lost output since no repayment is required.[15]

However, this ignores the strategic use of loan repayment to strengthen the hand of finance ministers over the "spending" ministers and subnational authorities. If broad conditions require loan repayment to be financed by general cuts in unproductive transfer programs or government consumption, then repayment becomes a blessing. "Forcing" the finance minister to cut noninvestment spending will "tax" rent-seeking and increase national output. Thus, when rent-seeking is a major issue, whether to favor grants of loans will depend on the particulars of the country. If the finance minister and the central government are committed to increasing national output and are looking for leverage to cut transfer spending, then loans with the right broad conditions can be an advantage.

We can conduct a formal comparison of grants and loans in our model. To make the comparison meaningful, we examine the effects of grants and loans of *equal* present value to donors. Suppose that the interest rate on the loan to the developing country is subsidized. The forgone interest income due to the subsidy implies that there is a "grant component" to the loan. Instead of a subsidized loan, consider giving the country a pure grant in the first period that is equal to the present value of the loan subsidy. Furthermore assume that the grant is conditional—that is, it is used to increase investment spending dollar-for-dollar, as we assumed in the analysis of a loan. Would the grant make the recipient country better off than the loan? As mentioned above, the answer depends on whether or not donors are able to impose broad conditions successfully on the recipient.

If the broad conditions on fiscal policy *can* be imposed, then subsidized loans are preferred to grants for two reasons. First, under the broad conditions, loan repayment is used as a way of taxing and reducing rent-seeking in the second period. Conditional grants, by contrast, would not require any fiscal adjustment in the second period and therefore rent-seeking will be unaffected (recall that any increase in second-period human capital has no effect on the decision to seek rents or work productively). Second, for a given present value, loans always provide more funds for first-period investment because the grant is equal to the present value of the *subsidized portion* of the loan. The increase in

[15]We say potentially, because we are assuming that grants are used exclusively to finance public investments, as are loans, and the other features of public expenditures and taxes remain fixed. However, in practice, grants have been associated with a decline in tax revenue collection (Clements and others, 2004). Thus, at least some of the grants are used to lower taxes. Because grants are not permanent sources of revenue, it is unlikely that taxes are lowered by a legislated cut in tax rates, which would reduce rent-seeking in favor of productive spending. Instead, the loss in tax revenue may be due to special tax exemptions or lack of tax enforcement that favors particular interest groups and which are transfers that reward rent-seeking. These considerations caused Clements and others to recommend that broad conditions be placed on grants to prevent losses of tax revenues. This is equivalent to requiring that grants funds be used dollar-for-dollar to increase investment (as we have assumed here).

second-period investment income is unambiguously greater than the additional interest expense if the recipient country was credit constrained in autarky (see Appendix III).

However, if the broad conditions cannot be imposed on the country, grants may be superior to a subsidized loan. A subsidized interest rate reduces the likelihood that the recipient country will be made worse off by the loan but does not eliminate it. Taxes will have to be raised in the future to pay off the principal and the subsidized interest charges. If the decline in productive work is sufficiently great, then the country could be made worse off. However, the grant *must* make the country better off, even when the loan will not. There is greater human capital and greater income in the second period due to the grant. There is no need to raise taxes and therefore the level of productive/rent-seeking is unchanged. Thus a grant can outperform a subsidized loan of equal present value (when the rent-seeking response to higher taxes is sufficiently high).

VI. Optimal Loans

We have shown that loans to the central government can improve welfare in recipient countries racked by voracious politics provided that conditions can be imposed on these countries' fiscal policies. This raises the question of what the *optimal* loan looks like under these assumptions.

Suppose the donor community chooses the loan quantity to maximize the recipient country's welfare subject to the broad conditions (a)–(c) stated in Section III. The optimal loan quantity solves

$$\max_{l} \{\ln[(1-\tau_1)wh_1n_1 - x^*] + \beta\ln[(1-\bar{\tau}_2)wh_2n_2 + (1-\alpha)T_2]\},$$

subject to the government budget constraint

$$\bar{\tau}_2 wh_2n_2 = l(1+r) + (1-\alpha)T_2.$$

This problem is equivalent to solving $\max_{l} \{wh_2n_2 - l(1+r)\}$, that is, it is equivalent to maximizing second period consumption of the representative household. The first-order condition for the optimal loan is

$$wh_2'n_2 + wh_2\frac{\partial n_2}{\partial\alpha}\frac{\partial\alpha}{\partial l} = 1 + r. \tag{9}$$

The left-hand side can be interpreted as the marginal benefit of the loan. The first term is the marginal return to human capital investments for a given level of productive work. The presence of rent-seeking lowers this return, because $n_2 < 1$. The second term is the indirect effect of the loan on fiscal policy and productive work. Appendix III shows that larger loans raise the implicit tax on transfers ($\partial\alpha/\partial l > 0$) and that this raises productive work ($\partial n_2/\partial\alpha > 0$). Thus, voracious politics lowers the marginal returns to investment by lowering productive uses of human capital but also raises the marginal benefit of loans if repayment comes through cuts in transfers to rent seekers.

The right-hand side is the marginal cost of a loan, the repayment of principal and interest. If the marginal benefit is decreasing in the loan quantity the first-order conditions are also sufficient conditions. Our numerical analysis produced no examples where the marginal benefit was increasing in the loan quantity.

To create a reference point, define a *first-best* loan to be the loan that would remove the credit market imperfection and generate the efficient amount of human capital in *absence of voracious politics*. We find that optimal loans in the presence of voracious politics can be greater than the first best loan. Greater loans to countries with bad politics! This surprising result is obtained because conditional loans can be used to reduce rent-seeking (the addition of the second term in the marginal benefit expression), which more than offsets the fact that rent-seeking lowers the return to human capital investment (the appearance of $n_2 < 1$ in the return to human capital investment expression).

Whether or not the optimal loan exceeds the first-best loan quantity depends on whether rent-seeking is eliminated at a loan quantity that is less than the first-best quantity. When rent-seeking is eliminated, $n_2 \equiv 1$, the marginal benefit of loans converges to the first-best case. If productive work is responsive to the implicit tax on rent-seeking, then as the loan and the tax increases it becomes more likely that rent-seeking is eliminated before the first-best quantity is obtained. If rent-seeking is less responsive to the implicit tax, then the optimal loan will exceed the first-best because of the additional political advantage of further raising loans and the implicit tax on rent-seeking.

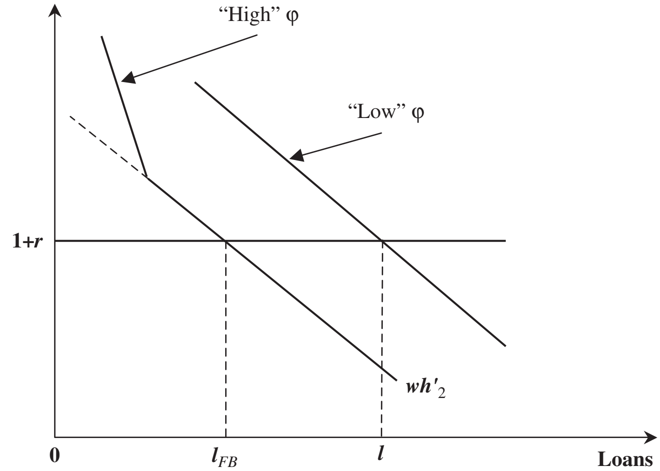
Figure 4 depicts the situation. We plot the declining marginal benefit of loans (the left-hand side of equation (8)) against the constant marginal cost (the right-hand side of equation (8)). If ϕ is high, then the diminishing returns associated with rent-seeking are modest. As the implicit tax on rent-seeking is increased, large changes in productive work are needed to raise the "after-tax" marginal benefit of rent-seeking to the (constant) after-tax marginal benefit of productive work. A low ϕ, on the other hand, results in small n_2 responses to changes in α. Numerical examples are provided in Appendix II to support this interpretation.[16]

VII. Concluding Remarks

Countries with voracious politics divert large amounts of public resources to unproductive transfers to powerful interest group. This is a source of poverty, because resources devoted to such redistribution reduce the efficiency of resource allocation and lower national welfare. Fractious politics also makes it more difficult for the international community to assist developing countries. Loans conditioned on public investment alone may not improve welfare in countries with voracious politics. The higher taxes needed

[16]Our numerical analysis only produced outcomes where the optimal loan was at least as large as the first-best loan.

Figure 4. Voracious Politics and the Size of the Optimal Loan



to pay back foreign debt may further increase rent-seeking, which could lower the returns to public investment below the cost of debt repayment.

When confronted with a highly contentious political environment in recipient countries, donors may require additional conditions to guarantee that their loans will increase the recipient's national welfare. One guiding principle for the donor community is to develop conditions that shift the burden of debt repayment away from the taxation of productive work and toward the taxation of rent-seeking. This may be accomplished by insisting that debt be repaid by cuts in programs susceptible to rent-seeking. Because the susceptible programs are likely to vary from country to country, it is critical that the recipient country's finance minister, and similar officials are involved in establishing the appropriate conditions in partnership with donors. In working with recipient country officials, donors should help weak central governments reduce the incentives to engage unproductive rent-seeking.

Consistent with our analysis, donors have sought to place conditions on their assistance to developing countries that aim at increasing government investment and reducing government consumption (which may also include corrupt public "investment" projects). Empirical evidence suggests that lowering government consumption raises economic growth (see, for example, Barro, 1997, Chapter 1; and Baldacci, Clements, and Gupta, 2003). Our analysis offers a new supporting argument for loan conditions that require cuts in government consumption, rather than increases in taxes, to finance loan repayment.

APPENDIX I

Noncooperative Behavior

Autarky

In the first stage of the noncooperative game, each interest group chooses (n_1,n_2) to maximize its representative agent's utility taking other groups and the central government's investment decision as given. Recall that everyone in the recipient country is credit-constrained, that is, they would like to borrow but cannot without the support of official donors and creditors. Thus, their choice of financial assets is constrained to be zero. Group k solves

$$\max_{(n_1,n_2)} \{\ln c_{1k} + \beta \ln c_{2k}\},$$

subject to

$$c_{ik} = (1-\tau_i)wh_i n_{ik} + T_{ik}$$

and

$$\sum_{j=1}^{m} T_{ij} + me_i = \tau_i wh_i \sum_{j=1}^{m} n_{ik},$$

for $i = 1, 2$. Note that each group takes into account how its decisions affect the spending levels and the tax base in the government budget constraint (given others' behavior). To solve the problem, note that the choice of n is a purely static decision in each period; that is, choose n to maximize income and consumption possibilities in each period. Substituting the government budget constraint into the group's budget constraint, we get

$$c_{ik} = wh_i n_{ik} - \frac{me_i + T_{ik} + \sum_{j\neq k}^{m} T_{ij}}{n_{ik} + \sum_{j\neq k}^{m} n_{ij}} n_{ik} + T_{ik}.$$

Differentiating with respect to n and setting the expression to zero gives

$$\frac{dc_{ik}}{dn_{ik}} = (1-\tau_i)wh_i + \frac{dT_{ik}}{dn_{ik}} + \left(\tau_i wh_i - \frac{dT_{ik}}{dn_{ik}}\right)\frac{n_{ik}}{\sum_{j=1}^{m} n_{ij}} = 0. \tag{A.1}$$

The first term is the effect of greater productive work on after-tax wages (holding the tax rate constant). The second term is the (negative) effect of greater productive work on transfer income. The third term is the net effect of greater productive work on the group's tax burden. Greater productive work lowers their tax burden because the tax base increases and government spending falls. However, this benefit is spread over all groups, and group k only receives a share of it.

Because the economic fundamentals of all groups are identical, their choices of n will be the same. Thus, the k-subscript can be dropped and $\sum_{j=1}^{m} n_{ij} = mn_i$. The first-order condition for n can then be written as equation (4) in the text.

In the second stage, the central government takes the rent-seeking choices of the groups as given and chooses x to maximize household utility. Again, we assume the government cannot borrow on behalf of its households without the intermediation of official international donors and creditors. The government solves $\max_x \{\ln c_1 + \beta \ln c_2\}$ subject to $c_i = (1-\tau_i)wh_i n_i + T_i$ and $\tau_i wh_i n_i = e_i + T_i$, $e_1 = x$, $e_2 = 0$. The solution to this problem gives $e_1 = x^* = \beta\theta wh_1 n_1/(1+\beta\theta)$.

Investment Project Conditionality

By design, the conditions require that tax-financed investments in the first period be kept at the same level as in autarky. Therefore, in the first period, behavior is unaltered. In the second period, we now have $e_2 = l(1+r) > 0$. This raises τ_2 but otherwise does not change the form of equation (A.1).

Extending Conditionality to Fiscal Policy

Because the broad conditions also require that tax-financed investments in the first period be kept as the same level as in autarky, behavior in the first period is, again, unaltered. However, under program conditionality, tax rates in the second period must be frozen, and the repayment of debt must be financed by spending cuts. This leads to the government budget constraint, equation (7), under which an interest group's second period problem generates the first-order condition

$$(1-\bar{\tau}_2)w + (1-\alpha)\frac{dT_{2k}}{dn_{2k}} + \frac{d\alpha}{dn_{2k}}T_{2k} = 0. \tag{A.2}$$

Differentiating equation (7) gives

$$\frac{d\alpha}{dn_{2k}} = \frac{(1-\alpha)(dT_{2k}/dn_{2k}) - \bar{\tau}_2 w h_2}{\sum_{j=1}^{m} T_{2j}} \tag{A.3}$$

Substituting into equation (A.2) and evaluating in the symmetric equilibrium gives equation (8) in the text.

APPENDIX II

Numerical Equilibria

We think of each period as lasting 30 years. The international interest rate is set to an annualized rate of 7 percent, a common value used for the return to capital in developed economies. If we assume a Cobb-Douglas production technology with a physical capital share of one-third, then our interest rate assumption yields a corresponding value for the rental rate on human capital, w, from the standard factor-price frontier.

We assume, unless otherwise specified, that $\phi = \theta = 0.5$. The implicit rate of time preference that defines β is set to an annualized value of 7 percent. The number of interest groups is set to 5 and the fraction of time spent on productive work in the first period is 0.75. These assumptions imply a value for ϕ_0 (to be consistent with an optimal choice of $n_1 = 0.75$), a first period tax rate of 0.54 (to balance the budget), and a level of x under autarky that is half the efficient level in a first-best world.

Equilibria under Autarky

To find the multiple equilibria in period 2 under autarky, we plot the Figure 1 corresponding to our numerical example and locate the approximate positions of the two equilibrium values of n_2. We then numerically solve for the two values of n_2 that approximately equilibrate the values of the convex and concave functions depicted in Figure 1 by altering the initial guesses of the iterative search. The two equilibrium solutions for n_2 in our example were 0.65 and 0.85.

Loans with Investment Project Conditionality

We next demonstrate that bad outcomes are possible from loans with simple investment conditions, that is, we provide a numerical example that corresponds to the outcome in the **equilibrium of Figure 2. We increased the loan by an amount sufficient to cover 10 percent of the financing gap between initial investment and first-best level (if the loans are much larger, the country would be unable to pay them back, that is, second period taxes would be driven to 100 percent). We computed the new **equilibrium with the tax rate endogenously increasing to balance the government budget in the second period. The new level of work effort falls from 0.85 to 0.82. The *social* or *pre-tax* rate of return on the human capital investment is only 1.2 percent. Thus, the rise in second period earnings is not nearly sufficient to cover the pincipal and interest (at 7 percent) repayment, and second period consumption in the recipient country must fall.

Loans with Extended Fiscal Policy Conditionality

Here we assume the same loan quantity as above, but now we fix second period taxes and make up the lost revenue by cutting transfer payments (raising α). The equilibrium α is 0.06 and n_2 rises to 0.87. The rate of return on the investment financed by the loan is now 8.8 percent. Thus, the recipient country's earnings rise more than enough to repay the loan and it is made better off.

APPENDIX III

Grants versus Loans

In this appendix, we establish the conditions under which a subsidized loan will increase the welfare of the recipient country more than a grant with the same present value as the loan subsidy.

For simplicity, begin with the special case of full interest subsidy (that is, consider an interest-free loan). It is also convenient to write the human capital production function more generally than in the text, as $h(x)$. Finally, let a superscript "0" denote the initial, or pre-policy, value of a variable.

Since we are considering conditional loans (that is, first period funds are used for increased investment dollar-for-dollar), no other first period variables are affected by either policy. Any welfare change for the representative household is due to a change in its second period after-tax income.

Second period income under a loan is $n_2h(1+x^0)-l$ and second period income under a grant is $n_2^0h(rl/(1+r)+x^0)$. Note that, as argued in the text, work effort may differ from its pre-policy value under a loan but not under a grant. Also note that the grant is $rl/(1+r)$, the present value of the loan subsidy. Now consider the effect of an increase in the loan (grant) on second period income,

$$Loan: \quad \frac{dn_2}{dl}h(x^0) + n_2^0h'(x^0) - 1.$$

$$Grant: \quad n_2^0h'(x^0)\frac{r}{1+r}.$$

We show first that with the initial level of work effort held constant, loans will increase second period income more than grants,

$$n_2^0 h'(x^0) - 1 - n_2^0 h'(x^0)\frac{r}{1+r} = \frac{n_2^0 h'(x^0)}{1+r} - 1 > 0,$$

provided that the developing country is actually credit constrained initially (that is, the rate of return to investment exceeds the market interest rate at the initial levels of investment and work effort).

To see that this result extends to less than full-interest subsidies, imagine reducing the interest subsidy by a dollar. The country loses $1/(1+r)$ in present value under the loan policy. Under an equal present value grant, the first-period grant and first-period investment fall by $1/(1+r)$ and thus second period income falls by $n_2^0 h'/(1+r) > 1 > 1/(1+r)$, or by more than under the loan. Thus, interest subsidies that are less than complete only strengthen the result that loans outperform grants whenever (1) work effort is constant; and (2) the developing country is credit-constrained in the pre-policy position. In summary, when the recipient is credit-constrained it would rather pay more in interest costs than lose the present value of the interest cost in investment (since the investment return on the present value of that cost will raise second period income more than the cost will lower it). Grants save interest expense but reduce investment income by even more.

When the broad conditions are also imposed on the country, the case for loans is strengthened since $(dn_2/dl)\ h < 0$. However, if the broad conditions cannot be impose then $(dn_2/dl)\ h < 0$, since taxes will have to increase to pay back the loan. In this case, if the disincentive effect of tax on work is strong enough, then second period income may rise more under grants. In particular, when loans lead to a drop in second period income, grants can be nevertheless be guaranteed to increase second period income.

APPENDIX IV

Optimal Loans

In this appendix, we examine the properties of the optimal loan made under the assumption that the extended conditions from Section III are met.

Comparative Statics under Extended Fiscal Policy Conditionallity

With tax rates fixed, the amount of time devoted to productive work can be explicitly solved for as

$$n_2 = 1 - \left[\frac{(1-\alpha)\phi_0\phi\left(1-\frac{1}{m}\right)}{(1-\tau_2)w}\right],$$

which implies $\frac{\partial n_2}{\partial \alpha} = (1-n_2)((1-\phi)(1-\alpha)) \geqslant 0$, with a strictly positive value if there is some rent-seeking.

Implicitly differentiating the government budget constraint under extended conditions with respect to l gives

$$\begin{aligned}\tau_2 wh_2' n_2 + \tau_2 wh_2 \frac{\partial n_2}{\partial \alpha}\frac{\partial \alpha}{\partial l} = {} & 1 + r - \frac{\partial \alpha}{\partial l} T_2 \\ & - \frac{1-\alpha}{1-\phi}\phi_0 \phi (1-n_2)^{\phi-1} h_2 \frac{\partial n_2}{\partial \alpha}\frac{\partial \alpha}{\partial l} \\ & + (1-\alpha)\phi_0 (1-n_2)^{\phi} h_2'.\end{aligned}$$

Using the first-order condition for the optimal choice of n_2, and after some tedious but straightforward algebra, one can get

$$1 + r > wh_2 \frac{\partial n_2}{\partial \alpha}\frac{\partial \alpha}{\partial l} = (1+r)\left[\frac{\left(1 - \frac{\theta}{1+\frac{x^*}{l}}\right)}{\frac{1-\frac{\tau_2}{m}}{1-\frac{1}{m}} + \frac{(1-\phi)(1-\tau_2)}{\phi(1-\frac{1}{m})}}\right] > 0,$$

since $\theta < 1$ and $\tau_2 \leqslant 1$.

Numerical Examples of Optimal Loans

Note that the results in Appendix II confirm that $wh_2 \frac{\partial n_2}{\partial \alpha}\frac{\partial \alpha}{\partial l}$ is decreasing in l when there is some rent-seeking and is zero otherwise. One cannot guarantee that the expression $wh'_2 n_2$ is decreasing in l because while wh'_2 is decreasing in l, n_2 is increasing in l (via α). Thus n_2 cannot increase too fast in order to guarantee a declining marginal benefit. As mentioned, we found the marginal benefit declined in all cases we examined.

Second, note that α may reach 1 before l reaches l_{FB}. In this case the marginal benefit converges to wh'_2 (the marginal benefit in a first-best world) before l reaches l_{FB}, and thus the optimal loan is l_{FB}. This is the outcome for the calibration used to produce the numerical examples in Appendix II.

However, if we make rent-seeking less responsive to the implicit tax on rent-seeking then we can generate the second type of outcome depicted in Figure 4 of the text. For $\phi = 0.5$ rent-seeking is sufficiently unresponsive to changes in α that the optimal loan exceeds l_{FB}. When ϕ is increased to 0.75, then n_2 is driven to 1 at loan levels that are less than l_{FB}. In this case the optimal loan is equal to l_{FB}.

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What Explains the Rapid Growth in Transition Economies?

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This paper analytically explores and empirically tests a number of hypotheses to explain the rapid growth in transition economies. Using the latest panel data, the paper finds that growth in transition economies has been higher because of the recovery of lost output, progress in market reforms, and favorable external conditions. These results are consistent with estimates from the global sample that includes 123 countries, and are robust to instrumental variable estimations and other robustness tests. A general implication of the findings is that some of the factors behind the rapid growth are unlikely to continue for a very long time and that the challenge would be to further improve the investment climate, which will require broadening the scope of macroeconomic reform into a second generation of reforms encompassing structural and institutional areas. [JEL O57, O43, F43]

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Transition economies are in a resurgent phase. From a 17-year perspective—that is including the sharp fall in the early 1990s—the record is not better than the average for developing countries. In the past decade, however, growth in most of the transition economies compares very favorably with the fastest growing economies in East Asia. In particular, the average

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unweighted average annual growth of the Commonwealth of Independent States (CIS) was about 7 percent in 1996–2006. This strong performance is in welcome contrast to the first half of the 1990s, when cumulative output contracted from 30 to 60 percent in these economies. Therefore, determining whether the strong rebound from the posttransition setbacks is more the result of conditions, including progress in market reforms that will support continuing growth requires an examination of the underlying influences.

Transition countries have been typically excluded from cross-country studies of the long term because of the short historical span, and because earlier data collection methods were unreliable. However, significant improvements have been made in data quality in recent years. In the economic literature there are mainly two approaches with regard to sources of growth, namely the cross-section growth accounting approach and the panel regression approach. This paper uses both approaches to analyze the sources of the recent rapid growth in transition economies, with particular emphasis on the CIS. The research questions include the following:

- Is either investment or total factor productivity (TFP) growth responsible for the major shifts in economic growth?[1]
- Is the recent strong growth explained by a bounce back from the initial posttransition setbacks (recovery of lost output)?
- Do market reforms explain the variance in relative output performance?
- To what extent is the recovery of growth driven by favorable external conditions?

The growth-accounting exercise in this paper suggests that the recent strong growth has been driven largely by growth in TFP. On average, capital accumulation made a modest contribution and employment rates continued to drop through 2004 in most transition countries. Assuming that TFP growth slows down, other sources of growth will be essential to sustain a rapid catch-up. In this connection, the recent trend of faster capital accumulation and the recent improvement in employment in some of the transition countries are expected to play a more important role in future growth.

The results of the panel regressions suggest the following: (1) transition countries that experienced larger declines in output during the early 1990s tended to grow at faster rates; (2) improvements in macroeconomic policies and market reforms explain about half of the total growth; and (3) the growth acceleration payoff to reforms in 2001–06 was enhanced by the favorable external environment (positive terms-of-trade shock, and global technological innovation). These external factors have accounted for about one-fourth of the average annual growth in transition economies.

[1]TFP is a measure of elements such as managerial capabilities and organizational competence, research and development, intersectoral transfers of resources, increasing returns to scale, embodied technical progress, and diffusion of technology.

I. Growth Accounting Approach

This section uses the growth accounting approach to determine the sources of growth in transition economies. Given the relatively short time span since the breakdown of the former Soviet Union, there are only few papers that have analyzed the determinants of growth in transition economies. Broeck and Koen (2000) analyzed the determinants of the sharp fall in output in transition economies in the 1990s. Loukoianova and Uigovskaya (2004) extended the work of Broek and Koen through data covering the period 1991–2000 for the low-income CIS. Schadler and others (2006) focused on the Baltic and Central European countries (Box 1).

Box 1. Review of Literature on Growth, Reforms, and Institutions in Transition Economies

Although there is agreement that stabilization policies are important, no consensus has yet been reached on the role of reforms in the recent strong recovery of transition economies. Below are the main papers and their findings relevant to the current study.

- Abed and Davoodi (2000) find that progress on structural reforms is both statistically more significant and economically more important than corruption in explaining differences in economic performance in transition economies.
- Broeck and Koen (2000) and Loukoianova and Uigovskaya (2004), using the growth accounting approach, found that most of the decline in 1991–97 and the subsequent recovery in output are explained by the movements in total factor productivity growth. In the absence of factor prices they assumed shares of 0.3 for capital and 0.7 for labor.
- Campos and Coricelli (2002) assess the implications of Broeck and Koen's growth accounting estimates; emphasize the role of reforms and institutions in dictating the path of transition process; and note that isolating reallocation from accumulation and technological progress remains a major challenge in transition economies.
- Fidrmuc (2003) cast doubts on the benefits of reform and Lawson and Wang (2004) failed to find a strong and positive effect of reforms on growth.
- Falcetti, Lysenko, and Sanfey (2005) found a positive and strong link between progress in market-oriented reforms and economic growth.
- Campos and Horváth (2006) constructed objective measures for three main reform areas (internal liberalization, external liberalization, and privatization) in transition economies for the period from 1989 to 2000.
- Beck and Laeven (2006), using natural resource reliance and the years under socialism to extract the exogenous component of institution building, showed the importance of institutions in explaining the variation in economic development and growth across transition economies during the first decade of transition.
- Schadler and others (2006) examined the progress toward income convergence achieved by the five Central European and the three Baltics countries and the policy challenges that these countries will face in facilitating the catch-up process. In the panel regression approach, the main variables used to explain growth were population growth, partner country growth, the relative price of investment goods, schooling, openness, government taxation, and institutional quality.
- Babetskii and Campos (2007) reviewed 43 econometric studies and found that the existing subjective measures of reform, controlling for institutions, and initial conditions were the main factors in decreasing the probability of reporting a significant and positive effect of reform on growth.

Methodology and Data Issues

The data set suffers from various serious weaknesses due to underreporting by private enterprises to avoid taxes and regulations, particularly in the early years of transition. The decline in output during the first half of the 1990s could be overstated because the statistical system was designed to collect information only on publicly owned enterprises. Beyond the mid-1990s, the information on the emerging private sector gradually became available and was incorporated into the statistical system.

One major concern about the measurement of the capital stock is that a significant portion of the communist capital stock may have been permanently scrapped. If so, this would cause the contribution of capital accumulation to be underestimated during the subsequent recovery. In order to address this concern, a one-time adjustment for the permanent scrapping of a significant portion of the capital stock during the communist era is applied; that is, the capital stock is reduced by the same rate as output between 1990 and 1995 so that the capital-output ratio is not allowed to rise during the course of the sharp contraction in output. Also, the quality of capital is treated the same over time and across countries.

Obviously, workers in different countries have different levels of skills. Typically, education and number of hours worked are emphasized as key components of effective labor. Such data, however, are not available for most transition countries. Few growth-accounting studies on nontransition economies made adjustments to labor quality by including education, age, or gender (Boseworth and Collins, 2003). Such information is available only for selected years and is limited for industrial countries and some emerging and developing countries. More importantly, the education level in transition economies, as measured by secondary school attainment or average years of study, is relatively high as compared with other developing and emerging economies, and there is little variation among them. Thus, the correlation between the education level and growth is expected to be weak in this case. In the absence of adequate indicators that reflect changes in the quality of labor over time and across countries, the growth in TFP will therefore be overestimated.

Another major challenge in using the growth accounting approach is the proper estimate of the shares of capital and labor in output. There are several approaches in the growth literature to estimating the shares of capital and labor in output. The first approach assumes that factor markets are perfectly competitive so that earnings of the factors are proportional to their productivities. However, this approach cannot be used for some transition countries due to lack of detailed national accounts statistics. The second approach uses a priori measure of capital share in the range of 0.3 to 0.4. Aiyer and Dalgaard (2005) establish that the standard Cobb-Douglas methodology of assuming a constant capital share of one-third for all countries is a very good approximation to a more general formulation under which countries have different aggregate production functions that do not require a constant elasticity of substitution among factors. The third approach estimates the coefficients of the production function by regressing the growth rate of output on the growth rates in capital and labor. The intercept then measures the growth in TFP, and the coefficients on the factor growth rates measure the shares of capital and labor, respectively. The main advantage of this process is that it dispenses with the assumption that factor social marginal products coincide with the observable factor process. The disadvantage of the regression approach, however, is that the growth of capital and labor cannot usually be regarded as exogenous with respect to variations in TFP (Barro, 1999). The fourth approach, used by Hsieh (2002), is the dual exercise. The advantage of using the dual is that factor prices, primarily wages and interest rates, are observed as an equilibrium outcome in a marketplace. However, a number of assumptions and estimates have to be made in order to construct the data on quantities of output and capital needed for a primal growth accounting exercise (Hsieh, 2002, p. 519). Shigeru, Khan, and Murao (2003) proposed a fifth approach that does not need the assumption of perfectly competitive factor markets nor assumes any particular functional form of the aggregate production function. Their approach is based on nonparametric kernel derivative estimation techniques developed in the statistics and econometrics literature (Pagan and Ullah, 1999). This approach estimates much lower elasticity of output with respect to capital (around 0.20) for several East Asian countries, thus emphasizing even more the role of the residual (growth in TFP) in explaining growth.

Results for Transition Economies

Given the data limitations for transition economies, particularly the low income CIS economies, this section uses the third approach mentioned above. Table 1 shows the regional estimates of the shares of capital and labor. The estimated TFP growth by regions, which here is the residual, was the highest for the Baltics, followed by the CIS.[2] Ideally, separate production function for each country should be used. But given the short historical span (1991–2006), it is assumed that the production functions are similar for each region. The countries included in each region share some common characteristics and are likely to have similar production functions. The estimated sum of the capital and labor elasticities are not far away from unity for the three Baltics and the five Central and Eastern European countries, but slightly higher than 1 for the CIS. The endogeneity problem is partially addressed by using two-stage least squares (2SLS), although finding good

[2]The majority of existing literature shows capital elasticity of 0.3 to 0.5 in industrial countries. The share of physical capital for industrial countries is likely to be lower than for developing countries where the marginal product of capital is higher (Boseworth and Collins, 1996, p. 155).

Table 1. Estimates of Input Shares, 1996–2006

Sample	Total Factor Productivity Growth (in percent)	Elasticity Estimates		Observations
		Capital	Labor	
Baltics—3	3.0	0.49	0.52	33
Commonwealth of Independent States—12	2.3	0.63	0.51	132
Central Europe—5	1.6	0.40	0.62	51
Southeast Europe—6	0.8	0.73	0.35	60

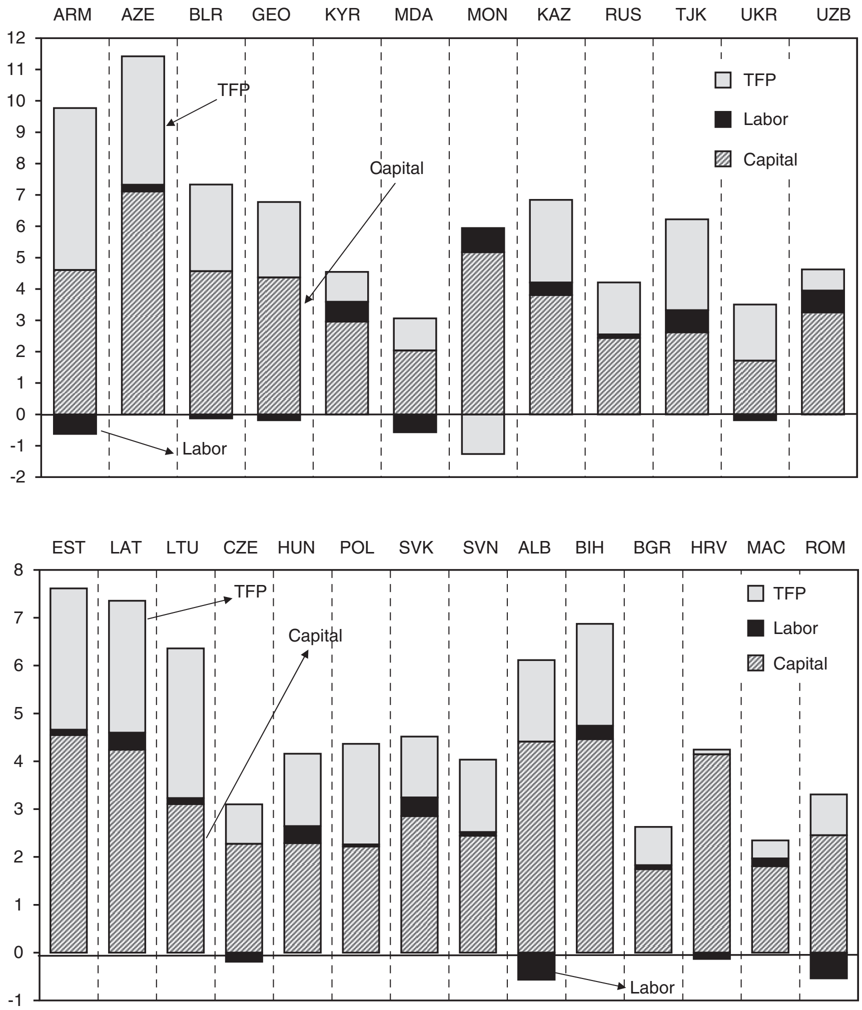
Source: Author's calculations.

instruments remains a challenge. The estimates of the shares of capital and labor are found to be robust to different estimation techniques.

Using the estimated shares of capital and labor from Table 1, the sources of growth are derived for the 26 transition countries (Figure 1). The average annual TFP growth in the CIS was higher than in Central European and six Southeast European economies, but lower than in the Baltics. A natural question is then, what were the factors that led to this high TFP growth? It is most likely that the inefficiencies inherited from central planning left much scope for managerial improvements, labor shedding, and gains from interindustry resource reallocation. Higher TFP growth could also be explained by the scale of some of the transition economies, which are relatively poor economies with very low endowment of technology. Hence for a given technological innovation, the smaller the initial endowment, the higher is the growth of TFP. When capital is scarce, its marginal productivity is considerable. Therefore, for similar investment rates, the contribution of capital deepening should be larger in economies with less capital. Increases in capacity utilization could also raise TFP growth. The sharp fall of economic growth in most transition economies before 1995 is largely attributed to steep decline in TFP. In the second phase (1996–2000) there was improvement in TFP growth in almost half of these economies.

The sensitivity of the TFP growth estimates were also examined under different assumptions of initial capital to output ratio (k) and elasticities of output with respect to capital. An increase in k from 2.0 to 2.5 raises the estimated TFP growth for the 15 countries of the former Soviet Union from 2.4 to 3.2 percentage points, and for the six Southeast European economies from 1.0 to 1.7 percentage points. A decrease in the capital share from 0.6 to 0.4 (close to the share of capital used in the literature) increases average annual TFP growth from 2.4 to 3.5 percentage points for the former Soviet Union countries (equivalent to 60 percent of output growth in 1996–2006). In general, countries with higher capital shares will tend to have lower TFP growth; a higher elasticity of output with respect to capital would result in a rise of the contribution of physical capital and a decline in the contribution of

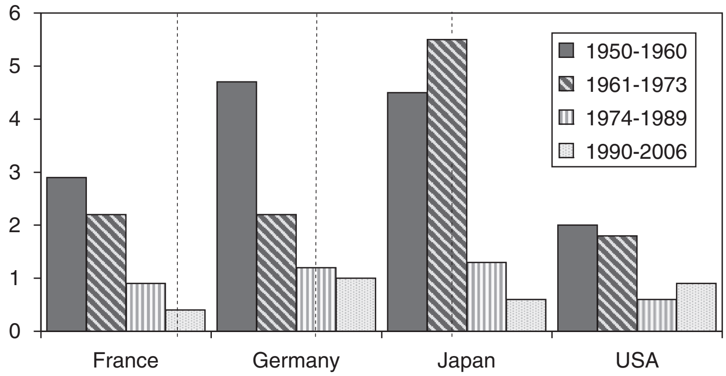
Figure 1. Contributions to Average Real GDP Growth, 1996–2006
(In percentage points of GDP)



Sources: Author's own calculations based on the following assumptions: Initial capital stock to GDP ratio of 2 and annual depreciation rate of capital stock of 5 percent. The elasticity of output with respect to capital and labor are derived from the regional production functions in Table 1.

Note: TFP = total factor productivity. First panel: ARM = Armenia; AZE = Azerbaijan; BEL = Belarus; GEO = Georgia; KYR = Kyrgyzstan; MDA = Moldova; MON = Mongolia; KAZ = Kazakhstan; RUS = Russia; TJK = Tajikistan; UKR = Ukraine; UZB = Uzbekistan. Second panel: EST = Estonia; LAT = Latvia; LTU = Lithuania; CZE = Czech Republic; HUN = Hungary; POL = Poland; SVK = Slovak Republic; SVN = Slovenia; ALB = Albania; BIH = Bosnia and Herzegovina; BGR = Bulgaria; HRV = Croatia; MAC = Macedonia; ROM = Romania.

Figure 2. Total Factor Productivity Growth, Historical Perspective
(In percentage points)



Sources: Author's estimates with the exception for the period 1950–60, which is based on Christiansen, Cummings, and Jorgenson (1980).

Note: The assumptions are that the initial capital stock to GDP ratio was 2; annual depreciation of 5 percent; and elasticities of output with respect to capital and labor of 0.3 and 0.7, respectively.

TFP growth. It should be noted that the methodology used here does not adjust factor inputs for quality changes. The implication is that the incremental effect on growth of embodied technological advancement is not attributed to capital but is rather measured as a higher level of TFP. The same measurement problem can also arise in the case of labor. Education and on-the-job training would improve the quality of labor. This would be reflected in higher TFP. This "mismeasurement" of TFP may well be significant in the case of transition economies, following the move from central planning to market economies.

Historical Perspective from Industrial and East Asian Economies

During the "Golden Age" (postwar period) in Western Europe and Japan, there were strong contributions to growth from TFP gains. Catching up, scale effects, and improvements in resource allocation made strong contributions to TFP during 1950–60 in the major industrial countries (Maddison, 1996).[3] These improvements stemmed from adjusting to trade liberalization, exploiting opportunities for mass production as larger and better integrated markets emerged, and from moving resources out of relatively low-productivity agriculture. As catch-up growth weakened, the magnitude of TFP growth fell markedly after 1973. The estimated TFP growth shown in Figure 2, using a simple growth accounting approach, is

[3]The United States saw per capita income growth averaging 2.4 percent a year between 1950 and 1973; over the same period, per capita income grew on average by 5 percent a year in Germany; and by slightly more than 8 percent in Japan.

close to other more sophisticated techniques in the growth literature. Consistent with the results found in this paper, Jorgenson and Yip (2001) found that between 1960 and 1973 TFP growth accounted for more than half of the growth in output for France, Germany, and Japan, but somewhat less than half of output in the United States. The relative importance of TFP growth declined substantially after 1973. Jorgenson (2005) shows TFP growth of 0.5 for the United States, –0.1 for Germany, and 0.85 for Japan for the period 1995–2001. Amador and Coimbra (2007), using a dynamic translog stochastic production frontier (computed through Bayesian Statistical methods), found that France, Germany, and the United Kingdom moved to a new lower floor of TFP contribution in the last two decades. In contrast, the United States and Canada recorded slightly higher TFP acceleration after the mid-1980s.

East Asian growth has relied much more heavily on factor inputs, both labor and capital, and less on TFP growth than that of "Golden Age" Europe and the current rapid growth in transition countries. The estimates in this section also show that TFP growth has accounted for about 40 percent of the output for South Korea and for slightly more than half for China in the past two decades. There are very few countries around the world that were able to sustain rapid growth for more than 15 years with relatively low shares of investment in GDP. In Chile, factor accumulation accounted for two-thirds of the growth in 1986–95, and about 90 percent in 1996–2006. Ireland's impressive economic performance over the past two decades was also driven largely by factor inputs. On the other hand, India achieved its growth with relatively little emphasis on capital accumulation and more substantial gains in TFP. In that mix of gains, India differs from the East Asian economies.

A key question for prospective growth is whether the TFP gains achieved thus far have already eliminated most of the inefficiencies of central planning—and will therefore soon fade away. Sustaining productivity growth rates such as those experienced recently in some of the transition countries is difficult. Underutilized labor combined with the recent trend of faster capital accumulation is expected to play a more important role in the medium-term growth.

II. Panel Regression Approach

This section reports estimates of spare specifications of growth equation including new variables that are of particular importance to transition economies, such as recovery of lost output, and improvements in market reforms. The aim is to use up-to-date data and experiences of a large number of countries over long periods to identify key determinants of growth for transition economies and to form a view of growth prospects that complements the growth accounting exercise in Section I.

This section differs from previous empirical studies on the determinants of growth in the following aspects. First, the main focus is on transition economies using the latest information (1991–2006). Second, it analyzes a new set of explanatory variables, including output recovery index, and different measures of market reforms and institutional qualities that may have influenced variation in output and TFP over time and across countries. Third, it tests for endogeneity of some of the explanatory variables so that appropriate econometric methods can be chosen. Fourth, it assesses the importance of period-specific effects (in the form of world economic conditions) based on a large sample of countries that includes developing and developed countries.

Methodology and Data Issues

Although the main focus of the paper is the transition countries, a heterogeneous data set is also used including most advanced and developing economies. Such an approach would improve the statistical reliability of the results. The transition sample includes 12 CIS countries, three Baltic countries, five Central European countries, and six Southeastern European countries covering the period 1991–2006. The global sample consists of 123 countries with data spanning the period 1980–2006 (data for the transition economies have a shorter span). The use of a large panel of countries over an extended period of time allows sufficient freedom to enrich the menu of variables used on the determinants of growth, and improves the statistical reliability of the results. Pure cross-section, four-year nonoverlapping averages, and annual data are used to estimate the regressions.

Averaging the data over time eliminates short-term disturbances as well as business cycle effects from the data, while allowing one to test for long-run market reform dynamics. Failure to eliminate short-run dynamics typically leads to highly correlated time series and to gross overestimation of coefficients. The choice of four-year periods is dictated by the data time span for transition economies (1991–2006), which gives four observations for each country. For the global sample, each country is represented by seven observations. Unlike a pure cross-country regression using long-period average data, a panel regression (with each observation representing four- or five-year averages) provides additional information because it captures both time-series and cross-sectional information. The definition and sources of data are described in the data appendix.

To examine the relationship between market reforms, recovery of lost output, and growth, this section follows the standard empirical growth literature,[4] and uses the following linear growth regression model:

$$g_{it} = \beta Z_{it} + \lambda X_{it} + \mu_i + \nu_t + \varepsilon_{it}, \tag{1}$$

where g_{it}, the dependent variable, is the per capita real GDP growth rate or TFP growth rate in country i during the period t, Z is the vector of "core explanatory variables" that are believed to have contributed to the rapid

[4]See Mankiw, Romer, and Weil (1992), Islam (1995), and Barro and Sala-i-Martin (2004).

growth in transition countries (including recovery of lost output index, market reforms, and institutional quality). X comprises a set of control variables that are often used in the growth literature, including terms of trade, level of development as proxied by initial GDP per capita, inflation rate, fiscal balance, government size in the economy (as captured by government consumption to GDP ratio), and investment. μ_i is a country-specific unobservable effect, ν_t is a time-specific factor, and ε_{it} is the disturbance term. The paper also controls for time-specific growth effects emanating from changes in the external economic environment by including world cycle dummies.

The panel regressions for annual and four-year average are estimated using three different methodologies: (1) pooled ordinary least squares (OLS) (fixed, random effects, or seemingly unrelated regressions); (2) 2SLS; and (3) the generalized method of moments (GMM).[5] For the pure cross-section, the OLS could be considered as an efficient estimation technique.[6]

In Equation (1) the core explanatory variables (vector Z, which includes measures of market reforms and quality of institutions) may not be entirely exogenous. If the causality runs mainly from these variables to growth then the problem may be benign, but if it runs from growth to these variables then the problem is more severe. The Durbin-Wu-Hausman test suggests that endogeneity is present, albeit not very strong. This problem is addressed, to a certain extent, by using the 2SLS and the GMM techniques.

We need good instruments for market reforms, quality of institutions, and investment. The following variables are used as instruments: (1) lagged values of the exogenous explanatory variables; (2) commodity exports as share of total exports; (3) distance from the capital of the respective country to Brussels; (4) an index measure of ethnic fractionalization as measured by Alesina and others (2003); and (5) a period trend. Using more appropriate instruments would yield more efficient instrumental variable (IV) estimates. But given the large number of explanatory variables and the relatively small cross-sectional dimension (particularly in the ‘‘transition'' sample) by the standards of common panel data, overfitting should be avoided by working with a reduced number of IVs.

Weak instruments correspond to weak identification of some or all of the unknown parameters. Weak identification leads to GMM statistics with nonnormal distributions, even in large samples, so that conventional IV or GMM inferences are misleading. A rule of thumb for checking for weak instruments is the F-statistics, which is used to test the hypothesis that the

[5]The GMM panel estimator extracts consistent and efficient estimates of the impact of reforms and institutions on growth. It exploits the time-series variation in the data, accounts for unobserved country-specific effects, and controls for endogeneity of all the explanatory variables.

[6]Based on Monte Carlo simulations, Hauk and Wacziarg (2004) argue that taking account of all the advantages and limitations of the different estimation procedures, the cross-section OLS estimator that averages data over longer periods might be the most efficient.

coefficients on the instruments equal zero in the first stage of the 2SLS (see Stock, Wright, and Yogo, 2002). When there is a single endogenous regressor, a first-stage *F*-statistics of less than 10 indicates that the instruments are weak, in which case the 2SLS estimator is biased (even in large samples), and 2SLS *t*-statistics and confidence intervals are unreliable.

Determinants of Growth

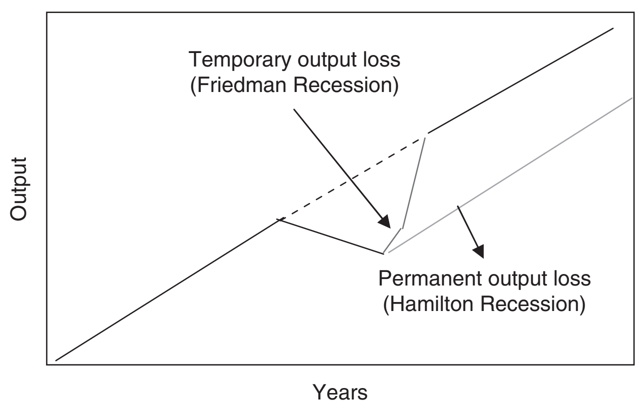
The explanatory variables used in this paper can be divided into four groups: (1) recovery of lost output or catch-up process (in the case of the "transition" sample); (2) market reforms and institutions; (3) other measures of macroeconomic and investment; and (4) external conditions.

Recovery of Lost Output

For the sample that includes only transition economies, the real GDP index (1990 = 100) of the previous period is used to test whether the amplitude of output recovery is influenced by the magnitude of the fall in output before recovery. The experiences of many countries show that usually sharp contractions in output due to crisis, wars, or other major shocks to the economy may be followed by strong growth that offsets the initial decline. This combined with corrective policies and structural reforms to reduce inefficiencies could spur strong economic recovery above the original trend line.

There are two models in the literature on the recovery of lost output following recessions, financial crisis, or a change in economic regime (Cerra and Saxena, 2005b). The first model (based on Friedman, 1993) suggests that recession can be characterized as a temporary fall in output. After the negative large "temporary" shock dissipates, output returns back to trend in a rapid-growth recovery phase (Figure 3). The second model (based on

Figure 3. Two Views of Lost Output Recovery



Hamilton, 1989) suggests that the stochastic trend in output undergoes regime switching between positive and negative states, resulting in a permanent output loss.

Negative shocks in theory impose only a temporary restraint on output, but may lead to rapid future growth that offsets the initial decline. First, negative shocks could stimulate political and economic reforms. Corrective policies could prompt an economic recovery above the original trend line if they reduce inefficiencies. Second, following Schumpeter's idea of "creative destruction," a sharp fall in output may cleanse the economy of inefficient firms, leading to higher productivity and economic growth (Caballero and Hammour, 1994). The core process of change comprises two elements: reallocation of resources from old to new activities (via closures and bankruptcies, combined with the establishment of new enterprises), and restructuring within surviving firms (via labor rationalization, product line change, and new investment). These can be thought of as the dynamic movements resulting from the establishment of new incentives and are reminiscent of the Schumpeterian concept of "creative destruction" by entrepreneurial activity, only with a much larger impact than what Schumpeter's model envisioned.

However, Cerra and Saxena (2005a) found that recessions or large contractions in output due to crisis, wars, or other reasons are in general not followed by high-growth recovery phases. They conclude that when output drops, it tends to remain well below its previous trend. The data used by Cerra and Saxena consisted of annual observations spanning 192 countries from 1960 to 2001, and thus their sample did not capture the recent strong growth in transition economies.

The recovery of lost output in the case of the transition sample is proxied by the real GDP index, with 1990 = 100. The index series are constructed using the real GDP growth rate estimates. For the global sample, the following indices are used: (1) 0 if initial real GDP was greater than 95 percent of its value in 1990; (2) 0.5 if initial real GDP was between 70 and 95 percent; and (3) 1 if initial real GDP was less than 70 percent.

Market Reforms and Institutional Development

Measures of market reforms for transition economies have been constructed by the World Bank (De Melo, Denizer, and Gelb, 1996), the European Bank for Reconstruction and Development (EBRD), and most recently by Campos and Horváth (2006). Critics have noted that proxies market reforms as measured by the EBRD are outcome-oriented rather than measuring inputs. In particular, there is a concern about the reliability of the EBRD scores, particularly during the early years of transition. In 2000, EBRD made an effort to backdate the indicators to 1990. This implies that the ratings for the early 1990s have to be treated cautiously, especially as these were the years in which information flows were limited. For robustness test, therefore, this paper also uses other measures of market reforms. In this regard, Table 2

Table 2. Evolution of Selected Growth Determinants in Transition Economies

	Catching-up Real GDP (1990 = 100)		Investment-to-GDP Ratio (in percent)		Market Reforms: EBRD index[1]			Market Reforms: Campos and Horváth[2]			Institutional Quality: ICRG[3]			Terms of Trade (2000 = 100)	
	1995	2006	1996–2000	2001–06	1995	2000	2006	1995	2000	2006	1995	2000	2006	1995	2006
Armenia	56	136	17	24	2.1	2.7	3.2	0.54	0.82	0.91	63	54	59	161	112
Azerbaijan	42	128	30	40	1.6	2.3	2.8	0.34	0.43	0.79	50	56	65	38	158
Belarus	66	138	25	25	1.9	1.7	2.0	0.32	0.24	0.33	60	56	61	99	111
Georgia	33	61	23	26	2.0	3.0	3.2	0.53	0.77	0.95	—	—	—	125	92
Kazakhstan	62	126	17	25	2.4	2.8	2.9	0.57	0.60	0.80	61	71	76	72	138
Kyrgyzstan	54	82	16	20	2.9	3.0	3.1	0.71	0.78	0.76	—	—	—	94	102
Moldova	38	52	19	21	2.2	2.9	2.9	0.23	0.40	0.57	65	62	62	86	83
Mongolia	89	143	32	38	1.8	2.4	2.7	—	—	—	69	69	68	117	130
Russia	60	96	17	19	2.8	2.7	3.0	0.27	0.42	0.62	52	54	68	78	155
Tajikistan	32	64	11	17	1.7	2.2	2.5	0.42	0.56	0.67	—	—	—	113	46
Ukraine	47	73	20	22	2.2	2.6	3.0	0.37	0.54	0.85	63	59	67	106	133
Uzbekistan	83	135	24	24	1.8	1.7	2.1	0.29	0.25	0.36	—	—	—	113	147
Estonia	76	162	27	30	3.4	3.7	3.8	0.63	0.78	0.87	71	75	77	76	162

Latvia	55	116	21	27	3.0	3.3	3.6	0.66	0.82	0.93	66	68	76	55	116
Lithuania	62	116	22	22	3.2	3.3	3.8	0.57	0.73	0.95	63	66	74	62	116
Czech Republic	99	130	29	29	3.5	3.6	3.8	0.63	0.76	0.85	80	78	78	99	102
Poland	90	139	21	24	3.4	3.6	3.7	0.44	0.63	0.90	78	74	76	98	105
Hungary	118	176	22	20	3.6	3.9	3.9	0.73	0.83	0.93	77	79	81	101	100
Slovak Republic	92	141	30	27	3.3	3.4	3.7	0.61	0.70	0.89	74	77	75	99	105
Slovenia	100	147	26	28	3.1	3.4	3.4	0.45	0.52	0.81	74	80	82	100	104
Albania	96	156	20	27	2.6	2.8	2.9	0.42	0.68	0.87	66	60	68	106	101
Bosnia & Herzegovina	28	95	—	22	1.2	2.1	2.6	—	—	—	—	—	—	—	99
Bulgaria	79	113	13	21	2.6	3.4	3.5	0.46	0.65	0.85	71	71	72	119	102
Croatia	81	119	23	27	2.9	3.3	3.5	0.51	0.74	0.85	66	69	76	94	96
Macedonia	77	98	16	16	2.2	2.7	2.9	0.45	0.74	0.85	—	—	—	105	98
Romania	93	119	20	23	2.5	3.2	3.4	0.37	0.57	0.80	68	66	69	93	95

Sources: IMF, World Economic Outlook database; EBRD, Transition Reports; Campos and Horváth (2006); and *International Country Risk Guide* (ICGR).

[1]Average of eight EBRD transition reform indicators (price liberalization, competition policy, banking reform, trade and foreign exchange system, large-scale privatization, small-scale privatization, governance and enterprise reforms, and infrastructure). The transition indicators range from 1 to 4.3, with 1 representing little or no change from a rigid centrally planned economy and 4.3 representing the standards of an industrialized market economy.

[2]Based on Campos and Horváth (2006). Simple average of internal liberalization, external liberalization, and privatization indices.

[3]The ICRG ratings reflect risk of the following components: government stability, socioeconomic conditions, investment profile, internal conflict, external conflict, corruption, law and order, ethnic tensions, democratic accountability, and bureaucratic quality.

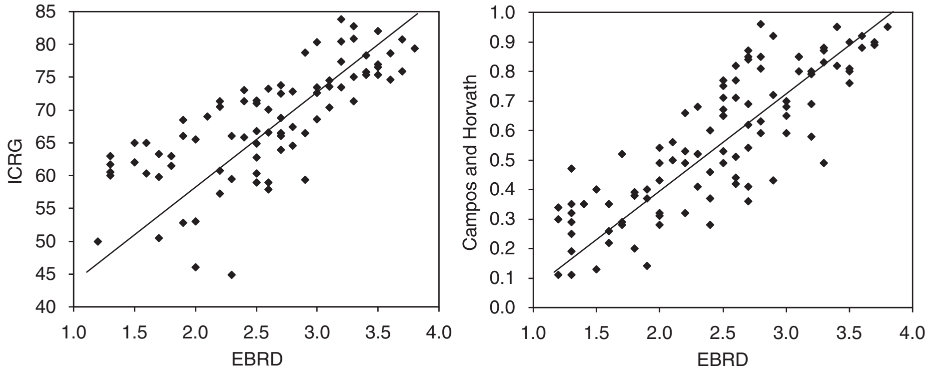
presents the development of market reforms in transition economies as measured by the EBRD and Campos and Horváth.

The overall EBRD score in Table 2 is the unweighted average of eight structural reform indicators: price liberalization, small-scale privatization, large-scale privatization, competition policy, trade liberalization, financial sector reform, governance and enterprise reforms, and infrastructure reform. The EBRD indicators range from 1 to 4.3, where 4.3 indicates that the country's structural characteristics are comparable to those prevailing on average in market economies, and 1 represents conditions before reform in a centrally planned economy with dominant state ownership of the means of production. However, it could be argued that the EBRD market reform scores are output-oriented rather than measuring inputs.

Campos and Horváth (2006) argue that the market reform indicators of the World Bank and EBRD are subjective and may be influenced by ex post reports of favorable or unfavorable performance raising the problem of possible endogeneity. But, in recent years, the EBRD Transition Reports have highlighted the main factors behind the change in their scores for each category of reform. They have constructed three categories of reform. The first captures internal liberalization reflecting the extent of price and wage liberalization. The second captures external liberalization efforts as reflected in the severity of trade barriers and capital controls. The third index captures the extent of privatization efforts. These three indices were constructed for 25 transition economies for all years between 1989 and 2000. Using Campos and Horváth's methodology this paper extended the market reform indices through 2006.

Institutional development is particularly important to investigate for the transition economies, where the institutions of central planning were a key constraint on growth in the early 1990s. Easterly (2003), and Rodrik, Subramanian, and Trebbi (2004) show that institutions are more robustly associated with faster growth than policies. Major institutional changes have taken place since the breakdown of the Former Soviet Union in 1991. Assessments of public institutions are mainly based on two indices: the political risk rating, *International Country Risk Guide* (ICRG), and the World Bank's governance indicators. The ICRG comprises 22 variables in three subcategories of risk: political, financial, and economic, for 143 countries and is available since 1984. The political risk rating, which is used in this paper, includes five indices of perceptions of government stability, democratic accountability, law and order, quality of bureaucracy, and corruption in government. The World Bank's governance indicators comprise six measures of institutional development: voice and accountability, political instability, government effectiveness, regulatory burden, rule of law, and control of corruption. Unfortunately, the World Bank's governance data do not exist prior to 1996, and therefore are not used in this paper. It should be noted that the market reform as measured by the EBRD and Campos and Horváth and the ICGR are highly correlated with each other (Figure 4).

Figure 4. Different Measures of Market Reforms and Institutional Quality are Correlated
(Four-year nonoverlapping averages, 1991–2006)



Based on both the EBRD and Campos and Horváth scores, Table 2 suggests that progress has been achieved in market reforms in most transition economies. But the CIS still remain far behind the five Central European and the three Baltic countries. In general, reform is most advanced in the privatization of small-scale enterprises, the liberalization of foreign trade and exchange, and the elimination of price controls. Structural reforms are least advanced in the regulation and supervision of the banking and financial sector, the development and enforcement of competition, and the reform of governance in both the private and the public sectors. Among the CIS countries Armenia and Georgia so far achieved an average market reform index as measured by the EBRD score of more than three.[7] Progress in market reforms has been particularly slow in Belarus and Uzbekistan. Turkmenistan virtually did not reform its economy with the exception of some small-scale privatization and price liberalization. An important question is whether the ICRG index picks some of the dimensions specifically related to transition. In this regard, Figure 4 compares the ICRG index with the EBRD or Campos and Horváth indices. It shows that institutional quality as measured by the ICRG is highly correlated with the measure of EBRD's market reforms, which is also highly correlated with market reforms measures based on Campos and Horváth.

Macroeconomic Stabilization, Investment, and Convergence

There is an array of policy determinants of growth (see Barro and Sala-i-Martin, 2004). In this paper the impact of macroeconomic

[7]The EBRD market reform index ranges from 1 to 4.3, where 1 represents conditions before reform in a centrally planned economy with dominant state ownership of the means of production, and 4.3 indicates that the country's structural characteristics are comparable to those prevailing on average in market economies.

stabilization is measured by the logarithm of the inflation rate,[8] and the overall fiscal balance as a ratio of GDP. Inflation is a policy result, but the fiscal balance refers more to the policy itself. It should be noted that the improvement in the overall fiscal position in Azerbaijan, Kazakhstan, Russia, and Uzbekistan was largely due to the substantial increase in government revenues from oil, gas, and other major commodities. Fiscal policy may influence growth through the size of government in the economy, as measured by the ratio of government consumption to GDP. Higher government consumption is believed to reduce growth prospects. This effect is normally associated with the crowding out of private sector investment, higher rent-seeking behavior, and distorted market incentives including higher taxation.

There is little disagreement in the general growth literature that investment is a major engine of growth. In the transition economies, with a history of excessive capital accumulation and inefficient use, the role of investment in the initial recovery phase (perhaps through the late 1990s) was relatively less important. In recent years, however, there has been some increase in the investment ratio, albeit from a very low level. Most of the increase in investment has been in the hydrocarbon and metallurgy sectors.

In the case of the full sample, the paper also considers convergence as one of the determinants of growth. Most neoclassical growth models have shown that the potential for economic growth rate also depends on a country's level of development as proxied by the initial per capita income (Barro and Sala-i-Martin, 2004). The coefficient for this variable is expected to be negative, implying that poor countries tend to grow faster than richer countries as each country converges toward its steady state.

External Conditions

Economic activity in a country is also affected by external conditions. The literature provides ample evidence of the transmission via international trade and external financial flows (see Mendoza, 1997). The change in the terms-of-trade index is included to account for possible exogenous shocks in international commodity prices that may have an impact on per capita growth. This index is derived from export prices relative to import prices. Terms-of-trade shocks capture changes in both the international demand for a country's exports and the cost of production and consumption inputs (Barro and Sala-i-Martin, 2004). This variable may also be included in

[8]A high rate of inflation is harmful to growth because it raises the cost of borrowing and thus lowers the rate of capital investment. At the same time, highly variable inflation makes it difficult and costly to forecast accurately costs and profits and hence investors and entrepreneurs may be reluctant to undertake new projects. Likewise, given that financial resources in the form of domestic savings and loans are limited, a larger fiscal deficit will mean that more of those limited resources must be devoted to financing the budget deficit. Fewer resources will thus be available for private sector investment.

the list of IVs because its movement depends primarily on world conditions and therefore is largely exogenous with respect to per capita growth for an individual country.

Several empirical studies have found a positive and significant link between improvement in the terms of trade and economic growth (Fisher, 1993; Mendoza, 1997). Barro (1997) notes that if the quantities of domestically produced goods do not change, then an improvement in the terms of trade raises real gross domestic income, but does not affect real GDP. Movements in real GDP occur only if shifts in the terms of trade bring about a change in domestic employment and output.

Estimation Results

The correlation matrix (Table 3) of the explanatory variables indicates high correlation (of more than 0.50) in the following explanatory variables: the market reform index, institutional quality, inflation rate, and per capita income. This implies that if these variables are included in the same regression, the estimated coefficients may not be individually reliable due to high multicollinearity.

First-Stage Estimation of Endogenous Variables

Table 4 shows the first-stage estimation results. Based on Stock and Yogo (2005) criteria, the IVs used to estimate the exogenous component of market reforms and institutions are strong (exceed the threshold values by a large margin).[9] The estimated coefficients for raw material exports, distances from Brussels, civil conflict or war dummy, and period trend have the expected sign and are highly significant in the regressions where measures of market reform are used as the dependent variables, columns (a) and (b). Together the exogenous variables explain about 70 percent of the variation in the scores for market reforms. In column (c), the exogenous variables explain about 50 percent of the variation in institutional development as measured by the ICGR scores. Also, specification tests confirm the validity of the IVs. The test of the overidentifying restrictions cannot be rejected, while the null hypothesis that the excluded exogenous variables do no explain market reform or institutional quality scores is strongly rejected.

[9]Stock and Yogo (2005) develop formal tests based on the *F*-statistic for the null hypothesis: (1) the bias of 2SLS is $>$ 10 percent of the bias based on OLS. The *F*-test rejects the null of weak instrumental variables at the 5 percent level if $F > 10.3$. They also consider the null hypothesis that (2) the null rejection rate of the nominal 5 percent 2SLS *t* test concerning $\hat{a}$ has a rejection rate 10 percent or greater. In this case, the *F*-test rejects the null of weak IVs at the 5 percent level if $F > 24.6$. The adverse effect of weak IVs on 2SLS depends on the degree of endogeneity present as measured by P_{uv}, the correlation between the structural and reduced form errors u and v. But P_{uv} is difficult to estimate precisely when the IVs are weak. In particular, P_{uv} cannot be consistently estimated under weak IVs asymptotics. The F test of Stock and Yogo (2005) is designed to be valid for any value of P_{uv}. For more details on the *F*-test, see Appendix II.

Table 3. Correlation Coefficient Matrix, Transition Sample

	Dependent Variables		Market Reforms		Institutions								
	Per capita growth	TFP growth	EBRD reform index[1]	Campos and Horváth reform[2]	World Bank institutional quality[3]	Political risk rating (ICRG)[4]	GDP Recovery Index	Investment to GDP	Fiscal Balance to GDP	Government Consumption to GDP	Inflation Rate	Terms of Trade	Per Capita Income
Per capita growth	1.00												
TFP growth	0.93	1.00											
EBRD reform index	0.63	0.54	1.00										
Campos and Horváth reform	0.67	0.61	0.83	1.00									
World Bank institutional quality	0.14	0.11	0.68	0.44	1.00								
Political risk (ICGR)	0.25	0.21	**0.72**	0.42	**0.74**	1.00							
GDP recovery index	−0.12	−0.12	0.28	0.11	0.44	0.47	1.00						
Investment to GDP	0.37	0.27	0.28	0.25	0.16	0.28	0.29	1.00					

Fiscal balance to GDP	0.62	0.54	0.40	0.31	0.14	0.23	0.01	0.28	1.00				
Government consumption to GDP	−0.09	−0.19	0.00	−0.22	0.18	0.01	0.10	0.11	0.16	1.00			
Inflation rate	−0.82	−0.74	**−0.72**	**−0.69**	−0.39	−0.37	−0.09	−0.32	−0.49	−0.01	1.00		
Terms of trade	0.35	0.40	0.18	0.23	0.03	0.16	0.05	0.11	0.33	−0.09	−0.24	1.00	
Per capita income	0.03	0.04	**0.52**	0.26	**0.70**	**0.67**	**0.64**	0.20	0.25	0.18	−0.24	−0.04	1.00

Source: Author's calculations based on 26 transition economies covering the period 1996–2006.

Note: TFP = total factor productivity.

[1]Simple average of eight EBRD transition reform indicators (price liberalization, competition policy, banking reform, trade and foreign exchange system, large-scale privatization, small-scale privatization, governance and enterprise reforms, and infrastructure). The transition indicators range from 1 to 4.3, with 1 representing little or no change from a rigid centrally planned economy and 4.3 representing the standards of an industrialized market economy.

[2]Based on Campos and Horváth (2006). Simple average of international liberalization index, external liberalization index, and privatization index.

[3]Average of six institutional concepts: voice and accountability, political stability, government effectiveness, regulatory burden, rule of law, and control of corruption. Each of these indicators is distributed normally, with a mean of zero and a standard deviation of one. The scores lie between –2.5 and 2.5, with higher scores corresponding to better outcome.

[4]ICRG = International Country Risk Guide. The ICGR ratings reflect risk ratings of the following components: government stability, socioeconomic conditions, investment profile, internal conflict, external conflict, corruption, law and order, ethnic tensions, democratic accountability, and bureaucratic quality.

Table 4. First-Stage Estimation Results
(Exogenous determinants of reforms, institutions, and investment)

	(a)	(b)	(c)
Dependent Variable →	Market Reform Index		Institutions
	EBRD	Campos and Horváth	ICRG
Log (commodity exports as percent of total exports)[1]	−0.18**	−0.19**	
Log (distance to Brussels)[2]	−0.36**	−0.15**	−3.01*
Civil conflict or war dummy	−0.44**	−0.13**	
Ethnic fractionalization[3]			−11.53**
Trend[4]	0.30**	0.14**	2.04**
R^2 (adjusted)	0.76	0.72	0.51

Source: Author's estimates.

Note: ICRG = International Country Risk Guide. * and ** indicate significant at the 5 and 1 percent confidence levels, respectively.

[1]Share of fuel, ores, metal, and agricultural raw exports in total exports.

[2]Measured in 1,000 km and refers to distance from capital of respective country to Brussels.

[3]Probability that two randomly selected individuals in a country are not from the same ethnic group.

[4]Values of 1 for 1991–94; 2 for 1995–98; 3 for 1999–2002; and 4 for 2003–06.

Resource endowments and entrenchment of communist elite have together influenced the degree of market reforms or institutional building during the transition period. Hoff and Stiglitz (2004) suggest that a higher ratio of natural resources to GDP or total exports decreases the political constituency for institutional development. In transition economies with less dependence on natural resources (Albania, Armenia, Moldova, Georgia, and the Baltics), the elites had fewer incentives to clinch to power and thus were more likely to allow the introduction of market reforms and good governance. Beck and Laeven (2006) show that transition countries that rely more on natural resources (Russia, Azerbaijan, Kazakhstan, Tajikistan, and Uzbekistan) experienced less market reform or institutional development during the transition process.

Also, distance from the capitals of the transition economies to the center of Western Europe (say, Brussels) is widely acknowledged to be a primary force shaping the opportunity for interaction among states including eagerness for market reforms and institutional development. Distance from Brussels is also highly and positively correlated with the years under socialism. In contrast to the Baltics and Central European economies, most of the CIS economies had long periods of communist rule and their capitals are relatively far away from the center of Western Europe. Old elites of the Central and Eastern European and Baltic economies with closer proximity to the capital of Western European countries had fewer possibilities to maintain their power. The prospects of future European Union (EU) membership have fostered institutional building, both through political incentives and through assistance from the major EU members (Roland and Verdier, 2003). Easterly and Levine (1997) show that ethnic diversity tends to reduce the provision of public goods, including the institutions that support business transactions. Also, Alesina and others (2003) conclude that ethnic fractionalization is likely to be an important determinant of economic success, both in terms of output and quality of institutions. Ethnic fractionalization fosters rent-seeking and might not be conducive to the building of strong market institutions.

Results of the Transition Sample

The estimation results for the transition sample are reported in Table 5. The *F*-tests demonstrate that the instruments are related to the endogenous variables that are instrumenting for, at high levels of statistical significance. Overall, the fit is good for this type of panel data. In all cases, the variables have the theoretically expected sign, but their magnitude and significance differs depending on the variables included, frequency of the data used (annual or period averages), and the estimation techniques. There are several interesting findings:

First, the estimated coefficient on the recovery of lost output is negative, as expected, and highly significant both in the per capita and TFP growth regression equations. The recovery of lost output effect is sizable: according to the point estimate, given that the average real GDP index in 1996 was about 50 for the CIS (1990 = 100) as compared with 100 in the Central European economies, the difference in per capita growth is expected to be about 3 percentage points in favor of the CIS, assuming other things are equal.

Second, there is a strong link between progress in market reforms as measured both by the EBRD and Campos and Horváth (2006) reform indices on the one hand, and growth in per capita real GDP or TFP on the other hand. Unlike Fidrmuc (2003) and Lawson and Wang (2004) but in agreement with Falcetti, Lysenko, and Sanfey (2005), the estimated coefficients for the EBRD reform index in this study are always positive and highly significant. The magnitude of the estimated coefficient implies that if the average EBRD score for the CIS countries in 2006 were close to the three Baltics then the average growth would have been about 3 percentage points higher than the outcome for 2001–06. The estimated coefficient of market reform (EBRD or Campos and Horváth scores) is larger in magnitude and highly significant when other stabilization policy variables (inflation rate or fiscal balance) are excluded from the estimated equation (see column 1 of Table 5).

Third, unlike previous studies on transition economies, the results suggest that investment is one of the variables that have contributed to the recent rapid growth. The regressions are also estimated without the investment

Table 5. Estimation Results for the Transition Sample, 1991–2006

Data →	Four-Year Nonoverlapping Averages							Annual Data	
Estimation method→	Fixed effects[1]			2SLS		GMM		GMM	GMM
	Second Stage: Dependent Variable is Per Capita Real GDP Growth								
Recovery of lost output[2]	−0.15**	−0.10**	−0.11**	−0.07**	−0.07**	−0.07**	−0.06**	−0.04**	−0.05**
Investment/GDP	0.26**	0.31**	0.33**	0.22**	0.24**	0.24**	0.31**	0.22**	0.22**
Log (inflation rate)		−1.54**	−1.93**	−2.22**	−2.08**	−1.93**	−1.97**	−1.35**	−1.48**
Fiscal balance/GDP		0.36**	0.31**	0.32**	0.36**	0.29**	0.32**	0.52**	0.24**
Government consumption/GDP		−0.23**	−0.26**	−0.18*	−0.16*	−0.14*	−0.09	−0.17*	−0.14*
EBRD reform index[3]	9.83**	3.63**		2.22**		1.99**			
Campos and Horváth reform index[4]			7.66**		5.27**		3.66**	5.03**	
Institutions ICRG index[5]									0.12**
Terms of trade growth	0.12**	0.06**	0.06**	0.08**	0.08**	0.07**	0.07**	0.07**	0.08**
F-test (*p*-value)[6]				0.000**	0.000**	0.000**	0.000**	0.000**	0.018*
Observations	110	110	99	110	99	110	99	326	275
R^2 (adjusted)	0.83	0.90	0.91	0.79	0.80				
	Second Stage: Dependent Variable is TFP Growth								
Recovery of lost output[2]	−0.07**	−0.06**	−0.04**	−0.11**	−0.05**	−0.05**	−0.04**	−0.03**	−0.03**
Log (inflation rate)								−0.95*	−1.00**
Fiscal balance/GDP		0.37*	0.21*	0.31*	0.28**	0.18**	0.17**	0.23**	0.20**
Government consumption		−0.26*	−0.15*	−0.22*	−0.18*	−0.29**	−0.28*	−0.04	−0.12*

EBRD reform index[3]	5.01**	4.71**		3.31**		3.12**			
Campos and Horváth reform index[4]			10.22**		4.42**		4.66*	3.38**	
Institutions ICRG index[5]									0.13**
Terms of trade growth	0.09**	0.07**	0.06**	0.07*	0.07*	0.04*	0.06**	0.03**	0.04*
F-test (*p*-value)[6]				0.000**	0.000**	0.000**	0.000**	0.000**	0.022*
Number of observations	110	110	99	110	99	110	99	326	275
R^2 (adjusted)	0.49	0.62	0.65	0.58	0.61				

Source: Author's estimates.

Note: TFP = total factor productivity; GMM = generalized method of moments; 2SLS = two-stage least squares. * and ** indicate significant at the 5 and 1 percent confidence levels, respectively.

[1]The choice of a fixed effects versus a random effects specification is justified by a Hausman test.

[2]Initial real GDP index (1990 = 100), constructed from the real GDP growth rates.

[3]Average of eight EBRD transition reform indicators (price liberalization, competition policy, banking reform, trade and foreign exchange system, large-scale privatization, small-scale privatization, governance and enterprise reforms, and infrastructure). The transition indicators range from 1 to 4.3, with 1 representing little or no change from a rigid centrally planned economy and 4.3 representing the standards of an industrialized market economy.

[4]Average of internal liberalization index, external liberalization index, and privatization index.

[5]ICRG = International Country Risk Guide. The ICGR ratings reflect risk ratings of the following components: government stability, socioeconomic conditions, investment profile, internal conflict, external conflict, corruption, law and order, ethnic tensions, democratic accountability, and bureaucratic quality.

[6]The null-hypothesis of the *F*-test is that the exogenous excluded variables do not explain market reforms or institutional development in the first stage. In addition to the exogenous instrumental variables, lagged values of the policy variables are also used as instruments in the annual data regressions.

variable. The reason is that the interpretation of the role of this variable is problematic even after the endogeneity problem is addressed. Investment could be capturing the effects of structural reforms that are difficult to quantify, or are already included in the EBRD market reform index. Investment could also change for reasons other than those related to reforms (for example, the large investment in the oil and gas sectors in Azerbaijan and other resource-rich countries).

The estimated coefficient for the terms-of-trade index is also positive and highly significant. Its magnitude is larger when the fiscal balance variable is excluded from the estimated equation, due to the relatively strong correlation between terms of trade and fiscal balance (see Table 3). Favorable terms of trade in the commodity exporter transition countries have shifted their fiscal deficits to significant surpluses (particularly in Azerbaijan, Kazakhstan, and Russia).

As for other explanatory variables, sound macroeconomic policies (including smaller fiscal deficits and government size in the economy) are associated with higher growth in per capita and in TFP growth. It should be noted that the fiscal coefficient is quite large and robust to changes in the specification of the equation and the estimation technique.

Results of the Global Sample

Table 6 shows that all included variables have the right sign and are significant at the 1 percent level, except for the terms of trade and government which is marginally significant. The regional dummies were used to test the hypothesis that different regimes may have characteristics that affect growth differently. This is confirmed with respect to Southeast Asia, which, on average, performed better than did other regions in the period under consideration. The coefficients of the African and Latin American dummies are negative, but insignificant at the 5 percent level.

The coefficient of the output recovery index is highly significant and robust to different estimation techniques. The coefficient of 2.77 from the four-year nonoverlapping averages using the GMM technique implies that on average growth in the CIS economies was between 2 to 2.5 percentage points higher than in other countries, depending on the period under consideration. To recall, the output recovery index takes the value of 1 for countries whose initial level of real GDP was less than 70 percent of the real GDP of 1990; 0.5 if initial real GDP was between 70 and 95 percent; and 0 if initial real GDP was greater than 95 percent.

The estimated coefficient on the quality of institutions as measured by the ICRG political index is robust in magnitude and continues to be significant under different estimation techniques. Its significance increases when some of the macroeconomic policy variables are excluded from the right-hand side of the regression equation. This may reflect the impact of institutions on policy sustainability variables and indicate that institutions play a dominant role in explaining cross-country differences in growth. More efficient institutions

Table 6. Estimation Results for the Global Sample, 1980–2006

Data →	Pure Cross-Section		Four-Year Averages			Annual Data	
Method of estimation →	OLS	2SLS	OLS	2SLS	GMM	2SLS	GMM
Log (per capita income)	−0.84**	−0.67**	−0.54**	−0.51**	−0.45**	−0.32*	−0.37*
Output recovery index[1]	2.60**	2.48**	1.21**	2.68**	2.77**	1.88**	2.43**
Investment/GDP	0.14**	0.13**	0.11**	0.11**	0.11**	0.12**	0.13**
ICGR	0.11**	0.12**	0.15**	0.15**	0.14**	0.14**	0.10**
ICGR × log (per capita income)[2]	−0.02**	−0.01**	−0.01**	−0.01**	−0.01**	−0.01*	−0.01**
Fiscal balance/GDP	0.04*	0.05*	0.05*	0.05**	0.09**	0.13**	0.13**
Log (inflation rate)	−0.05	−0.02	−0.22*	−0.19*	−0.29*	−0.42**	−0.51**
Population growth	−0.64**	−0.71**	−0.77**	−0.75**	−0.62**	−0.46**	−0.48**
Terms of trade growth	0.08**	0.07**	0.05**	0.03*	0.03*	0.02**	0.03**
World cycle dummies							
1987–1990			0.39	0.60	0.54*	0.34	0.38
1991–94			−0.27	−0.25	−0.52	−0.30	−0.27
1995–98			0.23	0.06	0.28	0.36	0.72
1999–2002			−0.17	−0.19	−0.27	−0.33	−0.07
2003–06			1.46**	1.52**	1.30**	1.41**	1.39**
Regional dummies							
Latin America	−0.58	−0.42	−0.44	−0.49	−0.57	−0.19	−0.29
Sub-Saharan Africa	−0.52*	−0.45	−0.83	−1.01*	−1.03*	−0.35	−0.35
South and East Asia	1.41**	1.67**	1.38**	1.27**	1.14**	1.78**	1.69**
F-test (*p*-value)[1]		0.015*		0.023*	0.019*	0.013*	0.012**
Observations	122	122	638	610	610	2356	2356
R^2 (adjusted)	0.82	0.78	0.45	0.44		0.29	

Source: Author's estimates.

Note: OLS = ordinary least squares; 2SLS = two-stage least squares; GMM = generalized method of moments. * and ** indicate significant at the 5 and 1 percent confidence levels, respectively. The following indices are used: 0 for countries with initial real GDP of more than 95 percent of the real GDP in 1990; 0.5 if real GDP was between 70 and 95 percent; and 1 if it was less than 70 percent. When per capita income and institutional quality are both low, the ability to take advantage of growth opportunities is limited. This effect is captured by the composite convergence term. A negative coefficient for the composite term implies that institutional quality improves, convergence accelerates.

[1]The null-hypothesis of the *F*-test is that the exogenous excluded variables do not explain institutional development in the first stage. In addition to the exogenous instrumental variables, lagged values of the policy variables were used as instruments in the annual data regressions.

allow an economy to produce the same output with fewer inputs; bad institutions lower incentives to invest, to work, and to save. The magnitude of the estimated coefficient implies that if the average ICGR for the CIS countries was close to the average for Southeast Asian economies, then the average growth would have been about 1.5 percentage points higher than the outcome for the last decade. The interaction term of ICRG with the log of per capita income is negative and significant, suggesting that lower income countries would benefit more than higher income countries for the same improvement in institutional quality.

The estimated negative coefficient for the population growth implies that faster population growth is associated with slower per capita real GDP growth. Lower population growth means more capital accumulation per worker and hence higher productivity growth. However, this should be interpreted with caution. Continued slow growth in the CIS, particularly in Belarus, Russia, and Ukraine, will be accompanied by changes in age structures, which in the long term could adversely impact the growth prospects.

More importantly, changes in the external environment over the past few years—as captured by the world cycle dummy for 2003–06—show clear favorable growth effect that was substantial and statistically significant. The recent favorable environment explains on average at least 1 percentage point increase in per capita output in most countries. This latter could reflect the rapid progress in technological innovation worldwide, lower interest rate, and easier access to capital markets for most developing and transition economies.

Robustness of the Results

The findings in this paper are robust to different econometric specifications and estimation techniques, using different measures of market reforms, and different samples.

A first estimation problem faced in this study is the decision of which explanatory variables to include in the growth equation. Variables could be significantly correlated with growth depending on which other variables are held constant. This is because economic theories are still not precise enough to decide on the determinants of growth. The high cross-correlation among some of the explanatory variables is also a problem (Table 3). For example, combining different sets of variables one finds that x_1 is significant when the regression includes x_2 and x_3, but becomes insignificant when x_4 is included or x_2 excluded. In particular, when the inflation rate is included in the regressions, the estimated coefficients on the market reform measures become much smaller. In general, however, the conclusion that recovery of lost output, macroeconomic stabilization, and market reforms significantly contributed to the rapid growth in 2001–06 appears robust to the alternative estimation methodologies and the choice of control variables.

A second concern with the estimated coefficients is the possible sensitivity of the results to the assumption about the form of the growth regression. In particular, the explanatory variables in Equation (2) enter the growth equation regression linearly and independently. This reflects an ad hoc assumption that the marginal effect of a change in explanatory variable is constant, both across different levels of the variable and across different economies. In this regard, this paper tested for the robustness of the results by allowing for two types of nonlinearities for the explanatory variables of interest in the panel regression equation. Thus, the paper includes a squared term for the EBRD or the Campos and Horváth measure of market reforms and the proxy for the recovery of lost output variables in the regression specification:

$$g = \beta_1 Z + \beta_2 Z^2 + \lambda X + \mu_i + \nu_t + \varepsilon_{it}. \tag{2}$$

The question of interest is whether the coefficient estimate β_1 remains robust when a squared term is included (a secondary question is whether β_2 is itself robust). Allowing for the inclusion of a squared term, the results show that the EBRD or the Campos and Horváth market reform and the recovery of lost output variables remain robust. In addition, the coefficients of the squared terms of these two variable (β_2) are significant and have opposite signs as compared with β_1.

Another possibility is that the partial effect of a variable on growth varies over different levels of development. For example, the marginal effect of market reforms (as measured by the EBRD score) could be quite different in Armenia than in Slovenia. One way to capture such linearity is to include an interaction term between the variable of interest and a measure of the country's level of development (such as per capita income) in the regression specification. That is:

$$g = \beta_1 Z + \beta_2 Z * log(Y_0) + \lambda X + \mu_i + \nu_t + \varepsilon_{it}, \tag{3}$$

where Y_0 measures the initial GDP per capita in purchasing power parity (PPP) U.S. dollars. Again the key question is whether β_1 becomes robust when the interaction term is included. Again the core explanatory variables remain robust. Combined with the results from Equation (2), this suggests some important nonlinearities in the correlation between market reforms and growth. The negative interaction term indicates that market reform has less of an effect at higher levels of development.

As an additional robustness check, which also sheds light on differences across country groups, the regressions were run on different subsamples—developed countries, transition, and low-income developing countries. The resulting estimated coefficients do show variations across the groups using a core set of explanatory variables (market reforms, institutions, and terms of trade). Although the variations behave directionally the same across country groups, the size and significance of coefficients are modestly different. Thus, there is support both for the unity of growth drivers and for the variation of impact in terms of their magnitude.

Contribution to Growth Changes

Changes in Growth Rates over Time

With the exception of Kyrgyzstan, all transition economies grew faster in 2001–06 than in 1996–2000. On average, the region grew by 5.3 percentage points a year faster in the latter period. In this regard, the estimated regressions can provide a useful decomposition of the importance of the various factors in explaining differences in growth between the two periods. For this, both the estimated coefficients from the main regression (based on the results in Table 5 using the GMM) and the actual values of the explanatory variables for the two periods under consideration are used. The objective here is to assess the contribution of the change in each category of explanatory variables to a country's fitted growth equation. The difference between the average country growth performance in 2001–06, denoted by g_1, and average growth performance in the same country in the previous period (1996–2000), denoted by g_0, can be expressed as follows:

$$g_1 - g_0 = \beta[Z_1 - Z_0] + \lambda[X_1 - X_0] + \gamma[W_1 - W_0] + \varepsilon_{it}. \tag{4}$$

Z is the vector of "core explanatory variables" (recovery of lost output index, EBRD measure of market reforms, investment, and terms of trade). X comprises a set of control variables including fiscal balance, inflation rate, and government consumption. W can be interpreted as an exogenous world environment. It captures the extent to which unaccounted international exogenous factors related to growth (such as productivity of new inventions in 2001–06). ε_{it} is the residual (the difference between the actual and the predicted change in growth).

The results of this approach are reported in Table 7. The first two columns show the actual and the fitted changes in the growth rates between 2001–06 and 1996–2000. The effects of changes in external factors such as trade and other global favorable environment factors are shown in columns (3) to (4). The combined impact of macroeconomic stabilization and reforms (specifically lower inflation, improvement in the fiscal position, progress made in market reforms as measured by the EBRD, and smaller size of government in the economy as measured by the government consumption to GDP ratio) are reported in column (5).

The estimates predict changes in the per capita growth rates quite well for the CIS, and the Central European economies but less well for the Southeastern European economies. For the CIS as a whole, 2.9 percentage points of the 4.9 predicted increase in growth is explained by improvement in macroeconomic stabilization and reforms. Had the market reform been deeper in the CIS, its impact on growth would have been correspondingly larger when multiplied by the estimated marginal growth effect. For example, if reforms in the CIS had attained the levels observed in the Baltics or in the Central European economies, the resulting aggregate growth acceleration impact would have been about 2 percentage points higher.

Changes in Growth Rates Across Regions

Another key advantage of the panel sample in this paper is that it permits the employment of an alternative standard of comparison, relying on cross-regional comparative analysis, to supplement the country-by-country time-series dimension. In this case, the unit of analysis is a comparison of regional aggregates. Specifically, the focus is on explaining the sources of the growth

Table 7. Decomposition of Growth Increase between Periods

(2001–06 compared with 1996–2000, in percentage points)

			Contribution to Predicted Change in Growth Rates						
			External						
	Actual Change in Growth Rates	Predicted Change in Growth Rates	Terms-of-trade shock	World cycle	Total external factors	Stabilization and reforms	Investment	Recovery of lost output	Residuals
	(1)	(2)	(3)	(4)	(3) + (4)	(5)	(6)	(7)	(8)
CIS	5.2	4.9	0.6	1.0	1.6	2.9	1.3	−0.4	0.3
Armenia	7.2	5.5	−0.6	1.0	0.4	3.9	2.2	−0.6	1.7
Azerbaijan	8.9	7.2	2.1	1.0	3.1	1.8	2.9	−0.5	1.8
Belarus	1.5	1.9	0.6	1.0	1.6	1.5	0.0	−0.8	−0.4
Georgia	1.7	3.4	−0.4	1.0	0.6	3.0	1.1	−0.8	−1.7
Kazakhstan	7.9	7.3	1.6	1.0	2.6	3.1	2.4	−0.5	0.5
Kyrgyztan	−2.0	3.7	0.3	1.0	1.3	2.8	0.9	−0.8	−5.7
Moldova	9.0	5.1	−0.4	1.0	0.6	3.4	0.5	0.4	3.8
Mongolia	3.3	6.0	0.3	1.0	1.3	3.7	1.9	−0.6	−2.7
Russia	4.7	5.6	1.1	1.0	2.1	3.4	0.4	−0.2	−1.0
Tajikistan	8.5	5.1	−0.6	1.0	0.4	2.8	1.9	0.1	3.4
Ukraine	9.5	5.9	1.0	1.0	2.0	2.4	0.6	0.7	3.6
Uzbekistan	2.4	1.8	1.1	1.0	2.1	0.8	−0.1	−0.7	0.6

Table 7 *(concluded)*

Baltics	2.9	3.7	0.3	1.0	1.3	2.5	0.9	−0.7	−0.8
Estonia	2.6	4.0	0.3	1.0	1.3	2.9	1.1	−0.9	−1.4
Latvia	3.1	4.6	−0.1	1.0	0.9	1.7	1.8	0.2	−1.6
Lithuania	3.0	2.6	0.7	1.0	1.7	2.9	−0.1	−1.3	0.4
Central Europe	0.3	0.3	0.2	1.0	1.2	0.9	0.0	−1.2	0.0
Czech Republic	2.6	1.8	0.5	1.0	1.5	0.9	0.0	−0.4	0.2
Hungary	0.3	−0.1	−0.1	1.0	0.9	−0.1	0.9	−1.2	−1.3
Poland	−1.9	−1.6	0.3	1.0	1.3	0.7	−0.7	−1.9	−3.1
Slovak Republic	1.5	0.3	−0.1	1.0	0.9	2.0	−1.0	−1.0	−0.3
Slovenia	−1.0	1.0	0.2	1.0	1.2	1.0	0.7	−1.3	−4.0
Southeast Europe	2.5	4.0	0.1	1.0	1.1	2.3	1.4	−0.4	−1.5
Albania	−0.5	3.0	−0.2	1.0	0.8	2.5	2.1	−1.6	−3.5
Bulgaria	5.5	8.0	−0.3	1.0	0.7	4.3	2.7	0.2	−2.5
Croatia	1.3	1.8	−0.2	1.0	0.9	0.8	1.4	−0.8	−0.5
Macedonia	−1.2	−0.3	−0.2	1.0	0.8	0.1	0.0	−0.7	−0.9
Romania	7.3	6.8	1.2	1.0	2.2	3.1	1.0	0.3	0.5

Source: Author’s estimates.

difference between the 12 CIS economies on the one hand, and the eight Central European and Baltic countries, and 11 Southeast Asian economies on the other hand. The comparison here is based on the growth performance over eight years (1999–2006). The difference between the average growth for the CIS region, denoted by g_{CIS}, and average growth for the Central European and Baltic countries, denoted by g_{CEB}, is expressed in Equation (5) using the estimated coefficients from the transition sample, but the difference between the CIS and the Southeast Asian (g_{SEA}) economies is expressed in Equation (6), using the relevant estimated coefficients from the global sample.

$$g_{CIS} - g_{CEB} = \beta[Z_{CIS} - Z_{CEB}] + \lambda[X_{CIS} - X_{CEB}] + \varepsilon_{it}, \tag{5}$$

$$g_{CIS} - g_{SEA} = \beta[Z_{CIS} - Z_{SEA}] + \lambda[X_{CIS} - X_{SEA}] + \varepsilon_{it}, \tag{6}$$

where the core and the control explanatory variables are defined as before for Equation (5), in Equation (6) instead of the market reform measure the institutional quality measure is used.

The first half of Table 8 compares CIS with the Central European and Baltic countries. The factors in favor of the Central European and Baltic countries were: (1) reforms were more advanced; (2) investment was higher; and (3) inflation was significantly lower. However, the positive impacts of these factors were more than offset by other factors in favor of the CIS, including recovery of lost output (initial, 1998, real GDP in the CIS was 59 percent of its level in 1990 as compared with 95 percent in the Central European and Baltic countries), and the terms-of-trade shocks were in favor of the CIS. The second half of Table 7 shows that despite weaker institutions, lower investment, and higher inflation in the CIS as compared with the Southeast Asian economies, recovery of lost output, slower population growth, and more favorable terms-of-trade gave the CIS an edge of 2.3 percentage points each year. The average score for the institutional quality in the Southeast Asian economies was 75 as compared with 64 for the CIS. Assuming hypothetically that the institutional scores of the CIS countries were raised to Southeast Asian economies' level, then the resulting change in the fitted value for growth using the parameter estimates would be a gain of another 1.2 percentage points per year. These results indicate that institutional reforms could play a major role in sustaining the recent rapid growth in the CIS, particularly when the temporary factors (recovery of lost output and terms of trade) behind the recent rapid growth disappear.

III. Conclusions and Policy Challenges

This paper investigates the main factors behind the rapid growth in transition economies. The central conclusion from the growth accounting exercise is that most of the improvement came from higher TFP, which averaged about 3 percentage points of GDP per year in 2001–06, close to the TFP gains in Western Europe in 1950–60. During the initial years of transition, the disorganization or chaos resulting from the removal of central controls and coordination produced

Table 8. Sources of Regional Differences in Growth, 1999–2006

Variables	CIS vs. Central Europe and Baltics: Impact on Growth (in percent)	CIS vs. Central Europe and Baltics: Explanation	CIS vs. Southeast Asia[1]: Impact on Growth (in percent)	CIS vs. Southeast Asia[1]: Explanation
Actual difference in growth	2.0	Growth was higher	2.4	Growth was higher
Predicted change (from model)	2.1	Predicted difference in growth	2.2	Predicted difference in growth
Recovery of lost output	2.9	Initial real GDP index lower	2.3	Initial real GDP index lower
Investment	–0.3	Investment lower	–0.6	Investment lower
Market reforms	–2.3	Market reforms weaker		Reforms not available for Southeast Asia
ICGR (institutional quality)		(ICGR and reforms correlated)	–1.2	Institutions weaker
Fiscal adjustment	0.6	Greater fiscal adjustment	0.0	No difference in fiscal balance
Inflation rate	–0.2	Inflation higher	–0.3	Inflation higher
Terms of trade	1.4	Favorable terms of trade	1.4	Favorable terms of trade
Population growth	0.0	Same population growth	0.6	Populations growth lower

Source: Author's estimates.

Note: ICRG = International Country Risk Guide.

[1]Southeast Asian countries include China, Hong Kong, India, Indonesia, Malaysia, the Philippines, Singapore, South Korea, Thailand, and Vietnam.

negative TFP growth rates as output fell, and a large part of the capital stock lay idle. Subsequently, as the economies achieved macroeconomic stability and market reforms, the reallocation of resources to more productive activities allowed the economies to generate rapid growth, especially in the CIS, with low rates of investment so that TFP growth rates increased.

Using panel data regression, the results suggest that the rapid per capita real GDP and TFP growth rates since 1999 is explained by the extent of how much transition economies have contracted in terms of real GDP in the 1990s and the degree of progress made in market reforms. Other factors, such as the terms of trade and macroeconomic stabilization, have also contributed to rapid growth. The results are robust to different panel econometric techniques, and different specifications. The main findings are as follows:

- Transition countries that experienced larger declines in output during the early 1990s tended to grow at much faster rates. On average, of the 8 percent annual average growth rate for transition economies in 2001–06, about 2 percentage points are attributable to the recovery of lost output.
- The growth impetus associated with macroeconomic stabilization and market reforms has been substantial because of their effect on the overall productivity. Had the market reform been deeper, its impact on TFP growth would have been correspondingly larger.
- The growth acceleration payoff to reforms in 2001–06 was enhanced by the favorable external environment (positive terms-of-trade shock and global technological innovation). These factors have accounted for about two percentage points of the annual growth in transition economies. The global environment alone in recent years explains 1 percentage point of the annual average growth.

A key question for prospective growth is whether the gains achieved thus far have already eliminated most of the inefficiencies of central planning—and will therefore soon fade away. As the transition countries approach the world technology frontier, thereby exhausting the opportunity for further TFP growth from this source, alternative channels to improve TFP growth will need to be sought. Further improvement in policy and institutions would need to play a role in this endeavor. Also, greater labor use and the recent trend of faster capital accumulation are expected to play a more important role in the medium-term growth.

APPENDIX I. SAMPLE, DATA DEFINITION, AND SOURCES

List of Countries

The set of countries covered in this paper was determined by the availability of key variables; small countries (with population less than one million) were also excluded. The global sample consists of the following 123 countries during 1983–2006:

Transition sample: Armenia, Azerbaijan, Belarus, Georgia, Kazakhstan, the Kyrgyz Republic, Moldova, Mongolia, Russia, Tajikistan, Ukraine, Uzbekistan, Estonia, Latvia, Lithuania, Czech Republic, Hungary, Poland, Slovak Republic, Slovenia, Albania, Bosnia and Herzegovina, Bulgaria, Croatia, Macedonia, and Romania.

Global sample: Armenia, Azerbaijan, Belarus, Georgia, Kazakhstan, the Kyrgyz Republic, Moldova, Mongolia, Russia, Tajikistan, Ukraine, Uzbekistan, Estonia, Latvia, Lithuania, Czech Republic, Hungary, Poland, Slovak Republic, Slovenia, Albania, Bosnia and Herzegovina, Bulgaria, Croatia, Macedonia, Romania, Australia, Austria, Canada, Cyprus, Denmark, Finland, France, Germany, Greece, Ireland, Island, Italy, Japan, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, United Kingdom, the United States, Bangladesh, Cambodia, China, Hong Kong, India, Indonesia, Korea, Malaysia, the Philippines, Singapore, Sri-Lanka, Thailand, Vietnam, Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominican Republic, Ecuador, El-Salvador, Guatemala, Guyana, Honduras, Mexico, Nicaragua, Panama, Paraguay, Peru, Trinidad and Tobago, Uruguay, Venezuela, Algeria, Egypt, Iran, Israel, Jordan, Lebanon, Morocco, Oman, Pakistan, Tunisia, Turkey, Yemen, Kuwait, Saudi Arabia, Syria, the UAE, Angola, Botswana, Burkina Faso, Cameroon, Central African Republic, Ethiopia, Ghana, Gabon, Guinea, Gambia, Guinea-Bissau, Ivory Coast, Kenya, Madagascar, Malawi, Mali, Mauritius, Mozambique, Namibia, Niger, Nigeria, Senegal, South Africa, Sudan, Togo, and Tanzania.

Sample Period

Annual and period average data were used. The transition sample is divided into four subperiods: 1991–94, 1995–98, 1999–2002, and 2003–06. The global sample is divided into seven subperiods: 1980–82, 1983–86, 1987–90, 1991–94, 1995–1998, 1999–2002, and 2003–06. The resulting information was unbalanced because of data limitations for some countries.

Definition and Sources of Data

The data sources used are the databases of the IMF, World Economic Outlook (WEO), United Nations Economic Commission for Europe, and the World Bank.

- *Per capita growth* (dependent variable): Per capita real GDP growth rate calculated from national currencies in constant prices. Source: WEO database.
- *Convergence* as measured by the initial income per capita in PPP-adjusted (U.S. dollars). Source: World Bank.
- *Investment*: Fixed capital formation as percent of GDP. Source: WEO database.
- *Recovery of lost output for transition sample*: An index of real GDP for all transition countries is constructed with real GDP for 1990 = 100. Source: Author's calculations based on the annual real GDP growth rates.
- *Recovery of lost output for global sample*: The following indices are used: (1) 0 if initial real GDP was greater than 95 percent of its value in 1990; (2) 0.5 if initial real GDP was between 70 and 95 percent; and (3) 1 if initial real GDP was less than 70 percent.
- *EBRD market reform index* (only for transition countries): The unweighted average of eight EBRD structural reform indicators—price liberalization, small-scale privatization, large-scale privatization, competition policy, trade liberalization, financial sector reform, governance and enterprise reforms, and infrastructure reform. The EBRD indicators range from 1 to 4.3, where 4.3 indicates that the country's structural characteristics are comparable to those prevailing on average in

market economies, and 1 represents conditions before reform in a centrally planned economy with dominant state ownership of the means of production. The reform indices are not perfect and their assessment is sometimes influenced by the observed macroeconomic performance, which raises the problem of possible endogeneity. Source: EBRD, Transition Reports (various years).

- Campos and Horváth (2006) market reform index is based on three categories of reform. The first captures internal liberalization reflecting the extent of price and wage liberalization. The second captures external liberalization efforts as reflected in the severity of trade barriers and capital controls. The third category captures the extent of privatization efforts. These three indices were constructed for 25 transition economies for all years between 1989 and 2000. Using Campos and Horváth's methodology this paper extended the market reform indices through 2006.
- The measure on institutional quality is taken from the ICRG, compiled by the private consultancy firm Political Risk Services. This data set covers 143 countries, from 1984 to the present. The composite index is an aggregation of various subcomponents that measure factors such as government stability, democratic accountability, law and order, quality of bureaucracy, and corruption in government. Source: Political Risk Services Group (2007), Syracuse University.
- *Ratio of government consumption to GDP*: Source: IMF, International Financial Statistics (IFS) and WEO database.
- *Macroeconomic stabilization* as measured by overall fiscal balance as a ratio of GDP, and the logarithm of the inflation rate. Source: IMF, WEO database.
- *External conditions*: Terms-of-trade shocks: percentage change in the terms-of-trade index (2000 = 100). Source: IMF, WEO database. World cycle period-specific shifts: Time dummy variables for 1987–90, 1991–94, 1995–1998, 1999–2002, and 2003–06. Source: Author's construction.
- *Instrumental variables*: (1) Commodity exports as percent of total exports: is the share of fuel, ores, metal, and agricultural raw exports in total exports, UNCTAD; (2) distance to Brussels is measured in 1,000 kilometers and refers to the distance from the *y* capital of the respective country to Brussels; (3) ethnic fractionalization is presented as the probability that two randomly selected individuals in a country are not from the same ethnic group (Alesina and others, 2003); (4) a dummy variable for civil conflict or war; and (5) a trend with values of 1 for 1991–94, 2 for 1995–98, 3 for 1999–2003, and 4 for 2004–06.

APPENDIX II. TEST OF WEAK INSTRUMENTS

This appendix is based on Stock, Wright, and Yogo (2001, 2002). They proposed an approach to make interference about weak instruments.

$$y = Y\beta + u, \tag{A.1}$$

$$Y = Z\Pi + v, \tag{A.2}$$

where y and Y are $T \times 1$ vectors of observations on endogenous variables, Z is $T \times K$ matrix of instruments, and u and v are $T \times 1$ vectors of disturbance terms. The errors $[u_t\ v_t]$' $(t = 1, \ldots, T)$ are assumed to be iid $N(0, \Sigma)$, where the elements of Σ are σ_u^2, and σ_{uv}^2, and σ_v^2 and let $\rho = \sigma_{uv}/(\sigma_u\sigma_v)$. The reduced equation (A.2) related the endogenous regressor to the instruments.

The concentration parameter, μ^2, is a unitless measure of strength of the instruments and is defined as:

$$\mu^2 = \Pi' Z' Z \Pi / \sigma_v^2. \tag{A.3}$$

A useful interpretation of μ^2 is in terms of F, the F statistics for testing the hypotheses $\Pi = 0$ in (A.2) (that is, the "first-stage F statistics"). Stock, Wright, and Yogo (2002) derive the following equations:

$$\mu(\hat{\beta}^{2SLS} - \beta) = (\sigma_\upsilon/\sigma_\nu)\frac{(Z_u + S_{uv}/\mu)}{1 + 2Z_v/\mu^2 + S_{uv}/\mu^2}, \tag{A.4}$$

where

$$Zu = (\Pi' Z' u)/(\sigma_u \sqrt{\Pi' Z' \Pi}),$$

$$Z_v = (\Pi' Z' v)/(\sigma_v X \sqrt{\Pi' Z' \Pi}),$$

$$S_{uv} = (v' P_z u)/\sigma_v \sigma_u), S_{vv} = (v' P_z v)/\sigma_v^2.$$

Under weak-instrument asymptotics, a threshold value μ^2/K is implied. If the actual value of μ^2/K exceeds this threshold, then the instruments are strong (for example, 2SLS relative, bias is <10 percent). Otherwise, the instruments are weak. The first-stage F statistics must be large, typically exceeding 10, for 2SLS inference to be reliable.

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The Optimal Level of Reserves for Low-Income Countries: Self-Insurance against External Shocks

REGIS BARNICHON*

This paper develops an analytical framework that helps to quantify the optimal level of international reserves for a small open economy with limited access to foreign capital and subject to natural disasters or terms-of-trade shocks. International reserves allow the country to relieve balance of payments pressures caused by external shocks and to avoid large fluctuations in imports. The paper calibrates the model to two regions—the Caribbean and the Sahel region in sub-Saharan Africa—and assesses the sensitivity of the results. The conclusion is that popular rules of thumb, such as maintaining reserves equivalent to three months of imports, only give imprecise benchmarks. [JEL F30, F31, F32]

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What is the optimal amount of international reserves for countries with limited access to foreign capital? While the recent buildup in international reserves in Asia spawned a renewed interest in the appropriate level of reserves for emerging market economies, less developed countries have largely been ignored by the literature. As a result, policymakers rely on personal judgment or rules of thumb such as maintaining reserves equivalent to three months of imports to evaluate a country's needs.

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This paper develops an analytical framework that helps to quantify the level of reserves that can be rationalized in terms of insurance against large external shocks, such as natural disasters or terms-of-trade shocks. By calibrating the model, the paper estimates the optimal amount of international reserves for two groups of countries subject to different natural disasters: the Caribbean, subject to hurricanes, and the Sahel region in sub-Saharan Africa, subject to drought.[1]

Although there are a number of reasons to accumulate international reserves, many low- to middle-income countries have weakly diversified economies that are very vulnerable to natural disasters or terms-of-trade shocks. Indeed, less developed countries often rely on international trade to import large quantities of goods of prime necessity (such as food) and on a single export sector to generate most of the foreign-exchange inflows. In addition, and unlike middle- to high-income countries, they do not have fast access to private foreign capital and must rely on international and bilateral donors to meet emergency financing needs. Their reliance on such flows, however, has considerable disadvantages. It can take time before donor resources are committed and disbursed, and there may be competition for donor resources from other countries with relief needs at the same time. In this context, international reserves can play a critical role by allowing rapid access to foreign exchange to avoid large imports fluctuations due to balance of payment constraints.

This paper presents a simple model that helps to quantify the level of reserves that can be rationalized in terms of insurance against large external shocks. The model looks at the intertemporal optimization problem of a small open economy that can hold costly foreign reserves to smooth import fluctuations in the face of large external disturbances. Because of the balance of payment constraint, a country can only buy imports if it receives enough foreign exchange inflow. By suddenly disrupting the normal inflow of foreign exchange, a natural disaster or a terms-of-trade shock may prevent a country from importing the desired level of foreign goods, resulting in a welfare loss. With an appropriate amount of international reserves, a country can minimize the negative impact of such shocks. Under a few assumptions, one can simplify the problem and derive a closed-form solution for the optimal reserves-to-import ratio that depends on the frequency and duration of shocks, the economic damage, the economy's characteristics, and the opportunity cost of holding reserves. Using data on natural disasters and terms-of-trade shocks since 1960, this paper then calibrates and numerically solves the fully fledged model to estimate the optimal amount of international reserves for the Caribbean and the Sahel countries.

The popular rule of thumb of maintaining reserves equivalent to three months of imports gives only an imprecise benchmark, as small changes in key parameters such as the shock's persistence, the size of the export sector,

[1]See Table A1 for a list of countries for each region.

or the degree of risk aversion can have large consequences on the optimal reserves level. Although an average Caribbean country needs only about a month and half of imports, an average Sahel country needs over four months. Hence, rules of thumb can only give an imprecise benchmark, and a careful study of each country's characteristics is necessary to evaluate its needs.

It is also important to stress that these estimates constitute only a lower bound on the appropriate level of reserves, as countries need to accumulate reserves to achieve other objectives beyond self-insurance against natural disasters and terms-of-trade shocks. Providing liquidity when needed or limiting exchange rate volatility (or maintaining a fixed peg) are perfectly good reasons that this framework will brush aside.[2]

The normative literature on the optimal level of international reserves goes back to the 1960s with Heller (1966), whose main insights were later formalized in a Baumol-Tobin inventory framework with fixed costs of adjusting reserves, and in which the stock of reserves is being depleted by a stochastic current account deficit (for a review, see, Frenkel and Jovanovic, 1981; Flood and Marion, 2002). More recently, the buildup in international reserves in Asia spawned renewed interest in the optimal level of reserves for emerging market countries prone to sudden-stops in capital inflows. Jeanne and Ranciere (2008), Aizenman and Lee (2005), Caballero and Panageas (2007), and Durdu, Mendoza, and Terrones (forthcoming) present models of optimal international reserves, in which countries aim to self-insure against sudden stops in capital inflows.[3] However, less developed countries have largely been ignored by the literature given their limited access to foreign private capital.[4] The framework presented here borrows from the vast literature on precautionary savings by modeling a low-income country as a representative agent with no access to the (international) capital market.[5] In the face of income uncertainty—represented by the occurrence of natural disasters—the agent's only option is to self-insure by managing a stock of riskless assets to buffer its consumption against adverse shocks. However, unlike standard models of precautionary savings, goods are not storable, and a precautionary savings motive emerges because of the balance of payment constraint as the country accumulates international reserves to avoid low consumption levels of imported goods. The model is closest to Jeanne and

[2]A country may also accumulate reserves if it pursues an export-led growth by artificially maintaining an undervalued exchange rate.

[3]See also Alfaro and Kanczuk (2009), who study the joint decision of holding reserves *and* sovereign debt, whereas most of the aforementioned literature tends to take the level of international debt as given.

[4]An exception is Aslam and Kim (2007), who study the optimal amount of precautionary savings in the face of volatile aid flows.

[5]For studies of precautionary savings, see Zeldes (1989); Caballero (1990); Kimball (1993); Carroll and Kimball (1996); and Hugget and Ospina (2001).

Ranciere (2008) but with two main differences. It is tailored to low-income countries with no access to capital markets, and it is explicitly dynamic; that is, a country can face more than one external shock over time.

I. Natural Disasters in the Caribbean and the Sahel

Developing countries are vulnerable to terms-of-trade shocks because they typically rely on a concentrated export sector to generate most of their foreign-exchange inflows. However, for many of these countries, the problem is exacerbated by geographical location. Figures 1 and 2 respectively display the number of people affected by natural disasters over 1963–2007 for the Caribbean and the Sahel.[6] Although the former is regularly hit by hurricanes, the latter suffers frequently from droughts. By disrupting the export sector and the normal inflow of foreign exchange, natural disasters can trigger balance of payments pressures in the same way that terms-of-trade shocks do. This section estimates the economic impact of natural disasters in each region, and calculates the average behavior of real output growth, real export growth, real import growth, and the change in the nominal exchange rate in a five-year event window centered on a shock.[7]

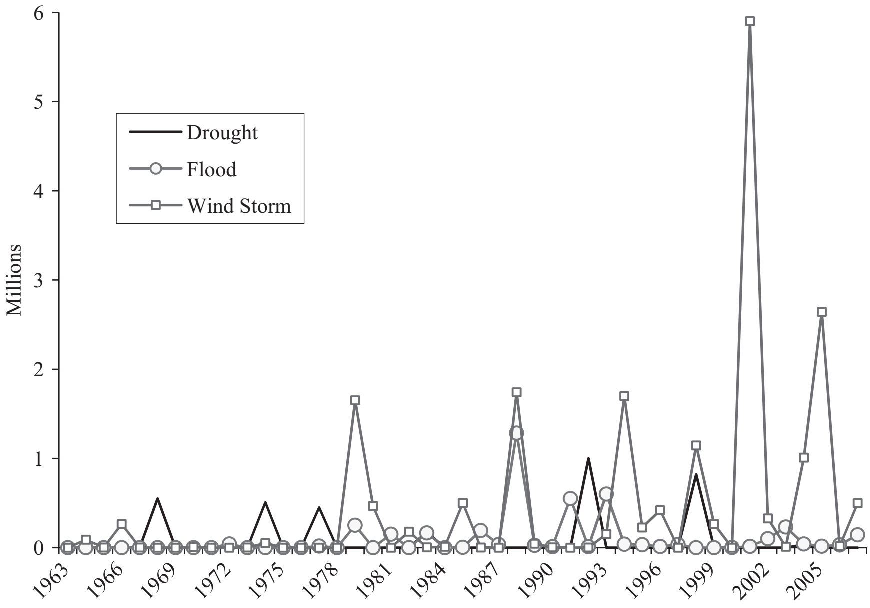
Hurricanes in the Caribbean

In September 2004, a Category 3 hurricane hit the island of Grenada and caused estimated damage of over US$800 million—or twice Grenada's GDP. Just as it required additional resources to finance relief, cleanup and emergency rehabilitations, the island experienced a dramatic decline in revenues and export earnings. Tourism and agriculture, the two major sources of foreign exchange earnings, were hit hard. Most tourism facilities could not reopen for the next six months, while the nutmeg crop, the principal export commodity, was largely destroyed. The government sought donor assistance, but despite over $150 million in pledges, only $12 million was available to address the immediate liquidity needs. Instead of focusing on recovery and reconstruction, the government was distracted by the need to

[6]Data on natural disasters are drawn from the Emergency Events Database (EM-DAT) published by the Centre for Research on the Epidemiology of Disasters (CRED) (www.em-dat.net).

[7]A good proxy for a hurricane's strength is the destruction of capital that it generates. Hurricanes are defined here as "major" when the estimated damage amounts (reported in EMDAT) represent more than 10 percent of a country's GDP. Droughts do not generate direct damage but rather hurt the population by disrupting production and/or triggering episodes of famine. Hence, a drought is classified here as "major" when either 10 percent of the population is affected *or* when at least five droughts occurred during the year. These thresholds allow for capturing most natural disasters with major consequences while ignoring smaller and more localized disasters that only had a minor impact on production and exports. With these threshold values, one obtains 30 observations (that is, disasters) for the Caribbean and 41 for the Sahel. All results are robust to alternative thresholds.

Figure 1. Total Number of People Affected by Type of Natural Disaster in the Caribbean



Note: Data on natural disasters are drawn from the Emergency Events Database (EM-DAT) published by the Centre for Research on the Epidemiology of Disasters (www.em-dat.net). People are considered "affected" when injured, homeless, displaced, evacuated, or requiring immediate assistance during a period of emergency.

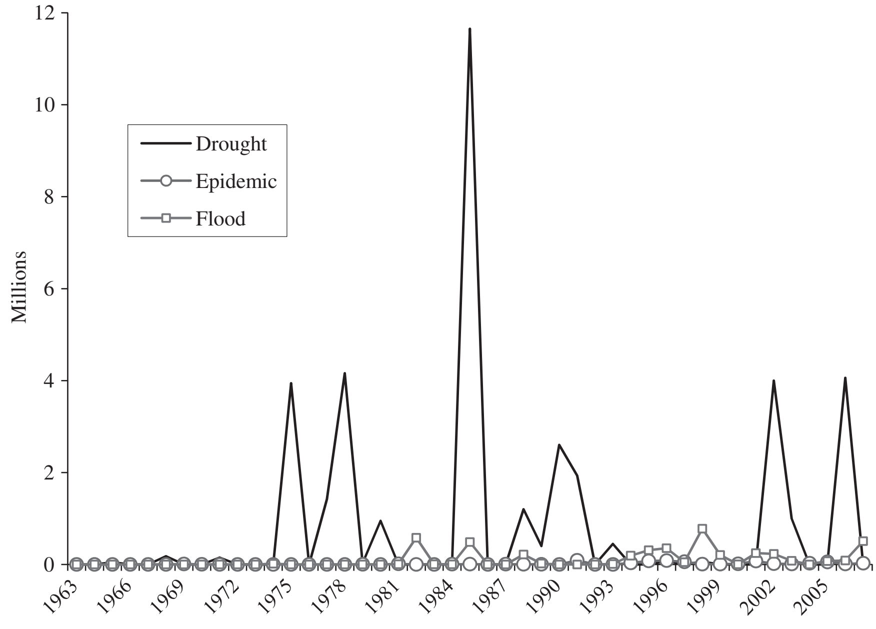
finance the emerging resource gap. This led to delays in the recovery and reconstruction periods.[8]

This episode illustrates the fate of many island nations from the Caribbean that suffer regularly from hurricanes, with enormous costs to the stock of capital and disruptions to the productive apparels. Figure 3 presents the average economic impact of major hurricanes in the Caribbean with one standard-deviation error band. On average, a major hurricane hits a Caribbean country every 20 years, that is, with a probability $\pi^{nd} = 0.95$ each year.[9] Output growth falls by 3 percentage points while exports fall by 5 percentage points. Despite the shortfall in foreign-exchange earnings, imports

[8]Despite the fact that many Caribbean islands are not, strictly speaking, low-income countries and theoretically have access to international capital markets, their high debt level limits their ability to access credit in the aftermath of a disaster. In addition, access to catastrophe insurance is limited due to the high transaction costs resulting from the relatively small amount of business island nations bring to these markets.

[9]Because natural disasters data are only available at an annual frequency, this data set cannot be used to estimate the disasters duration.

Figure 2. Total Number of People Affected by Type of Natural Disaster in the Sahel



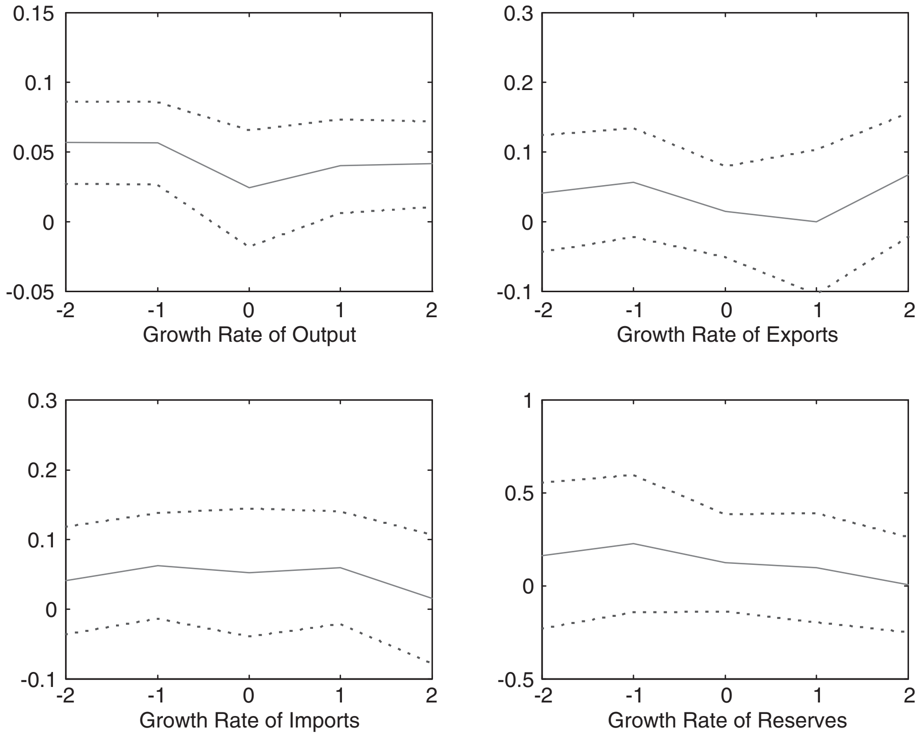
Note: Data on natural disasters are drawn from the Emergency Events Database (EM-DAT) published by the Centre for Research on the Epidemiology of Disasters (www.em-dat.net). People are considered "affected" when injured, homeless, displaced, evacuated, or requiring immediate assistance during a period of emergency.

do not decline; in fact, the country's import needs for reconstruction purposes are particularly large. As a result, import growth is relatively stable, declining by an average of only 1 percent, while the growth rate of reserves declines by an average of 10 percentage points on impact and by an additional 10 percentage points in the second year after the shock.

Droughts in the Sahel

Figure 4 presents the average economic impact of major droughts in the Sahel region with one standard-deviation error band. An average country from the Sahel faces one major drought every 12 years, that is, with a probability $\pi^{nd} = 0.92$ each year. Unlike hurricanes, droughts tend to develop over the course of several years. Although the behavior of real economic variables resembles qualitatively that of Caribbean countries, there are quantitative differences. Output growth drops only marginally by 0.3 percentage points on impact before rebounding the next year, but export growth drops by 8 percentage points. Imports remain roughly constant while the growth rate of reserves falls by 16 percentage points.

Figure 3. Impact of a Major Hurricane in the Sahel

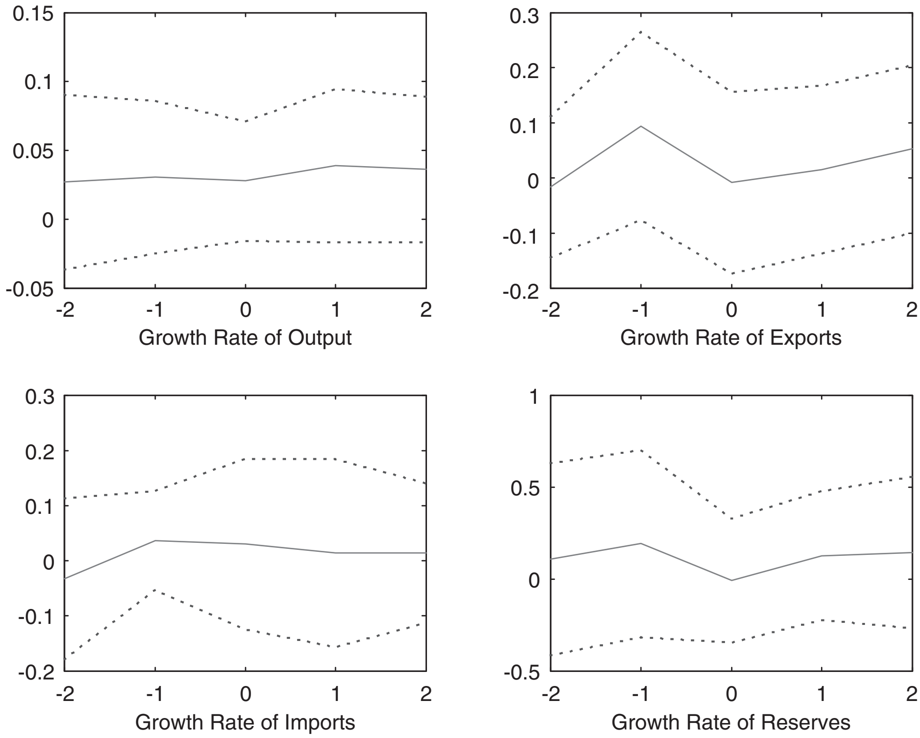


Note: The five-year event window is centered around a natural disaster occurring at time 0. A drought is considered "major" when either 10 percent of the population is affected or at least five droughts occurred during the year. Dashed lines represent the one standard-deviation error bands.

II. A Model of Optimal International Reserves

This section presents and calibrates a simple model of a small open economy that cannot access international capital markets but can hold costly international reserves to smooth consumption fluctuations in the face of large terms-of-trade shocks or large disturbances to exports. The optimization problem of this small open economy is analogous to the optimization problem of a single individual in the heterogenous-agents models of precautionary savings (for example, Hugget, 1993; Aiyagari, 1994). As in those models, Constant Relative Risk Aversion (CRRA) utility with incomplete markets implies that the marginal utility of consumption goes to zero from above, making the economy "extremely averse" to a savings plan that would leave it exposed to the risk of "very low" consumption at any date and state of nature. However, unlike standard models of precautionary savings, goods are not storable, and a precautionary savings motive emerges because of the balance of payment constraint. The model economy consumes

Figure 4. Impact of a Major Drought in the Sahel



Note: The five-year event window is centered around a natural disaster occurring at time 0. A drought is considered "major" when either 10 percent of the population is affected or at least five droughts occurred during the year. Dashed lines represent the one standard-deviation error bands.

two types of goods—domestically produced or imported—but without access to international capital markets; it can only buy imports if it receives enough foreign exchange inflow. More specifically, a country can import foreign goods by (1) exporting home goods, (2) borrowing or receiving grants from abroad, or (3) using foreign exchange reserves. By suddenly disrupting exports and the normal inflow of foreign exchange, a natural disaster or a terms-of-trade shock may prevent a country from importing the desired level of foreign goods, resulting in a welfare loss. By holding an appropriate amount of international reserves, a country can minimize the negative impact of such shocks.[10] However, this self-insurance comes at a price because of the opportunity cost of accumulating low-yield securities such as U.S. government bonds.

[10]Middle- to high-income countries can address the immediate liquidity needs by borrowing abroad but this is not the case for less developed countries with no immediate access to private foreign capital.

The Model

There are two countries, "Home and Foreign." Home is a small open economy consisting of a representative agent that consumes two types of goods: home goods c_H and foreign goods c_F. Both goods are not storable. At each period t, the Home consumer receives an endowment Y_t of home goods that can either be consumed or exported. The economy grows at the rate g so that $Y_t = (1+g)Y_0$ and in this nonstationary economy, all variables are expressed as a share of "normal" output (that is, the output level prevailing in a nondisaster state). In order to import foreign foods, Home must pay in foreign currency so that at each date t, it must satisfy the balance of payments constraint

$$c_{F,t} \leq \varepsilon_t c^*_{F,t} - ((1+g)R_{t+1} - R_t) + Tr_t,$$

where $c_{F,t}$ is the import-to-output ratio, $c^*_{F,t}$ is the export-to-output ratio, ε_t is the terms of trade, R_t is the reserves-to-output ratio (the amount of international reserves scaled by output) and Tr_t is a generic term for foreign transfers (private remittances or official grants) and loans as a share of output.[11]

Holding low-yield reserves presents an opportunity cost to Home modeled as a payment rR_t, payable in home goods. Indeed, Jeanne (2007) argues that instead of accumulating reserves in the form of, say, low-yield U.S. bonds, economies could receive a higher rate of return by investing in the domestic business sector or in the building of public infrastructure.[12] Hence, the aggregate resource constraint (rescaled with the output level) takes the form

$$c_{H,t} + \frac{1}{\varepsilon} c_{F,t} = y_t - \frac{1}{\varepsilon}(rR_t + ((1+g)R_{t+1} - R_t) - Tr_t),$$

with y_t denoting the country's endowment as a share of its income in "normal" times.

The representative agent seeks to maximize its expected utility by consuming home and foreign goods subject to the aggregate resource constraint and the balance of payment constraint. At date 0, Home's problem can be written

$$\max_{\{c_{H,t}, c_{F,t}\}} E_0 \sum_{t=0}^{\infty} \beta^t u(c_{H,t}, c_{F,t})$$

$$\text{s.t.} \begin{cases} c_{H,t} + \frac{1}{\varepsilon} c_{F,t} = y_t - \frac{1}{\varepsilon}(rR_t + (R_{t+1} - R_t) - Tr_t) \\ c_{F,t} \leq \varepsilon c^*_{F,t} - ((1+g)R_{t+1} - R_t) + Tr_t \\ R_{t+1} \geq 0 \end{cases}.$$

[11]Home takes ε_t as given as the low-income country is too small compared to the rest of the world to affect world markets and the terms of trade.

[12]Jeanne (2007) also measures the opportunity cost of reserves by using the spread between the interest rate on external debt and the return on reserves. However, it is difficult to apply this approach to low-income countries with no or little access to private capital market and whose external debt consists mostly of subsidized loans from international organizations or foreign governments.

To capture the occurrence of rare disasters and their impact on the economy, the country's endowments of home goods as well as the value of its exports follow a two-state Markov process with time-invariant transition probabilities. In a "normal" state, the representative agent receives an endowment Y^n and exports a fraction of output $c_F^{*n} = \delta$. However, with probability π^{nd}, a natural disaster hits the economy in a "normal" state and disrupts output production, exports capacities, and the terms of trade such that $Y^d = \eta_Y Y^n$, $c_F^{*d} = \eta_X c_F^{*n}$ and $\varepsilon^d = \eta_\varepsilon \varepsilon^n$ with η_Y, η_X, $\eta_\varepsilon < 1$. Once in a "disaster" state, the economy returns to its "normal" state with probability π^{dn} so that $1/\pi^{dn}$ is the expected duration of the disaster.

Because foreign donors typically decide unilaterally on the aid amount they provide to less developed countries, it is assumed that Foreign provides an exogenous and constant aid-to-GDP ratio Tr each period. Hence, in the aftermath of a disaster, Home can only cover its foreign exchange losses by using international reserves.[13] By denoting $\tilde{c}_{F,t} = c_{F,t} - Tr$ the imports that are only paid for with international reserves or the proceeds of exports, and $\mathscr{S}_{\{s=d\}}$ an indicator function equal to zero in a normal state and one in a disaster state, one can rewrite Home's problem at date t as follows:

$$V(R_t) = \max_{R_{t+1}} \left[u(c_{H,t}, \tilde{c}_{F,t}) + E_t \beta V(R_{t+1}) \right]$$

$$\text{s.t.} \begin{cases} c_{H,t} = (1-\delta) y_t - r R_t / \varepsilon_t \\ \tilde{c}_{F,t} \leq \varepsilon_t \delta_t y_t - (1+g) R_{t+1} + R_t \\ y_t = 1 - (1-\eta_Y) \mathscr{S}_{\{s_t = d\}} \\ R_{t+1} \geq 0 \end{cases}.$$

Home will choose its level of international reserves to satisfy the first-order condition $u'_{C_F,t} = \frac{\beta}{1+g} E_t \left(u'_{C_F,t+1} - \frac{r}{\varepsilon_{t+1}} u'_{C_H,t+1} \right)$. By accumulating one more unit of reserves in period t, Home gives up on consumption of foreign goods at t and of home goods at $t+1$ because of the opportunity cost of reserves, but it also enjoys a higher expected utility of foreign goods consumption at $t+1$. Higher income growth has the same effect as a higher discount rate as the representative agent would like to increase consumption and borrow in anticipation of higher future income.

In the steady-state of the "normal" state, Home reaches its optimal reserves to output ratio R^* and its consumptions of home and foreign goods are

$$\begin{cases} c_H = 1 - \delta - r R^* / \varepsilon \\ c_F = \varepsilon \delta^n - g R^* + Tr \end{cases}.$$

[13]This assumption remains valid for a time horizon of a few months after the shock. While the IMF, World Bank, and bilateral donors do provide emergency assistance for countries hit by natural disasters, the process can be lengthy, and the funds are usually not available until a few months after the shock.

An Approximated Closed-Form Solution for the Reserves-to-Imports Ratio

This subsection studies the problem analytically by considering a simpler case with the log-utility specification

$$u(c_{H,t}, c_{F,t}) = \theta \ln(c_{H,t}) + (1-\theta)\ln(c_{F,t}) \text{ with } \theta \in [0,1].$$

The country's first-order condition can be simplified by noting that $r \ll 1$, $g \ll 1$, $R \ll 1$ and $\pi^{nd} \ll 1$. Indeed, the opportunity cost of reserves will typically be smaller than 10 percent, and developing countries rarely have reserves-to-output ratio in excess of 10 percent. We also saw in Section I that disasters are extremely rare events occurring less than once every 10 years, so that π^{nd} is less than 1 percent. In that case, the Appendix shows that Home's first-order condition implies

$$R^* - (1+g)R' \approx \frac{\pi^{nd}}{\frac{1}{\delta\varepsilon}\left(\frac{1+g}{\beta} - (1-\pi^{nd})\right) + (1-\pi^{nd})\frac{r\theta/\varepsilon}{(1-\delta)(1-\theta)}} - \varepsilon\delta\eta_\varepsilon\eta_X\eta_Y,$$

with R^* the optimal reserves-to-output ratio and R' the level of reserves immediately after a disaster. Further, assuming that the agent uses almost all of its reserves at once so that $R' \ll R^*$, the resulting expression for the optimal reserves-to-import ratio is[14]

$$\frac{R^*}{c_F} \approx \varepsilon\delta\left[\frac{\pi^{nd}}{\left(\frac{1+g}{\beta} - (1-\pi^{nd})\right) + (1-\pi^{nd})\frac{\delta}{1-\delta}\frac{r\theta}{1-\theta}} - \eta_\varepsilon\eta_X\eta_Y\right] \times \frac{1}{\varepsilon\delta + Tr}. \tag{1}$$

Looking at Equation (1), one can draw a number of intuitive conclusions on the determinant of the optimal amount of reserves. A higher shock probability (that is, a higher π^{nd}) or a larger drop in the value of exports (that is, a lower η_Y, η_X or η_ε) raises the optimal reserves-to-import ratio. On the other hand, a higher opportunity cost of holding reserves (that is, a higher r), a higher discount rate (lower β) or higher growth rate (higher g) lowers the reserves-to-import ratio.[15] The share of imports covered by foreign grants or loans Tr influences the level of optimal level of reserves: a higher level of transfers (official loans or grants, and private remittances) in steady-state lowers the optimal reserves-to-import ratio as transfers or loans

[14]This assumption is clearly a restriction, but simulations show that with log-utility, it holds for country groups such as the Caribbean that face very disruptive but rare and short-lived disasters.

[15]As mentioned previously, higher income growth has the same effect on optimal reserves as a higher discount rate. Indeed, the model economy is equivalent to an economy with zero growth but with a discount rate $\tilde{\tilde{\beta}} = \frac{\beta}{1+g}$.

are not sensitive to natural disasters.[16] Finally, there is an inverse U-shape relationship between the size of the export sector δ and the optimal reserves-to-output ratio R^*. This is due to the interaction of two factors: the utility cost of accumulating reserves and the opportunity cost of maintaining a given level of reserves. Given the concavity of the utility function, higher foreign-exchange inflows (that is, a larger export sector) make reserves accumulation relatively easier, and the optimal reserves-to-import ratio increases (as captured by the first term $\varepsilon\delta$ on the right-hand side of equation (1)). However, above a certain level, the country exports and imports such a large share of its GDP that the steady state level of reserves gets large relative to GDP, and the opportunity cost of holding reserves becomes nonnegligible (that is, $\frac{\delta}{1-\delta}r$ becomes large in equation (1)).

III. Calibration and Numerical Solution

This section calibrates the model, calculates the optimal reserves-to-imports ratio, and conducts a sensitivity analysis exercise for each group of countries. From now on, the constant-elasticity-of-substitution (CES) utility function is used:

$$u(c_{H,t}, c_{F,t}) = \frac{\sigma}{\sigma - 1}\left[\theta^{\frac{1}{\gamma}}(c_{H,t} - \underline{c}_H)^{1-\frac{1}{\gamma}} + (1-\theta)^{\frac{1}{\gamma}}(c_{F,t} - \underline{c}_F)^{1-\frac{1}{\gamma}}\right]^{\frac{1-1/\sigma}{1-1/\gamma}}, \qquad (2)$$

with $\gamma > 0$ the elasticity of substitution between home and foreign goods, θ the preference for home goods, $1/\sigma > 0$ the coefficient of relative risk aversion, and c_H and c_F the subsistence consumption level of home and foreign goods. Note that when $\gamma = \sigma = 1$, this utility function reduces to the Cobb-Douglas utility used in the previous section. In addition, a Stone-Geary preference specification will be useful when calibrating the model to sub-Saharan African countries whose consumption is close to subsistence levels.[17]

In order to capture the urgency posed by some disasters, a monthly frequency is used for the calibration. Indeed, the main disruptions caused by a natural disaster such as a hurricane do not happen over the course of a year but over a few weeks or months. Hurricanes are sudden and short-lived events and the shortage of foreign exchange may materialize in the first weeks after the shock, not the next quarter or year. As a result, imports may drop to close to zero in the immediate aftermath of the shock with an arbitrarily large utility loss if a country does not hold any international reserves. A yearly frequency would smooth out the import loss and mask the utility loss given the concavity of the utility function.

[16] Again, a higher level of aid in the immediate aftermath of a disaster would lower the optimal reserve level even further. However, as argued in a previous footnote, a rapid response on a large scale is unlikely in the first months after the shock.

[17] This specification is consistent with the evidence from Ogaki and Zhang (2001) and Ogaki, Ostry, and Reinhart (1996) that the relative risk aversion coefficient is a decreasing function of wealth in poor countries.

Using the evidence from Section I on the impact of disasters and terms-of-trade shocks, the parameters of the model can be calibrated and the optimal level of international reserves estimated for countries from the Caribbean and the Sahel region. Table 1 presents the calibration parameters used for each country group. The monthly probability of shocks is set to match the estimates from Section I, and the output and export loss parameters are fixed to match the empirical ones. The size of the export sector is chosen to match the average exports-to-GDP ratio of the group, and the preference for foreign goods is set accordingly. The monthly discount factor β is set to 0.9966 and the coefficient of risk aversion to 5. Absent strong evidence regarding the elasticity of substitution between home and foreign goods, a value of 0.3 is used, as in Agenor, Bayraktar, and Aynaoui (2008) for the case of low-income countries. To calibrate the opportunity cost of holding reserves, one of Jeanne's (2007) suggestions is followed and the difference between higher-yielding domestic investment opportunities and the return on 10-year U.S. treasury bonds is considered. Caselli and Feyrer (2007) compute the return to capital in a sample of high- and low-income economies and document an average annual real return close to 7 percent in low-income countries. Because 10-year treasury bonds averaged a real rate of return of roughly 3.5 percent over 1963–2007, this approach leads to an opportunity cost of reserves of roughly 3.5 percent a year. Finally, unless otherwise noted, the subsistence levels of consumption are fixed to zero.

Finally, to numerically solve the Bellman equation $V(R_t)$, value function iteration on a grid is used for reserves holding spanning zero to five months of imports.[18]

Self-Insurance against Natural Disasters

Hurricanes in the Caribbean

In a "normal" state, the average Caribbean country exports and imports, respectively, 30 and 40 percent of its output. "Transfers" provide the remaining 10 percent of the financing. Consistent with Section I, a major hurricane hits every 25 years. Given that hurricanes are sudden events with little persistence and maximum disruption in impact, it is assumed here that a natural disaster brings exports to a full stop for some time. To estimate that time, the number of months with zero imports necessary to match the total exports loss of 10 percent identified in Section I is calculated. That way, one can estimate that a hurricane disrupts exports for an average of one month and a quarter while output drops by 36 percent.[19] Given that countries from

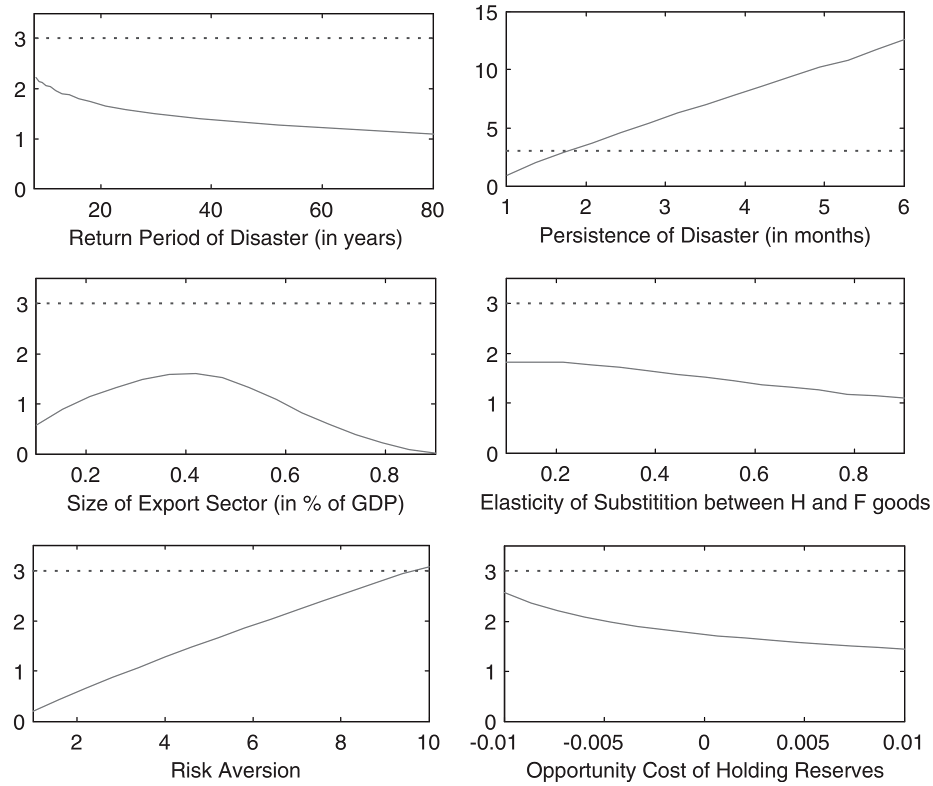
[18]For each country, I choose a grid size large enough (150 grid points) to ensure that the optimal reserves level (rounded at the second decimal) does not change with an increase in the grid's density. I start with an initial guess $V_0 = 0$ and stopping criterion $e = 10^{-5}$.

[19]A more realistic assumption would be to assume a gradual recovery phase starting one month after the shock. The present calibration exercise is mostly illustrative but could be easily extended to a richer setting.

Table 1. Calibration (*Monthly frequency*)

		Caribbean		Sahel	
		Hurricanes	Terms-of-trade shocks	Droughts	Terms-of-trade shocks
Probability of disaster (Return period)	π^{nd}	0.33% (25 years)	0.55% (17 years)	0.66% (12 years)	0.88% (10 years)
Persistence of disaster (Duration)	π^{dn}	83% (1.25 months)	8% (1 year)	16% (6 months)	8% (1 year)
Output loss	$1-\eta_Y$	36%	0	1%	0
Exports loss	$1-\eta_X$	100% (10% yearly)	0	16% (8% yearly)	0
Terms-of-trade loss	$1-\eta_\varepsilon$	0	10%	0	15%
Exports-to-GDP ratio	δ	0.4	0.4	0.2	0.2
Risk aversion	σ^{-1}	5	5	5	5
Elasticity of substitution		0.3	0.3	0.3	0.3
Growth rate	g	0.20%	0.20%	0%	0%
Subsistence level of imports (percent of initial GDP)	c_F	0%	0%	26%	26%
Transfers (percent of initial GDP)	Tr	10%	10%	10%	10%
Optimal reserves-to-imports ratio (in months of imports)	R^*/c_F	1.42	<0.01	1.93	2.43
Shocks combined		1.52		4.10	

Figure 5. Optimal Reserves-to-Import Ratio (in months of imports) in the Caribbean: Sensitivity Analysis



the Caribbean have fixed exchange-rate regimes in the majority of cases, the terms of trade is kept constant during disasters. Finally, GDP per capita in the Caribbean grew at an average rate of 2.5 percent per year over the past 40 years so the monthly growth rate g was set to 0.20.

The estimated optimal level of reserves covers 1.42 months of imports, that is, slightly more than the expected duration of the disaster. Figure 5 illustrates the sensitivity of the results to key parameters: the return period of disasters (that is, the disaster's probability), the disaster's persistence, the size of the export sector, the opportunity cost of holding reserves, the subsistence level of imports, and the coefficient of risk aversion. In each case, the analysis here starts from the baseline calibration and vary one parameter at a time to draw a number of conclusions. First, the optimal reserves-to-imports ratio increases with the shocks' probability (or return period) and the shock's persistence. Second, as we previously observed in the simpler case with log-utility, there is an inverse U-shape relationship between the size of the export sector and optimal reserves due to the interaction of two factors: the utility cost of accumulating reserves and the opportunity cost of maintaining a given level of reserves. Third, as the elasticity of substitution between domestically produced goods and imported goods increases from 0.1 to 0.9, the optimal reserves-to-import ratio decreases by about a month. Fourth, looking at a range of plausible values of risk aversion, optimal reserves reach three months of imports only for large degrees of risk aversion. Finally, given the wide range of opportunity costs spanned by the simulation (from –10 to +10 percent per year), the optimal R^* varies comparatively little. This happens because the import sector represents only 40 percent of GDP in the baseline calibration. The total cost of holding reserves rR^* is not large and has thus only a small impact on the level of reserves.[20]

Droughts in the Sahel

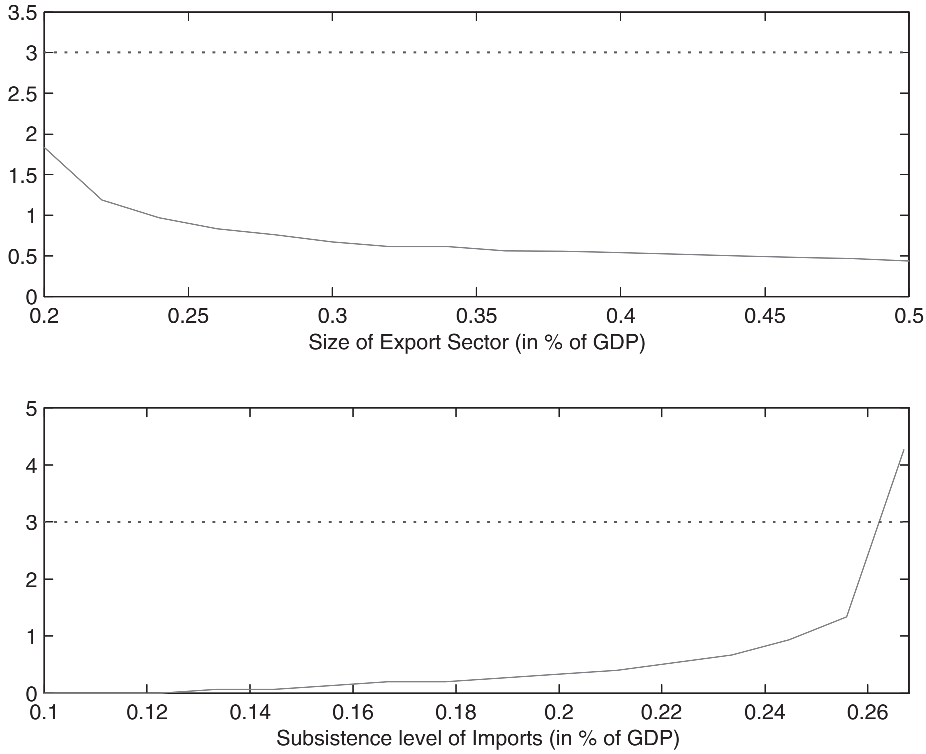
The average country from the Sahel region faces one major drought every 12 years. Droughts are more frequent than hurricanes and cause little physical damage but are also more persistent. Accordingly, it is assumed that annual output growth is unaffected while exports drop by 16 percent for six months, consistent with the 8 percent annual decline documented in Section I. The average Sahel country exports and imports, respectively, 20 and 30 percent of its output. Transfers provide the remaining 10 percent of the financing. Because most countries from the Sahel are part of the CFA Franc zone, the terms of trade is kept constant during disasters. Finally, to capture the situation in sub-Saharan Africa, where most of the population lives close to subsistence levels and where a small drop in consumption can have disastrous consequences, I postulate a nonzero subsistence level of imports.[21] Indeed, the 2008 riots in sub-Saharan Africa following the price increase of a number of food products shows that consumption in normal times is very close to subsistence levels. To calibrate $\underline{c}_F$, it is assumed that the 2008 riots were the result of consumption reaching subsistence levels. Given that the riots were triggered by an increase in food prices of 50 percent in one year and that the food basket represents roughly 40 percent of imports for an average Sahel country, the subsistence level of imports is set at 20 percent of "normal" imports, that is, 26 percent of "normal" GDP. However, the sensitivity of the results to a range of subsistence levels is also presented. Finally the GDP per capita growth rate in sub-Saharan Africa was zero or negative over the past 40 years so the monthly growth rate g is set to 0.

The estimated optimal reserves-to-import ratio is 1.93 months. Although the shock is less violent than with hurricanes, its duration makes it costly as it brings the population close to subsistence levels (that is, close to famine levels) for a long time. This provides a strong rational for holding

[20]In countries with a larger import sector (for example, 80 percent of GDP), simulations show that the opportunity cost of holding reserves plays a much more important role.

[21]Given the small effect of a drought on exports, the optimal reserves-to-import ratio is close to zero at 0.08 months if the subsistence level of imports is set to zero.

Figure 6. Optimal Reserves-to-Import Ratio (in months of imports) in the Sahel: Sensitivity Analysis



international reserves. When a drought occurs, reserves are used progressively to minimize the decline in imports over the expected duration of the drought. Figure 6 presents the sensitivity analysis and gives similar conclusions to the ones drawn for the Caribbean. Although the optimal reserves-to-import ratio depends on a country's characteristics, its level is always below three months except when parameters such as the subsistence level of imports, risk aversion, or the drought's persistence take very high values.

Self-Insurance against Terms-of-Trade Shocks

This subsection considers the impact of terms-of-trade shocks on the optimal level of international reserves. Caribbean economies tend to be less concentrated than in the Sahel, and the primary sector represents a smaller share of GDP. As a result, Caribbean countries are less sensitive to fluctuations in prices of raw materials, agricultural products, and staples. Using data on major terms-of-trade shocks since 1960, I find that while an average Caribbean country faces a 10 percent decline in its terms of trade every 17 years, an average Sahel country faces a 15 percent decline every 10 years.[22]

By calibrating the model to these transition probabilities and assuming that a terms-of-trade shock lasts for a year with no other economic impact on output and exports than the depreciation in ε, I find that the optimal reserves level represents less than 0.01 months of import for a Caribbean country but 2.43 months for a Sahel country.[23] Again, by bringing consumption of foreign goods close to subsistence levels for a long time, terms-of-trade shocks provide a strong rational for holding international reserves in the Sahel. In the Caribbean, however, a terms-of-trade shock lowers imports from 50 percent of GDP to slightly less than 45 percent, a small welfare loss given the concavity of the utility function.

Self-Insurance against Natural Disasters and Terms-of-Trade Shocks

This subsection estimates the optimal reserves-to-import ratio for each region by taking into account the possibility of natural disasters *and* terms-of-trade shocks. To do so, the two-state Markov process from Section III is generalized to three states. Home can either be in a ‘‘normal’’ state, facing a terms-of-trade shock, or facing a natural disaster, and the transition probabilities are the ones used previously.

The reserves target represents 1.52 months of import for a Caribbean country, only slightly more than that described earlier as self-insurance against natural disasters is the main motive for holding international reserves. For a country from the Sahel, however, the optimal reserves level stands at 4.10 months of imports, as droughts and terms-of-trade shocks are equally important disturbances.

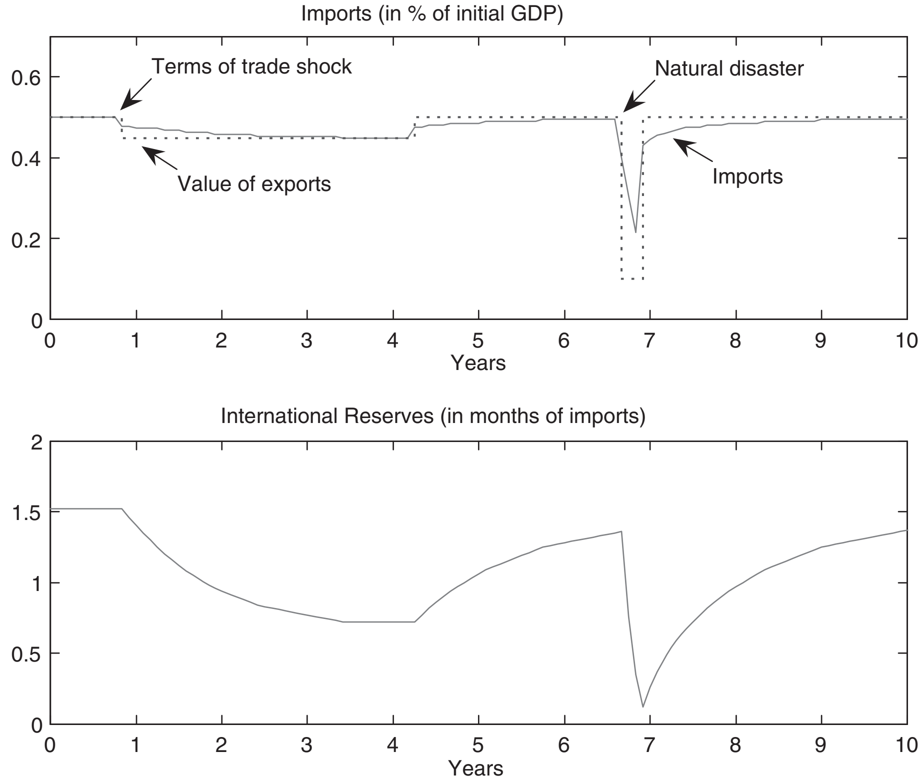
Figure 7 shows the impact of a terms-of-trade shock and a natural disaster on exports, imports, and international reserves in the Caribbean. Although Home initially keeps imports close to normal after the terms-of-trade shock, it progressively slows down the use of its reserves and reduces imports to avoid using too much of foreign exchange. Because the shock lasts much longer than expected (3.5 years instead of 1 year), Home stops using reserves after some time so as to keep enough reserves to respond to a hurricane. After the hurricane, exports drop by 50 percentage points, but imports decline by only 10 percentage points on impact thanks to the quick use of international reserves.

Figure 8 simulates the evolution of international reserves in a Sahel country. although Caribbean countries face the possibility of large, but rare,

[22]A worsening in the terms of trade is considered to be major when it is above 10 percent, based on data drawn from the IMF’s *World Economic Outlook*. Terms-of-trade shocks that coincide with natural disasters are ignored.

[23]These assumptions are consistent with estimates of the economic impact of a terms-of-trade shock in each region. The empirical analysis (similar to the one conducted in Section I) is available upon request.

Figure 7. Impact of External Shocks in the Caribbean

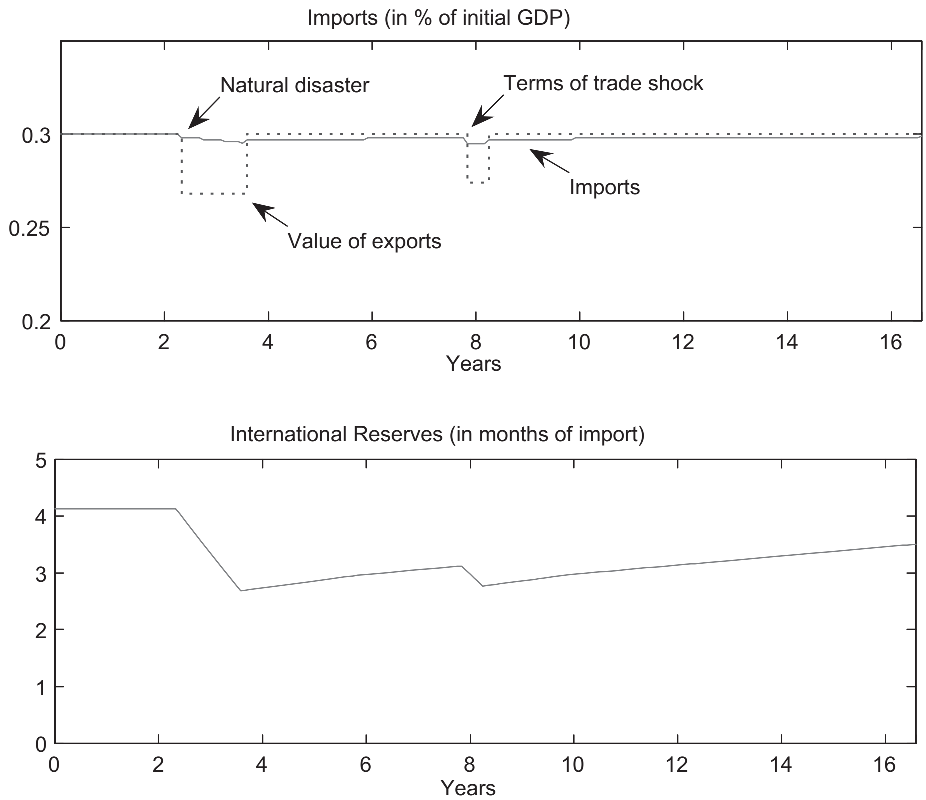


disruptive shocks, droughts and terms-of-trade shocks induce only mild, albeit frequent, declines in foreign exchange inflows. When hit by a terms-of-trade shock, Caribbean countries still face the possibility of very disruptive shocks and cannot afford to use too much reserves. This is not the case in the Sahel however, and in Figure 8 Home does not progressively slow down the use of its reserves. Similarly while Caribbean countries have to accumulate reserves at a fast pace, Sahel countries can smooth reserves accumulation.

Policy Implications

The main conclusion of this exercise is that the optimal reserves level is very sensitive to the parameters calibration. As a result, rules of thumb such as maintaining reserves equivalent to three months of import give only imprecise benchmarks. Although an average Caribbean country only needs one-and-a-half months of imports, an average Sahel country needs over four months. First, small parameter changes can have large consequences on the optimal reserves level. For example, depending on the size of the export

Figure 8. Impact of External Shocks in the Sahel



sector, the optimal reserves level in the Caribbean can take values between zero and two months of imports in the baseline calibration. Similarly, a shock's persistence of two months calls for roughly three months of imports in reserves but a persistence of three months already calls for five months.[24] Second, while some parameters play a critical role in a country, they can be almost irrelevant in another one. For example, the opportunity cost of holding reserves is negligible in a country with a small sector but it becomes determinant when exports represent a large share of GDP.

Finally, note that the average reserves level over time can be very different from the optimal level in steady-state, and one cannot evaluate a country's target by simply looking at its historical average. For example, the optimal reserves-to-import ratio is above four months for countries from the Sahel but the average reserves level is only at 3.25 months over the 16 years of the simulation.

[24]Note that the persistence of the shock can also be interpreted as the time taken by the international community to intervene and provide assistance that compensates for the loss in foreign exchange inflows.

IV. Conclusion

This paper develops an analytical framework that helps to quantify the level of reserves that can be rationalized in terms of insurance against large external shocks, such as natural disasters and terms-of-trade shocks. By calibrating the model, it is estimated that the optimal amount of international reserves for two groups of countries hit by different natural disasters: the Caribbean, hit by hurricanes, and the Sahel, hit by drought.

The calibration exercise shows that the optimal reserves level can be very sensitive to the parameters calibration, and the model needs to be carefully calibrated to evaluate each country's needs. As a result, rules of thumb such as maintaining reserves equivalent to three months of imports can only give imprecise benchmarks. Although an average Caribbean country only needs one-and-a-half months of imports, an average Sahel country needs over four months. Indeed, small changes in key parameters such as the size and persistence of shocks hitting a country, the importance of the export sector, or the degree of risk aversion, can have large consequences on the optimal reserves level.

An interesting extension would be to use a similar framework to evaluate the optimal size of sovereign wealth funds for economies relying mostly on primary commodities. Although the income provided by natural resources can provide large foreign exchange inflows, price volatility as well as uncertainty about the exact amount of natural resources available call for the accumulation of reserves to smooth price fluctuations and to provide an alternative source of revenue.

APPENDIX

Approximated Closed-Form Solution for the Reserves-to-Import Ratio

One starts by writing-up the agent's first-order condition:

$$\frac{1-\theta}{c_F}=\frac{\beta}{1+g}(1-\pi^{nd})\left(\frac{1-\theta}{c_F'^{n}}-\frac{r\theta}{c_H'^{n}}\right)+\beta\pi^{nd}\left(\frac{1-\theta}{c_F'^{d}}-\frac{r\theta}{c_H'^{d}}\right)$$

or

$$\frac{1-\theta}{\delta\varepsilon-gR}=\frac{\beta}{1+g}(1-\pi^{nd})\left(\frac{1-\theta}{\delta\varepsilon-gR}-\frac{r\theta/\varepsilon}{1-\delta-rR^{*}/\varepsilon}\right)$$
$$+\frac{\beta}{1+g}\pi^{nd}\left(\frac{1-\theta}{\varepsilon\delta\eta_{\varepsilon}\eta_{X}\eta_{Y}+R^{*}-(1+g)R'}-\frac{r\theta/\eta_{\varepsilon}\varepsilon}{(1-\delta\eta_{\delta})\eta_{Y}-rR^{*}/\eta_{\varepsilon}\varepsilon}\right),$$

with R^{*} the steady-state optimal reserves to output ratio and R' the level of reserves one period after the shock.

This expression can be simplified by noting that $r\ll 1$, $R\ll 1$, $g\ll 1$ and $\pi^{nd}\ll 1$. Indeed, the opportunity cost of reserves will typically be smaller than 10 percent, and developing countries rarely have a reserves-to-output ratio in excess of 10 percent. An annual growth rate in the order of 4 percent translates into a monthly g of only 0.33 percent. In addition, we saw in Section I that disasters are extremely rare events, occurring less than once every

Table A1. List of Countries in Each Group

Caribbean	Sahel
Antigua and Barbuda	Benin
The Bahamas	Burkina Faso
Barbados	Cape Verde
Belize	Côte d'Ivoire
Dominica	The Gambia
Dominican Republic	Guinea
Grenada	Guinea Bissau
Haiti	Mali
Honduras	Mauritania
Jamaica	Niger
St. Kitts and Nevis	Senegal
St. Lucia	Sierra Leone
St. Vincent and the Grenadines	Togo
Trinidad and Tobago	

10 years, so that π^{nd} is less than 1 percent. Hence, to a first order, one can approximate $1-\theta/\delta\varepsilon-gR$ with $1-\theta/\delta\varepsilon$, because gR is second order, and one can approximate $r\theta/1-\delta-rR^*$ with $r\theta/1-\delta$, because rR is second order, and given the parameters' calibration for η_Y and η_ε (see Table 1), the fourth term on the right-hand side is of second order ($\pi^{nd}r$ is second order) so that the first-order condition can be rewritten as

$$\frac{1-\theta}{\delta\varepsilon}\left(1-\frac{\beta(1-\pi^{nd})}{1+g}\right) \approx -\frac{\beta}{1+g}(1-\pi^{nd})\frac{r\theta/\varepsilon}{1-\delta} + \frac{\beta}{1+g}\pi^{nd}\frac{1-\theta}{\varepsilon\delta\eta_\varepsilon\eta_X\eta_Y + R^* - (1+g)R'}.$$

Rearranging, one has

$$\varepsilon\delta\eta_\varepsilon\eta_X\eta_Y + R^* - (1+g)R' \approx \frac{\frac{\beta}{1+g}\pi^{nd}(1-\theta)}{\frac{1-\theta}{\delta\varepsilon}\left(1-\frac{\beta(1-\pi^{nd})}{1+g}\right)+\frac{\beta}{1+g}(1-\pi^{nd})\frac{r\theta/\varepsilon}{1-\delta}},$$

so that

$$R^* - (1+g)R' \approx \frac{\pi^{nd}}{\frac{1}{\delta\varepsilon}\left(\frac{1+g}{\beta}-(1-\pi^{nd})\right)+(1-\pi^{nd})\frac{r\theta/\varepsilon}{(1-\delta)(1-\theta)}} - \varepsilon\delta\eta_\varepsilon\eta_X\eta_Y.$$

Assuming that the agent uses almost all of its reserves at once so that $R' \ll R^*$, one obtains the expression for the steady-state reserves-to-output ratio:

$$R^* \approx \frac{\pi^{nd}}{\frac{1}{\delta\varepsilon}\left(\frac{1+g}{\beta}-(1-\pi^{nd})\right)+(1-\pi^{nd})\frac{r\theta/\varepsilon}{(1-\delta)(1-\theta)}} - \varepsilon\delta\eta_\varepsilon\eta_X\eta_Y,$$

and the reserves-to-import ratio

$$\frac{R^*}{c_F} \approx \left[\frac{\pi^{nd}}{\left(\frac{1+g}{\beta}-(1-\pi^{nd})\right)+(1-\pi^{nd})\delta\frac{r\theta}{(1-\delta)(1-\theta)}} - \eta_\varepsilon\eta_X\eta_Y\right]\frac{1}{1+\frac{Tr}{\varepsilon\delta}}.$$

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Are Africa's Currency Unions Good for Trade?

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This paper explores and quantifies several aspects of the performance of Africa's currency unions. It benchmarks Africa's experience with that of the world using an augmented version of the gravity model and applying a comprehensive set of robustness checks. The empirical findings suggest that membership in a currency union should benefit Africa as much as it does the rest of the world. In addition, for both samples, we find evidence that (1) there is a significant currency union trade-generating effect; (2) currency unions are associated with trade creation and increased price comovements among member countries; and (3) the duration of currency union membership matters for trade: longer duration brings about greater benefits, and vice versa, however with some diminishing returns. [JEL F14, F15, F33]
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Newton's theory of universal gravitation relates the force of attraction between two objects to their respective masses and the distance between them. Tinbergen (1962) and Pöyhönen (1963) proposed that the same principle may be applied to explain bilateral trade flows where trade is estimated as an increasing function of the trading partners' incomes, and a

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decreasing function of the distance between them. Although simply specified, gravity models have performed well empirically, and are widely applied in empirical trade analysis.[1]

In his seminal work, Rose (2000) uses gravity models to empirically investigate the impact of currency arrangements on bilateral trade. He concludes that the effect of common currency on international trade is "large": two countries sharing a currency tend to trade roughly three times as much as they would otherwise. Rose and van Wincoop (2001), Frankel and Rose (2002), and Glick and Rose (2002) confirm this result, and show that it is robust to various specifications and estimation techniques. Frankel and Rose (2002) further find that the growth-enhancing benefits of currency unions are achieved through increased trade only and not through other channels (such as reduced inflation).

These results have generated immense interest and controversy in the academic community, and numerous studies have followed. Most of these studies point out methodological or data limitations in the earlier analyses and find a smaller impact of currency unions on trade (see, for example, Tenreyro, 2001; Nitsch, 2002; and Baxter and Kouparitsas, 2006). In general, these studies challenge the magnitude of the estimated effect of currency unions in the research conducted by Rose, but do not question the validity of its existence or the direction of the effect.[2]

From a theoretical standpoint, currency unions promote trade through various channels. For example, they reduce transaction costs between trading partners and facilitate exchange; create a larger market in which there are potential gains from economies of scale and production efficiency (assuming factor mobility and flexible wages); bring macroeconomic stability by signaling the central bank's commitment to reduce inflation; enhance the credibility of the monetary authority; and reduce uncertainty. Less uncertainty about prices can help allocate resources more efficiently in a region, while the absence of exchange rate risk in countries sharing a common currency encourages investment, facilitates capital mobility, and helps to generate trade.

Nonetheless, the potential gains from forming a currency union must be compared with the potential losses from membership. The potential losses depend on the extent to which countries in the union face asymmetric shocks, and whether countries in the region are sufficiently flexible to absorb or mitigate such shocks. Flexibility implies that factors are mobile and an

[1]Initially, these models were criticized for lacking a proper theoretical justification. Anderson (1979) and Bergstrand (1985) were the first formal attempts to address this criticism and derived the gravity equation theoretically. Deardorff (1998) provides a comprehensive overview of the gravity model and shows that a variety of theoretical models can be tied to it. Similarly, Feenstra, Markusen, and Rose (2001) argue that gravity models are consistent with several theoretical models of trade.

[2]See Baldwin (2005) for a comprehensive review of the literature and relevant issues on currency unions and trade.

appropriate system of transfers between countries is in place to act as a shock absorber and reduce costs of asymmetric shocks even if the shocks are large.

This paper extends the existing literature on how currency unions affect trade in three important directions. First, the main focus of the paper is Africa. We concentrate on Africa because it has a rich history of currency unions, although its integration in world trade has remained limited.[3] With ongoing debates over existing and proposed monetary unions in Africa, the benefits of forming a currency union are at the forefront of policy agendas and debates across the continent. We empirically investigate certain stylized facts of currency unions in Africa—including trade diversion and the impact of currency unions on price and output comovements, and on trade stability—and compare them with the world estimates to examine whether Africa's experience is different. Second, we investigate the time dimensional aspects of currency unions by estimating the effects of dissolution of a currency union on trade *and* the effect of the duration of currency union membership on bilateral trade flows. As the effect of joining a currency union on trade may evolve over time, it is important to quantify the changes (if any) in the trade-generating effect of currency unions. Third, we conduct a comprehensive set of robustness checks to address the methodological and econometric concerns often highlighted in estimating gravity models.

Our empirical findings suggest that African countries stand to benefit at least as much from currency union membership as other countries in the world; therefore, currency union benefits are not region-specific. Specifically, our results for both the world and Africa show that (1) countries belonging to a currency union trade, on average, about one-and-a-half times more with each other than with other comparable countries that do not share a currency; (2) currency unions are associated with trade creation and increased comovement of prices, but not with the comovement of output among members; (3) the effect of currency unions on trade stability is ambiguous; and (4) the duration of sharing a common currency matters significantly for bilateral trade: the longer a country participates in a currency union, the greater the benefit is, albeit with some diminishing returns. Although the trade-stimulating effect of currency unions holds significantly for Africa, given the continent's small trade base and the protracted period necessary following the establishment of a currency union to increase trade, our analysis suggests that currency union membership may not be a primary tool to achieve high levels of trade in the region, especially in the short run.

I. Literature Review

Baldwin's (2005) survey of the evidence on how currency union has an impact on trade correctly places Andrew Rose's research at the center of this literature. Rose and his coauthors find that membership in a currency union

[3]Figure A1 lists African regional economic integration arrangements. See Masson and Pattillo (2004) for a detailed discussion of currency unions in Africa.

promotes trade substantially. In his seminal work, Rose (2000) shows that trade between countries belonging to the same currency union is about three times larger than trade between comparable countries that do not share a currency. His study also shows that the effect of lower exchange rate volatility is positive, although the coefficient is much smaller than that associated with currency unions. In fact, the increase in trade associated with currency unions is much larger in order of magnitude than that associated with the complete elimination of exchange rate volatility.

Rose presents several arguments for his results, such as a common currency can induce financial integration that has consequences for trade. He also argues that by entering a currency union, a government signals its commitment to long-term integration, thereby promoting trade.[4] Rose and van Wincoop (2001) investigate further the trade-generating effect of currency union membership by using Anderson and van Wincoop's (2003) structural model to address country-specific idiosyncrasies. This approach, which was applied only to countries with complete bilateral data, somewhat reduces the effect of currency unions on trade to about two-and-a-half times. Glick and Rose (2002) explore the time-series dimension of the data, and introduce country-pair specific intercepts in the gravity model. They find a relatively smaller, but still large, impact of currency unions on trade: joining a currency union almost doubles bilateral trade. Although Rose's subsequent research and other studies on the subject have established different magnitudes of the estimated currency union effect, there seems to be general agreement in the literature on the existence of a common-currency effect, which, as suggested by Frankel (2005), "is probably substantially smaller than a tripling."[5]

A growing body of evidence describes the potential channels through which currency unions may affect trade. For example, Alesina and Barro (2002) show that besides country size and the volume of trade exchange with a potential anchor, the comovement of prices, outputs, inflation, and the volatility of inflation matter for currency unions. Thus, countries that stand to gain the most from joining currency unions are those having the largest comovements of outputs and prices with a potential anchor, and those with a history of high and volatile inflation. Tenreyro (2001) tests the predictions of Alesina and Barro (2002) using a probit model, and finds that countries with a higher comovement of prices have a higher propensity to form a currency union, but that the comovement of output has no effect on a country's decision. Alesina and Tenreyro (2002) and Tenreyro and Barro (2003) analyze the impact of currency unions on the comovement of output and prices between trading partners and investigate how comovements of outputs

[4]However, he acknowledges that the effect may be smaller for modern industrial countries; most currency unions in Rose (2000) comprise small or poor countries or both.

[5]Rose (2004) performs a "meta-analysis" of the currency union effect by combining estimates from 34 other studies and estimates a range of 30 to 90 percent for the currency union effect.

and prices would respond to the formation of a currency union. They find that sharing a currency enhances trade and increases price comovements, but decreases the comovement of shocks to real gross domestic product (GDP) (that is, increases specialization).

Keeping in view the potential and estimated impact of currency unions on trade, Africa presents an interesting case to assess the relative impact of currency unions and free trade agreements (FTAs) on intraregional and international trade. This is because Africa has a rich history of currency unions and preferential regional trade agreements, but its participation in world trade remains limited. Various explanations have been proposed for Africa's marginalization in global trade activity, including slow economic growth, unfavorable geographical and exogenous factors, poor infrastructure, ill-planned trade policies, weak governance and institutions, barriers to intraregional trade, and constraints on factor mobility. Further, the substantial savings on transaction costs that accrue from a monetary union and imply an increase in trade benefits may be limited in Africa because of lower diversification and a heavy dependence on primary commodities.[6] The loss of nominal exchange rate flexibility makes real adjustments to asymmetric shocks more difficult, especially in view of the poor systems of fiscal transfers and the limited development of the banking and financial sectors in Africa.

In this context, Debrun, Masson, and Pattillo (2005) show that gains from adopting a common currency depend, among other factors, on the correlation of terms-of-trade shocks. This, in turn, is connected to the countries' dependence on primary commodities and their prices. They also show that the existence of interest groups affects incentives to join a currency union or accept a new member in a multilateral union. This effect is noteworthy as it implies that differences in government spending propensities may be more important than asymmetric shocks for the benefits/losses arising from joining a currency union.

Coe and Hoffmaister (1999) analyze North-South trade and show that, on average, Africa trades more with the rest of the world (ROW) than other developing countries. However, Subramanian and Tamirisa (2003) stress the importance of distinguishing between Anglophone and Francophone countries while assessing the integration of African countries in global markets. Using Glick and Rose's (2002) specification, Masson and Pattillo (2004) examine the impact of currency unions on trade in Africa. Their estimated effect of currency unions on African bilateral trade with the ROW is almost the same as for the world: currency unions increase trade threefold in both Africa and the world.[7]

[6]See Collier (1995), Rodrik (1998), Yeats (1998), Collier and Gunning (1999), Limão and Venables (2001), and Subramanian and Tamirisa (2003) for a detailed discussion on this issue.

[7]However, the currency union variable used by Masson and Pattillo (2004) uses the FTA definition from Glick and Rose (2002), which does not distinguish between FTA and currency union effects. We overcome this limitation by constructing separate variables for FTA and currency unions, and identify their impacts separately.

This paper contributes to the existing literature by quantifying a series of "stylized facts" relating to the trade and currency union nexus in Africa, and investigating whether Africa's experience is different from that of the rest of the world. In particular, we examine the trade-generating impact of currency union membership, the impact on trade creation, and the effect of currency unions on trade stability and on the comovement of prices and outputs. Importantly, we also investigate whether the duration of currency union membership has any significant effect on trade. To explore these issues, we apply an augmented gravity model of trade that includes variables for FTAs, years of currency union membership, and trade diversion, and use an extended data set that covers more countries and years than previous empirical work on the subject.

II. Methodology

Analytical Framework

Traditionally, gravity models represent trade between two economies as a function of their respective economic masses and obstacles to trade such as the distance between them. However, this basic model has been extended in recent years to incorporate a variety of other factors that may hinder or promote trade, for example, common language, historical ties, common border, geographical location, and so forth.

In line with recent literature, we begin by investigating the effect of currency unions on trade by defining the following augmented gravity model:

$$\log(X_{ij}) = \beta_0 + \sum_{k=1}^{N} \beta_k Z_k + \gamma CU_{ij} + \nu_{ij}, \tag{1}$$

where i and j denote the exporting and importing countries, respectively; X_{ij} denotes the value of bilateral trade between i and j; CU_{ij} is a binary variable that is unity if i and j share the same currency;[8] γ is the estimate of currency union's trade-generating effect; and Z_k is a vector consisting of other variables that includes (log of) product of real GDP, real GDP per capita, land areas of the trading partners, and the distance between them, and dummy variables that are equal to 1 if the countries share an FTA, historical ties, language, or a border; are part of the same nation; or were colonies of the same colonizer after the year 1945, and 0 otherwise.

[8]The definition of "currency union," following Glick and Rose (2002), implies that money is interchangeable between the two countries at a 1:1 par for an extended period of time, so that there is no need to convert prices when trading between a pair of countries. Under this definition, hard fixes are not identified as currency unions. Further, the definition of currency union is transitive: if country pairs X, Y and X, Z are in a currency union, then Y and Z are in a currency union.

Next, we investigate the possibility that the stimulus to trade among members of a currency union comes at the expense of trade diversion with nonmembers. To do this, we follow Frankel and Rose (2002) and define a dummy variable that is unity if the trading partners are not in the same currency union but (at least) one is in a currency union with another country. A negative (and significant) coefficient of this variable would indicate the existence of potentially harmful trade diversion, and could be interpreted as implying that increased trade among members of a union comes at the expense of reduced trade with nonmembers.

Then, to investigate the impact of currency unions on comovements of output and prices, we construct the variables measuring comovement of prices and output, as in Alesina and Tenreyro (2002), and use them as dependent variables in Equation (1). Specifically, the price comovement between countries is determined by estimating a second-order autoregressive equation of annual price data for every pair of countries with more than 20 observations:

$$\ln\left(\frac{P_{it}}{P_{jt}}\right) = b_0 + b_1 \ln\left(\frac{P_{i,t-1}}{P_{j,t-1}}\right) + b_2 \ln\left(\frac{P_{i,t-2}}{P_{j,t-2}}\right) + \varepsilon_{tij}. \tag{2}$$

Estimated residuals from Equation (2) are then used to obtain a measure of comovement of prices with higher VP_{ij} representing greater synchronization of prices between countries i and j:

$$VP_{ij} = -\sqrt{\frac{1}{T-3}\sum_{t=1}^{T}\hat{\varepsilon}_{tij}^2}. \tag{3}$$

In a similar fashion, we construct a measure for the comovement of output, with $\hat{u}$ denoting estimated residuals from the following autoregressive process:

$$\ln\left(\frac{Y_{it}}{Y_{jt}}\right) = b_0 + b_1 \ln\left(\frac{Y_{i,t-1}}{Y_{j,t-1}}\right) + b_2 \ln\left(\frac{Y_{i,t-2}}{Y_{j,t-2}}\right) + u_{tij}, \tag{4}$$

and

$$VY_{ij} = -\sqrt{\frac{1}{T-3}\sum_{t=1}^{T}\hat{u}_{tij}^2}. \tag{5}$$

Next, to assess the currency union impact on trade stability, we follow Rose (2005) and estimate an equation similar to the gravity equation, with the coefficient of variation of log of real trade as the dependent variable of Equation (1). We calculate values for the dependent variable for the periods 1950–76 and 1976–2003, so we have two observations per pair. In addition, as a robustness check of the dependent variable, we use (1) the maximum absolute value (during the 27-year sample period) of the difference between the log of real trade and the sample average of trade of every country, scaled by the sample average; (2) the mean absolute value of the difference between exports and their sample average of every country (scaled by average exports); and (3) the standard deviation of the residual from a conventional gravity equation of exports in levels. All the explanatory variables are averaged over the corresponding time periods.

Finally, to investigate the effects of membership duration, we construct another variable of interest—the number of years that a given trading partner has shared a common currency. We modify Equation (1) to include this variable as follows:[9]

$$\log(X_{ij}) = \beta_0 + \sum_{k=1}^{N} \beta_k Z_k + \upsilon(yearsCU)_{ij} + \mu_{ij}. \tag{6}$$

Estimation Issues

While estimating trade flows using the gravity model, several relevant methodological issues need to be discussed. These issues are derived from various critiques of the estimation of the gravity equation and include the three "classic gravity model mistakes" pointed out by Baldwin (2005), as well as the critique on the correct functional form of the gravity equation pointed out by Santos Silva and Tenreyro (2006).[10] To the extent that these critiques relate to the analysis presented in this paper, we discuss our attempts to address them through robustness checks of the estimated results.

First, we begin with the issue of the omitted variables bias stemming from the correlation of any protrade omitted variables with the currency union dummy. This has been labeled as the "gold medal mistake" in Baldwin's (2005) critique. Research following Rose (2000) attempts to control for this bias by introducing country-specific idiosyncrasies in the model, both in the context of cross-section and in panel estimations. In cross-section analysis, country fixed effects (using country-specific dummy variables) can be introduced to account for Anderson and van Wincoop's (2003) "multilateral resistance" factor, according to which trade between two countries does not only depend on the characteristics of the countries but also on the barriers between them and the ROW.[11] However, given that there is a time-series element to the potential bias that is not eliminated with this procedure,

[9]Because the sample period of our data set begins in 1948, we ignore years spent in a currency union before 1948.

[10]For details on the "classic gravity model mistakes," see Baldwin (2005) and Frankel (2005).

[11]Anderson and van Wincoop (2003) use the national price indices (P_i and P_j) to account for "multilateral resistance" between countries i and j, which can be estimated using an iterative process. However, because the estimation process is complex, they propose an alternative method that is preferable for empirical work: namely, estimating implicit prices by fixed effects, that is, by including country-specific dummy variables.

we employ a panel data fixed effects procedure (the fixed effects "within" estimator) that adds country-pair specific effects to the equation, and thus exploits the time-series dimension of the data around country-pair averages.

Second, as forming a currency union (or continuing to stay in a currency union) may also be an endogenous choice, some of the large trade-creating effects of currency union may actually be a reflection of reverse causality. The use of instrumental variables could be a solution to the potential endogeneity problem. However, an appropriate instrument for a currency union is hard to find, which is further complicated by the fact that currency union membership is proxied by a dummy variable. Nevertheless, attempts by Alesina and Tenreyro (2002) to address the endogeneity problem using an instrumental variable based on client-anchor relationship have shown that the effect of currency union on trade remains high even after accounting for this potential endogeneity.[12] In addition, Rose and van Wincoop (2001) argue that "reverse causality also does not explain away the findings; there is little evidence in the political science literature that countries join currency unions to increase trade, and instrumental variables only increase the impact of currency unions on trade." This political dimension is particularly important for the case of Africa. Masson and Pattillo (2004) underscore the political aspect of the decision to form or participate in a currency union and argue that the experience of Africa shows that political objectives are important to the formation of monetary unions.[13] Hence, in our analysis we choose to treat currency unions as an exogenous variable with respect to trade.

Finally, there are several issues relating to model misspecification. These include (1) the aggregation of exports and imports as the dependent variable (the "silver medal mistake" of Baldwin); (2) inappropriate deflation of nominal trade values by the U.S. aggregate price index (the "bronze medal mistake" of Baldwin); (3) possible nonlinear effects entering the gravity equation; and (4) the treatment of zero-trade observations in the sample.[14] We attempt to address all these issues in our robustness checks.

On the aggregation issue, some critics argue that although theory supports the use of bilateral exports as the dependent variable, the use of bilateral trade as the dependent variable without properly aggregating imports and exports can seriously bias the results. We address this critique by taking the dependent variable as the sum of the logarithms of exports and imports in addition to the logarithm of the sums, and re-estimating

[12]However, the instrument applied by Alesina and Tenreyro (2002) is not designed for multilateral currency unions.

[13]The CFA franc zone and the Common Monetary Area were formed, in large part, due to the political self-interest of the major powers (France in the former case, and South Africa in the latter).

[14]The issue of nonlinearities is also discussed in Baldwin (2005). See Frankel (2005) for a justification of using the "pooled" export-import specification. The treatment of zero-trade observations in the estimation is discussed in detail in Santos Silva and Tenreyro (2006).

Equation (1). To account for the potential bias arising from inappropriate deflation by the aggregate U.S. price index, we add time dummies. This procedure corrects for global trends in inflation rates, as every bilateral trade flow is divided by the same price index adjusted for time effects. To address the possibility of nonlinear effects operating in gravity equation estimations (for example, due to sample nonhomogeneity), we add quadratic terms for both output and output per capita as in Glick and Rose (2002).

The issue of zero-trade observations arises because many observations in bilateral trade data sets appear as zeros either because some pairs of countries did not trade, or because of rounding errors and missing observations. Using the log-linear form of the gravity equation as in Equation (1) implies including only those observations for which the dependent variable is positive. Given that the value of trade flows between some pairs of countries—typically pairs of small countries—tends to be zero, this may lead to a sample selection problem. The truncation at zero may result in inconsistent estimators when ordinary least squares (OLS) are used.[15] We check the sensitivity of our results to the inclusion of zero-trade observations mainly in two ways. First, we apply the Tobit estimation method to account for the censored nature of the dependent variable of the model. Second, we apply the pseudomaximum likelihood (PML) approach applied by Santos Silva and Tenreyro (2006), which takes the real value of trade as the dependent variable, and includes zero observations. An additional advantage of using the PML approach is that it may have a superior functional form than the log-linear gravity model. This is because, as noted by Santos Silva and Tenreyro (2006), Jensen's inequality can have important implications for log-linear models in the presence of heteroscedasticity: if the error term is heteroscedastic with the variance depending on the regressors, then the parameters estimated by OLS can be severely biased.[16] Also, as commonly done in empirical gravity model literature, we avoid the truncation of observations by adding a positive constant to all trade observations and taking the log, that is, we use $\log(\text{constant} + X_{ij})$ as the dependent variable.

Data

The data set used in this paper is an extended version of Glick and Rose's data set. It includes 217 countries and political units over the time period

[15]Greene (1981) shows that the size of the "truncation bias" when the variables are distributed normally is inversely proportional to the "proportion of nonlimit observations in the sample," but this bias decreases when the fit of the model improves or the regressors have a skewed distribution.

[16]Jensen's inequality implies that even if the expected value of the error term obtained from Equation (1) is 0, $E[\log X_{ij}|Z_{ij}]$ is not essentially equivalent to $\exp(E[X_{ij}|Z_{ij}])$. The possible bias in the presence of heteroskedasticity can be mitigated with the use of heteroscedasticity-robust standard errors in the estimations.

1948–2003, which constitute the world sample in our analysis.[17] The Africa sample is a subset of the world sample, consisting of the bilateral trading patterns of 49 African countries. The intra-Africa and African trade with the ROW (Africa-ROW) samples include those pairs of countries where both trading partners are in Africa, and where only one partner is in Africa, respectively.

The data for the paper have been compiled from various sources. The annual bilateral trade observations are obtained from the IMF's *Direction of Trade Statistics* and are expressed in real U.S. dollars using the U.S. consumer price index for all urban consumers; GDP and prices (purchasing power parity of GDP) data have been taken from University of Pennsylvania's World Tables 6.1 and the World Bank's *World Development Report 2005*; the terms-of-trade data are from the IMF's *World Economic Outlook* 2005; information on colonial past, distance, and language has been compiled from the Central Intelligence Agency (CIA) *World Factbook 2004*; data on currency unions for the period 1998–2003 are from the IMF's *Annual Report on Exchange Arrangements and Exchange Restrictions*; and data on FTAs and currency unions are from Glick and Rose (2002). Table A1 provides details on data construction and sources. Tables A2–A4 present the countries, currency unions, and FTAs used in the study as well as the summary statistics for the variables of interest.

III. Results

The Trade-Generating Effect

We begin by investigating the general aspects of bilateral trade in the world and in Africa by using the extended data set and applying the gravity model to establish some benchmark results. In addition to the FTA dummy used by Glick and Rose (2002), we construct another FTA dummy variable that addresses an important limitation of Rose and Glick's data set because it takes into account the FTAs with Africa, and uses it as an alternate variable. In the next step, we estimate the characteristics of bilateral trade in Africa and trade creation, and address robustness issues.

Benchmark results

Table 1 shows the benchmark results of estimating Equation (1) using the extended sample. Following the literature, we use simple pooled OLS estimation (columns 1–3) and then employ the panel fixed effects estimation technique to address the potential omitted-variable bias by introducing the

[17] Political units include overseas territories, parts of kingdoms, possessions, self-governing territories in free association with another country, unincorporated territories, and crown dependencies.

Table 1. Benchmark Results

Sample:	World	World	Africa	World	World	Africa	World	World	Africa
Estimation:	OLS	OLS	OLS	Fixed effects	Fixed effects	Fixed effects	PML	PML	PML
Specification:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Currency union	1.00***	0.84***	0.99***	0.55***	0.60***	0.54**	0.15**	0.15**	0.65***
	(8.55)	(7.75)	(6.57)	(14.83)	(16.15)	(7.91)	(2.49)	(2.51)	(5.05)
Log distance	−1.08***	−1.05***	−1.06***						
	(52.68)	(52.08)	(19.77)						
Log product real GDP	0.92***	0.93***	1.01***	0.43***	0.40***	0.14**	1.01***	1.01***	1.14***
	(104.20)	(105.38)	(61.62)	(27.34)	(25.14)	(3.78)	(16.52)	(16.52)	(12.62)
Log product real GDP/capita	0.42***	0.44***	0.34***	0.36***	0.40***	0.45**	0.00	0.00	−0.19**
	(32.28)	(33.80)	(14.77)	(23.66)	(26.73)	(11.83)	(0.03)	(0.03)	(2.55)
Common language	0.36***	0.35***	0.29***						
	(9.25)	(9.31)	(4.66)						
Common land border	0.52***	0.40***	1.22***						
	(4.71)	(3.85)	(7.30)						
Free trade agreement (FTA)	1.14***			0.75***			0.47***		
	(10.40)			(33.41)			(3.38)	0.47***	0.61***
FTA (with Africa)		1.19***	0.92***		0.33***	0.15**		(3.40)	(3.96)
		−14.32	−7.46		(13.36)	(3.19)			
Number landlocked in the pair	−0.23***	−0.25***	−0.35***						
	(8.17)	(8.82)	(8.86)						
Number islands in the pair	0.03	0.03	−0.31***						
	(0.92)	(0.73)	(5.00)						
Log product of areas	−0.08***	−0.08***	−0.18***						
	(10.75)	(11.26)	(14.70)						
Common colonizer	0.55***	0.51***	0.30***						
	(8.78)	(8.27)	(3.48)						
Current colony	0.95***	0.99***	−0.39	0.34***	0.32***	−0.10	0.94**	0.94**	0.10
	(3.91)	(4.02)	(0.86)	(7.98)	(7.50)	(1.25)	(2.34)	(2.33)	(0.55)

Table 1 *(concluded)*

Ever colony	1.31***	1.32***	2.14***						
	(10.75)	(10.97)	(14.42)						
Same nation	−0.20	−0.23	1.63***						
	(0.20)	(0.22)	(3.30)						
Observations	265,262	265,262	100,597	265,262	265,262	100,597	380,512	380,512	176,712
R^2 (within)				0.14	0.14	0.03			
R^2 (between)				0.6	0.59	0.30			
R^2 (overall)	0.68	0.68	0.52	0.55	0.54	0.24			
Ramsey F-test (p-value)	0.00	0.00	0.00						
Fixed effects F-test (p-value)				0.00	0.00	0.00	0.00	0.00	0.00
Hausman test (p-value)				0.00	0.00	0.00	0.00	0.00	0.00
Wald χ^2 (p-value)							0.00	0.00	0.00

Source: Authors' calculations.

Note: This table reports pooled ordinary least squares (OLS), fixed effects, and Poisson pseudomaximum likelihood (PML) estimation results for the gravity model of bilateral trade using the world and Africa (at least one country in the pair is in Africa) samples. The dependent variable is log of real trade between trading partners ($\log(X_{ij})$) in the OLS and fixed effects regressions, and real trade between trading partners (X_{ij}) in the PML regressions. Real trade is defined as the average of exports and imports between the trading partners deflated by the U.S. consumer price index. The independent variables are currency union (dummy variable equal to 1 if money is interchangeable between the two countries at a 1:1 par for an extended period of time, and 0 otherwise); log of geographical distance between the trading partners; log of the product of real GDP of the trading partners; log of the product of real GDP per capita of the trading partners; common language (dummy variable equal to 1 if the two countries share a language, and 0 otherwise); common land border (dummy variable equal to 1 if the two countries share a border, and 0 otherwise); free trade agreement (Glick and Rose's (2002) dummy variable equal to 1 if the two countries share a free trade agreement, and 0 otherwise); free trade agreement with Africa (a more comprehensive version of the variable free trade agreement that also takes into account free trade agreements with African countries); number of landlocked countries in the pair (for example, 0, 1, or 2); number of islands in the pair (for example, 0, 1, or 2); log of product of land areas of the two countries; common colonizer (dummy variable equal to 1 if the trading partners share the colonizer, and 0 otherwise); current colony (dummy variable equal to 1 if one country in the pair is colonized by the other country, and 0 otherwise); ever colony (dummy variable equal to 1 if one trading partner has ever been a colony of the other, and 0 otherwise); and same nation (dummy variable equal to 1 if both partners are part of the same nation, and 0 otherwise). Year dummies included in all regressions. The table reports robust t-statistics in parentheses of columns 1–6, and robust z-statistics in columns 7–9. *, **, *** denote significance at the 10, 5, and 1 percent levels, respectively.

country-pair specific idiosyncratic effects (columns 4–6).[18] The specification in column 1 uses the world sample and replicates the benchmark pooled results of Glick and Rose (2002) very closely: countries sharing a common currency trade about two-and-a-half times more than countries not involved in a currency union. The coefficients on the standard determinants of the gravity models, such as income, population, and distance, have the correct sign, are statistically significant, and yield plausible elasticity estimates broadly in line with those obtained in earlier literature. In column 2, we allow the free trade area dummy to include additional agreements, especially those operating in Africa. The coefficient of currency union decreases slightly to 2.3 compared with the original Glick and Rose specification.

Column 3 presents the estimation results for Africa. Interestingly, the currency union trade-generating effect is larger than for the world sample, and the marginal impacts of other determinants change.[19] The effect of sharing a border is now much larger, reflecting the poor transportation links between many African countries and the tendency to trade more with neighboring countries. The impact of having a common language is however less important in Africa, as the currency unions in the sample are the Francophone countries of West and Central Africa and the South Africa Common Monetary Area. The variable reflecting having been a colony (ever colony) has a greater impact that reflects the colonial ties of African countries. Surprisingly, there is a reversal in the sign of the coefficient of number of islands, which is hard to interpret. However, this may be a result of the omitted-variable bias caused by ignoring country-specific fixed effects.

The Ramsey regression specification error test for omitted variables confirms the existence of omitted variables in the three pooled regressions. Thus, in columns 4–6, we re-estimate the earlier specifications as a panel and add a set of country-pair specific intercepts to the equation.[20] The *F*-test for the joint significance of the fixed effects shows that the fixed effects are significant in each case. The magnitude of the trade-generating coefficient changes, which suggests that not accounting for country-specific idiosyncratic effects may lead to biased estimates. For both the world and Africa, the size of the currency union coefficient falls, and the trade-generating effect becomes 1.8 and 1.7 for the world and Africa, respectively.

[18]Henceforth, unless stated otherwise, our panel and PML estimates include country-pair fixed effects along the lines of Glick and Rose (2002), as well as time effects. In addition, country fixed effects were also used to account for the Anderson and van Wincoop (2003) "multilateral resistance" factor in pooled OLS estimations and are available from the authors on request. In the interest of clarity, we do not show results of all the pooled OLS country fixed effects estimates in Tables 1–3, but some are summarized in Table 4.

[19]Masson and Pattillo (2004) also estimate the gravity equation for Africa. However, their results may not be directly comparable to ours because they do not take into account free trade areas operating in Africa.

[20]Although we estimate both fixed effects "within" and random effects, we rely on the robust fixed effects within estimator as suggested by the Hausman test.

Turning to the issue of zero-trade observations, columns 7–9 in Table 1 present the results when the PML approach is used. Not eliminating the zero-trade observations increases the world and Africa samples by about 45 and 31 percent, respectively. The results show that the effect of currency unions is still positive and significant in all cases but much smaller in magnitude. Currency unions now increase trade by a factor of about 0.2 and 0.7 in the world and in Africa, respectively. The inclusion of zero observations reduces the trade-generating elasticity for the world sample by more than the elasticity for Africa; hence, the effect of currency unions on trade is now larger for Africa than for the world.[21]

Characteristics of trade in Africa

Next, we investigate the characteristics of Africa's trade in more detail. Table 2 presents the results of the relative trade performance of African countries by introducing dummies for intra-Africa trade, and trade between Africa and the ROW in Equation (1). We first estimate the augmented equation for the world sample for comparison purposes and observe that the effect of currency unions on trade changes only marginally: the trade-generating effect of a currency union is now 2.2. Both the Africa-ROW and intra-Africa trade dummies are positive and highly significant, indicating that if at least one of the bilateral trading partners is in Africa, trade increases by no less than 1.3 times.

Columns 2 and 3 of Table 2 present the results when the sample is confined to those trading pairs where one partner country is in Africa, and where both partner countries are in Africa, respectively. In both cases, membership in a currency union increases trade by about 1.7 times. In columns 5–7 panel fixed effects ‘‘within’’ estimates are shown for the world, Africa-ROW, and intra-Africa samples, respectively. The results confirm that currency union participation in Africa has beneficial intraregional trade-generating effects as well as Africa-ROW trade-generating effects: the trade-generating effect is 1.7 for intra-African trade and 1.9 for Africa-ROW trade. Columns 8–10 employ the PML estimation technique. Once again, the magnitude of the coefficient for currency union is smaller than for pooled OLS and fixed-effects estimators.

[21]To explore this issue in detail, we ‘‘decompose’’ the trade-generating coefficient for the world sample into the coefficients of two homogeneous groups—the industrialized and nonindustrialized countries—using fixed effects within estimator and PML. The results suggest that the PML technique is more sensitive to sample homogeneity than the other estimator. Thus, for example, the estimated coefficient for the currency union variable obtained from PML (fixed effects) is 0.14 (0.37) and 1.07 (0.51) for the industrialized and nonindustrialized groups, respectively. The world weighted average of 0.15 under PML appears to be influenced by the industrialized group average, and hence is smaller in magnitude. It is worth pointing out here that some other studies that estimate the gravity model using PML also obtain smaller estimated coefficients vis-à-vis other techniques, such as Tobit (see, for example, Amurgo-Pacheco and Pierola, 2008).

Table 2. Africa Trade Details

Sample:	World	Africa-ROW	Intra-Africa	World	World	Africa-ROW	Intra-Africa	World	Africa-ROW	Intra-Africa
Estimation:	OLS	OLS	OLS	OLS	Fixed effects	Fixed effects	Fixed effects	PML	PML	PML
Specification:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Currency union	0.79***	0.53	0.55***	0.76***	0.60***	0.55***	0.43***	0.15**	0.60***	0.62***
	(7.03)	(1.18)	(2.69)	(6.23)	(16.15)	(7.77)	(4.94)	(2.51)	(4.30)	(2.70)
Africa-ROW	0.27***									
	(7.90)									
Intra-Africa	0.42***									
	(4.58)									
Log distance	−1.06***	−0.96***	−1.29***	−1.06***						
	(52.00)	(15.64)	(9.92)	(−52.27)						
Log product real GDP	0.94***	1.06***	0.62***	0.93***	0.40***	0.11***	0.88***	1.01***	1.18***	0.60***
	(104.26)	(64.47)	(9.83)	(104.11)	(25.14)	(2.78)	(4.59)	(16.52)	(12.34)	(5.05)
Log product real GDP/capita	0.48***	0.33***	0.50***	0.47***	0.40***	0.51***	−0.46**	0.00	−0.13	−0.10
	(35.94)	(13.95)	(5.80)	(35.04)	(26.73)	(11.79)	(2.44)	(0.03)	(1.54)	(0.79)
Common language	0.33***	0.27***	0.40**	0.34***						
	(8.76)	(4.26)	(2.01)	(8.76)						
Common land border	0.43***	1.10***	0.98***	0.42***						
	(4.17)	(3.21)	(4.27)	(4.04)						
Free trade agreement (with Africa)	1.15***		0.78***	1.18***	0.33***		0.31***	0.47***		0.30**
	(13.08)		(5.24)	(13.81)	(13.36)		(5.56)	(3.40)		(2.09)
Number landlocked in the pair	−0.28***	−0.37***	−0.33***	−0.28***						
	(9.97)	(8.82)	(3.07)	(−9.82)						
Number islands in the pair	0.04	−0.39***	0.17	0.03						
	(1.09)	(6.07)	(0.83)	(0.88)						
Log product of areas	−0.09***	−0.20***	−0.01	−0.09***						
	(12.35)	(16.14)	(0.15)	(−11.88)						
Common colonizer	0.49***	0.32***	0.25	0.51***						
	(7.91)	(3.46)	(1.09)	(8.02)						
Current colony	0.98***	−0.03		1.02***	0.32***	−0.07		0.94**	0.11	
	(3.92)	(0.06)		(3.99)	(7.50)	(0.82)		(2.33)	(0.59)	
Ever colony	1.30***	2.12***		1.31***						
	(11.25)	(14.08)		(11.16)						

Table 2 *(concluded)*

Same nation	−0.31	1.47***		−0.30						
	(0.31)	(2.62)		(−0.30)						
Intra-Francophone				0.36**						
				(2.13)						
Intra-Anglophone				0.02						
				(0.10)						
Francophone-ROW				0.19***						
				(4.81)						
Anglophone-ROW				0.22***						
				(5.19)						
Observations	265,262	85,759	14,838	265,262	265,262	85,759	14,838	380,512	146,468	30,244
R^2 (within)					0.14	0.04	0.02			
R^2 (between)					0.59	0.27	0.16	0.00	0.00	0.00
R^2 (overall)	0.68	0.54	0.40	0.68	0.54	0.21	0.11	0.00	0.00	0.00
Hausman test (p-value)					0.00	0.00	0.00	0.00	0.00	0.00

Source: Authors' calculations.

Note: This table reports pooled ordinary least squares (OLS), fixed effects, and Poisson pseudomaximum likelihood (PML) estimation results for the gravity model of bilateral trade using the world, Africa (at least one country in the pair is in Africa), intra-Africa (both countries in the pair are in Africa), and Africa-rest of the world (ROW) (only one country in the pair is in Africa) samples. The dependent variable is log of real trade between trading partners ($\log(X_{ij})$) in the OLS and fixed effects regressions, and real trade between trading partners (X_{ij}) in the PML regressions. Real trade is defined as the average of exports and imports between the trading partners deflated by the U.S. consumer price index. The independent variables are currency union (dummy variable equal to 1 if money is interchangeable between the two countries at a 1:1 par for an extended period of time, and 0 otherwise); log of geographical distance between the trading partners; log of the product of real GDP of the trading partners; log of the product of real GDP per capita of the trading partners; common language (dummy variable equal to 1 if the two countries share a language, and 0 otherwise); common land border (dummy variable equal to 1 if the two countries share a border, and 0 otherwise); free trade agreement with Africa (a more comprehensive version of the variable free trade agreement that also takes into account free trade agreements with African countries); number of landlocked countries in the pair (for example, 0, 1, or 2); number of islands in the pair (for example, 0, 1, or 2); log of product of land areas of the two countries; common colonizer (dummy variable equal to 1 if the trading partners share the colonizer, and 0 otherwise); current colony (dummy variable equal to 1 if one country in the pair is colonized by the other country, and 0 otherwise); ever colony (dummy variable equal to 1 if one trading partner has ever been a colony of the other, and 0 otherwise); same nation (dummy variable equal to 1 if both partners are part of the same nation, and 0 otherwise); intra-Francophone (dummy variable equal to 1 if the both partners are African Francophone countries, and 0 otherwise); intra-Anglophone (dummy variable equal to 1 if both partners are African Anglophone countries, and 0 otherwise); Francophone-ROW (dummy variable equal to 1 if only one country in the pair is an African Francophone country, and 0 otherwise); and Anglophone-ROW (dummy variable equal to 1 if only one country in the pair is an African Anglophone country, and 0 otherwise). Year dummies included in all regressions. The table reports robust t-statistics in parentheses of columns 1–7, and robust z-statistics in columns 8–10. *, **, *** denote significance at the 10, 5, and 1 percent levels, respectively.

Finally, following Subramanian and Tamirisa (2003), we also investigate whether Anglophone and Francophone African countries differ in their trading characteristics (column 4 of Table 2). The results show that the trade-generating effect remains about 2, but there is some evidence of differences in the trade patterns of Francophone and Anglophone Africa. The intra-Francophone trade coefficient is significant, suggesting that trade among the Francophone Africa countries is about 43 percent more than the average. This suggests that even after controlling for the currency union effect (essentially the effects of the CFA franc zone) trade among Francophone Africa has increased. We also find evidence of increasing integration of both Francophone and Anglophone Africa with the ROW, confirming the recent pattern of increase in intra-African trade as well as Africa-Europe trade. The coefficients of trade with the ROW are positive and significant, suggesting that in Francophone and Anglophone Africa, overall trade is about 21 and 23 percent more than the average, respectively.

Trade creation or diversion?

In Table 3, we investigate the possibility that the stimulus to trade among members of a currency union comes at the expense of trade with nonmembers. To do so, we add the trade diversion dummy to the specifications of Tables 1 and 2. All the OLS and fixed-effects specifications show that the coefficient associated with the trade diversion dummy is positive. This suggests that there is a significant trade creation (rather than trade diversion) effect, but this effect is smaller for Africa than for the world. Moreover, although there is a significant trade creation effect in both intra-Africa and Africa-ROW specifications, the intra-Africa trade creation effect is about five times larger, which reflects the observed trade patterns in Africa's trade arrangements. The PML estimation for the world sample, however, gives a negative estimated coefficient for the trade diversion dummy, and insignificant coefficients for the Africa and Africa-ROW samples, suggesting that the earlier results may not be robust.[22] The evidence from all estimation methods suggests that currency unions are associated with trade creation within Africa, which is significant and almost identical in magnitude for the PML and fixed-effects estimates.

Sensitivity analysis

We check the robustness of our results presented in Tables 1 and 2, by changing our methodology in a number of ways and reporting estimates of the coefficients of interest. In particular, we conduct sensitivity checks of the

[22]However, when we decompose the world sample into homogeneous groups we find that for the nonindustrialized group the trade creation effect is significantly positive (0.22), but a trade diversion effect exists for the industrialized group. As discussed in footnote 21, it appears that the PML result for the world in this case is also influenced by the industrial group observations.

specification by estimating the dependent variable as the average of the logarithm of exports and imports (rather than the logarithm of the average); adding quadratic terms for output and output per capita to control for possible sample nonlinearities; modifying the dependent variable by adding a constant to trade observations before taking the logarithm to avoid truncation of the data;[23] and using Tobit to account for the censored nature of the dependent variable.

The results in Table 4 show the estimates of the currency union trade-generating effect γ for the various robustness checks. First, for the benchmark panel fixed effects estimates, adding nonlinear terms and time effects, and changing the specification of the dependent variable to the average of the logs has a marginal effect on the size of γ. The trade-generating effect is about 1.8 for the world and 1.7 for Africa (which is smaller than that estimated using pooled OLS). Next, including zero-trade observations makes a difference to the results. This is because excluding zero values tends to drive the overall elasticity upward, whereas their inclusion drives it downward.[24] The Tobit estimates that overcome the bias that may result from the censored nature of the data are sufficiently close to the cross-section results obtained in Table 1. However, they may suffer from the omitted-variable bias resulting from not controlling for country-pair specific fixed effects. The PML estimates that take into account all zero-trade observations as well as country-pair fixed effects find a positive and significant but smaller impact of currency unions on trade.

Despite the differences in the magnitudes among estimation methods, the results always show a statistically significant trade-generating effect for the Africa sample, which is comparable to the world sample trade-generating effect. Overall, the currency union trade-generating effect is strong for Africa, and may be larger than that for the world. However, the magnitude of the currency union trade-generating effect seems to be sensitive to the inclusion of zero-trade observations as well as the estimation methodology employed, particularly when zero observations are included and the PML approach is used. Taken together, all panel estimates (fixed effects "within" and PML) suggest that the average trade-generating effect is 1.2 for the world and 1.4 for Africa.

Years of Membership, Comovements, and Trade Stability

We now turn our attention to three different but related issues. In particular, we investigate the impact of the duration of currency union membership on bilateral trade, the currency union impact on output and price comovements, and the effect of currency union membership on trade stability.

[23]To compare our results with those obtained by Alesina and Tenreyro (2002) we set the constant equal to 100. The dependent variable therefore becomes $\log(100 + X_{ij})$.

[24]See Amemiya (1984) for a detailed discussion on this issue.

Table 3. Trade Creation or Diversion?

Sample:	World	Africa	Africa-ROW	Intra-Africa	World	Africa	Africa-ROW	Intra-Africa	World	Africa	Africa-ROW	Intra-Africa
Estimation:	OLS	OLS	OLS	OLS	Fixed effects	Fixed effects	Fixed effects	Fixed effects	PML	PML	PML	PML
Specification:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Currency union	0.97***	1.05***	0.61	0.59***	0.77***	0.75***	0.74***	0.91***	0.12**	0.72***	0.64***	0.91***
	(8.77)	(6.81)	(1.38)	(2.58)	(20.42)	(10.77)	(10.00)	(9.15)	(2.16)	(4.10)	(3.39)	(3.92)
Trade diversion	0.23***	0.10**	0.09	0.09	0.23***	0.22***	0.18***	0.52***	−0.20***	0.10	0.03	0.51***
	(8.66)	(2.35)	(1.92)	(0.55)	(20.81)	(11.66)	(9.22)	(8.52)	(3.63)	(0.80)	(0.46)	(3.53)
Log distance	−1.05***	−1.05***	−0.96***	−1.29***								
	(52.02)	(19.62)	(15.61)	(9.86)								
Log product real GDP	0.93***	1.01***	1.06***	0.63***	0.38***	0.13***	0.11***	0.80***	1.01***	1.14***	1.17***	0.64***
	(105.28)	(61.82)	(64.45)	(9.89)	(24.23)	(3.49)	(2.65)	(4.20)	(16.59)	(12.68)	(12.38)	(4.90)
Log product real GDP/capita	0.44***	0.33***	0.32***	0.50***	0.41***	0.45***	0.51***	−0.40**	−0.01	−0.18***	−0.13	−1.40
	(33.56)	(14.46)	(13.64)	(5.61)	(27.56)	(11.93)	(11.76)	(2.15)	(0.17)	(2.58)	(1.56)	(1.24)
Common language	0.33***	0.28***	0.27***	0.40**								
	(8.80)	(4.57)	(4.20)	(2.00)								
Common land border	0.42***	1.22***	1.11***	0.98***								
	(4.04)	(7.29)	(3.19)	(4.26)								
Free trade agreement (with Africa)	1.17***	0.93***		0.80***	0.34***	0.18***		0.40***	0.47***	0.62***		0.21
	(14.03)	(7.55)		(5.13)	(13.68)	(3.80)		(6.98)	(3.39)	(4.07)		(1.49)
Number landlocked in the pair	−0.24***	−0.35***	−0.36***	−0.33***								
	(8.38)	(8.71)	(8.70)	(3.03)								
Number islands in the pair	0.01	−0.31***	−0.39***	0.17								
	(0.24)	(5.01)	(6.07)	(0.86)								
Log product of areas	−0.08***	−0.18***	−0.20***	−0.01								
	(11.48)	(14.82)	(16.19)	(0.20)								
Common colonizer	0.52***	0.31***	0.33***	0.26								
	(8.42)	(3.60)	(3.51)	(1.15)								
Current colony	1.03***	−0.35	−0.02		0.38***	−0.08	−0.06		0.86**	0.10	0.10	
	(4.17)	(0.77)	(0.03)		(8.72)	(0.95)	(0.69)		(2.32)	(0.54)	(0.58)	
Ever colony	1.25***	2.10***	2.09***									
	(10.57)	(13.94)	(13.65)									

Table 3 *(concluded)*

Same nation	−0.25	1.60***	1.46***									
	(0.24)	(3.22)	(2.59)									
Observations	265,262	100,597	85,759	14,838	265,262	100,597	85,759	14,838	380,512	176,712	146,468	30,244
R^2 (within)					0.14	0.03	0.04	0.03				
R^2 (between)					0.59	0.27	0.26	0.17				
R^2 (overall)	0.68	0.52	0.54	0.40	0.53	0.22	0.20	0.12				
Hausman test (p-value)					0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Wald χ^2					0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00

Source: Authors' calculations.

Note: This table reports pooled ordinary least squares (OLS), fixed effects, and Poisson pseudomaximum likelihood (PML) estimation results for the trade diversion effect of currency unions using the world, Africa (at least one country in the pair is in Africa), intra-Africa (both countries in the pair are in Africa), and Africa-rest of the world (ROW) (only one country in the pair is in Africa) samples. The dependent variable is log of real trade between trading partners ($\log(X_{ij})$) in the OLS and fixed effects regressions, and real trade between trading partners (X_{ij}) in the PML regressions. Real trade is defined as the average of exports and imports between the trading partners deflated by the U.S. consumer price index. The independent variables are currency union (dummy variable equal to 1 if money is interchangeable between the two countries at a 1:1 par for an extended period of time, and 0 otherwise); trade diversion (dummy variable equal to 1 if the trading partners are not in the same currency union but at least one is in a currency union with another country); log of geographical distance between the trading partners; log of the product of real GDP of the trading partners; log of the product of real GDP per capita of the trading partners; common language (dummy variable equal to 1 if the two countries share a language, and 0 otherwise); common land border (dummy variable equal to 1 if the two countries share a border, and 0 otherwise); free trade agreement with Africa (a more comprehensive version of the variable free trade agreement that also takes into account free trade agreements with African countries); number of landlocked countries in the pair (for example, 0, 1, or 2); number of islands in the pair (for example, 0, 1, or 2); log of product of land areas of the two countries; common colonizer (dummy variable equal to 1 if the trading partners share the colonizer, and 0 otherwise); current colony (dummy variable equal to 1 if one country in the pair is colonized by the other country, and 0 otherwise); ever colony (dummy variable equal to 1 if one trading partner has ever been a colony of the other, and 0 otherwise); and same nation (dummy variable equal to 1 if both partners are part of the same nation, and 0 otherwise). Year dummies included in all regressions. The table reports robust t-statistics in parentheses of columns 1–8, and robust z-statistics in columns 9–12. *, **, *** denote significance at the 10, 5, and 1 percent levels, respectively.

Table 4. Sensitivity Analysis of Currency Union Effect
(Various Estimation Methods)

	World		Africa	
Sample:	γ	Effect	γ	Effect
Ordinary least squares				
With time effects (from Table 1)	0.84***	2.3	0.99***	2.7
	(7.75)		(6.57)	
Without time effects	0.91***	2.5	0.97***	2.6
	(7.69)		(5.70)	
With time and country fixed effects	0.85**	2.3	1.14	3.1
(Anderson van-Wincoop)	(7.86)		(7.82)	
Dependent variable $\log(100 + X_{ij})$ (with time effects)	0.57***	1.8	0.58***	1.8
	(4.85)		(4.05)	
Tobit estimation (with time effects)	0.79***	2.2	0.85***	2.4
	(20.30)		(16.09)	
Fixed effects				
With time effects (from Table 1)	0.60***	1.8	0.54***	1.7
	(16.15)		(7.91)	
Without time effects	0.58***	1.8	0.55***	1.7
	(15.54)		(7.94)	
Dependent variable calculated as average of logs	0.65***	1.9	0.60***	1.8
(with time effects)	(17.17)		(8.86)	
Nonlinearities added (with time effects)	0.45***	1.6	0.50***	1.7
	(11.83)		(7.23)	
PML panel estimates with fixed effects including zero trade observations				
With time effects (from Table 1)	0.15**	0.2	0.65***	0.7
	(2.51)		(5.05)	
Without time effects	0.09	0.1	0.65***	0.7
	(1.34)		(7.20)	

Source: Authors' calculations.

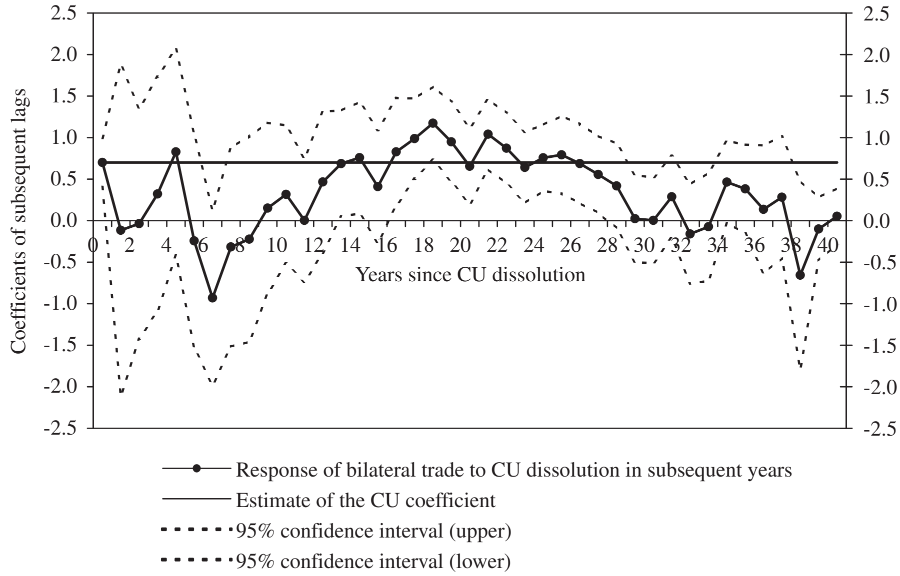
Note: This table presents the results of the currency union trade-generating effect for various robustness checks using the world and Africa samples. The dependent variable is log of real trade ($\log(X_{ij})$) in the ordinary least squares (OLS) and fixed effects regressions, and real trade (X_{ij}) in the Poisson pseudomaximum likelihood (PML) regressions. The table reports robust t-statistics in parentheses of OLS and fixed effects estimation results and robust z-statistics in parentheses of PML results.

*, **, *** denote significance at the 10, 5, and 1 percent levels, respectively.

Membership duration

To investigate the "time effect" of currency union membership, we examine how the effects of leaving a currency union evolve over time. To do so, we follow Glick and Rose (2002) and define a dummy variable that is equal to

Figure 1. Estimated Impact of Dissolution of Currency Union (CU) on Trade: World Sample



Notes: The horizontal lines in this figure correspond to the estimate of the coefficient of the currency union dummy (the γ coefficient for the gravity equation (1) extended to include the vector of lagged variables described in the text) for the world and Africa samples, respectively. The corresponding lag of the dummy variable associated with CU dissolution is statistically significant if the corresponding error bands exclude zero.

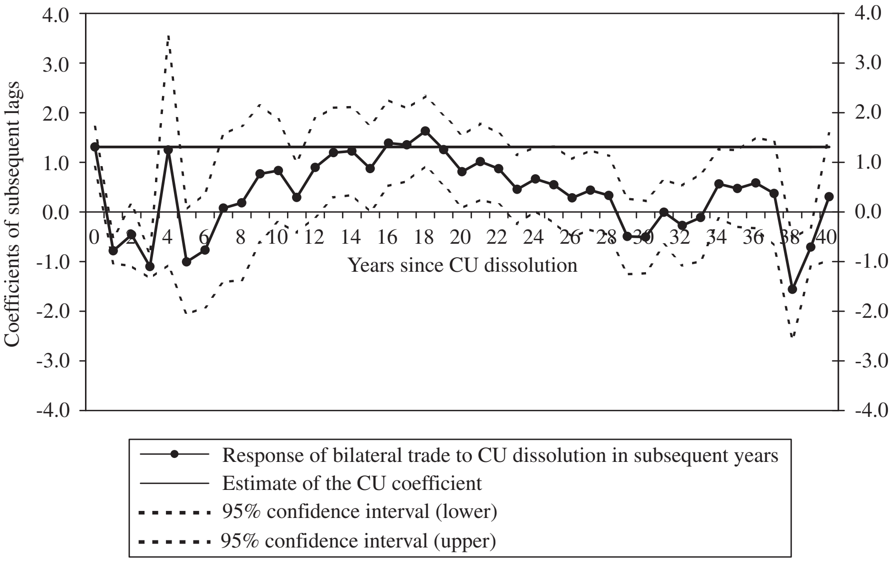
unity for the observations associated with the year a union is dissolved between a given pair of countries, and equal to zero otherwise.[25] Importantly, we also include lags of this dummy variable in the equation, that is, we add a dummy variable for one year after dissolution, for two years after dissolution, and so on.

Figures 1 and 2 plot the coefficients of subsequent lags to trace out the response of bilateral trade (the "response coefficients") to the dissolution of a currency union. This clearly illustrates the impact of currency union dissolution on trade over time.[26] As the time-dimension of our data set is

[25]See Tables A2–A4 for the number of identified currency union dissolutions for the world and Africa samples.

[26]The horizontal lines in Figures 1 and 2 correspond to the estimate of the coefficient of the currency union dummy (namely, the γ coefficient in equation (1) extended to include the vector of lagged variables described in the text) for the world and Africa, respectively. The 95 percent confidence intervals are also included.

Figure 2. Estimated Impact of Dissolution of Currency Union (CU) on Trade: Africa Sample



Notes: The horizontal lines in this figure correspond to the estimate of the coefficient of the currency union dummy (the γ coefficient for the gravity equation (1) extended to include the vector of lagged variables described in the text) for the world and Africa samples, respectively. The corresponding lag of the dummy variable associated with CU dissolution is statistically significant if the corresponding error bands exclude zero.

longer than that used by Glick and Rose, we show more convincingly that exit from a currency union is associated with a decline in bilateral trade.

For both samples, trade is mostly lower after the dissolution of a currency union compared with trade during the currency union. This is particularly true for Africa, where—with the exception of a small "blip" that occurs in the 18th year—the "response coefficients" are estimated to be substantially lower than before the dissolution. For the world sample, the "response coefficients" are around the level of trade during the currency union until about 27 years after the exit. Further, the adverse effect of exiting a currency union is smaller for Africa in the first few years but increases later. After 10 years of currency union exit, bilateral trade declines cumulatively by about 30 percent for Africa and 45 percent for the world, but 25 years after the exit, the cumulative decline in bilateral trade is greater in Africa than in the world sample.

Statistically significant effects of currency union dissolution are visible (in both samples) after about 13 years. However, after about 22 and 28 years for Africa and the world samples, respectively, these effects become gradually negligible (and insignificantly different from zero), suggesting that the effect of the currency union dissolution disappears. Overall, the results suggest that leaving a currency union has a significant impact on trade (perhaps more for Africa than the world), but this relationship is far from linear. This observation motivates the inclusion of a quadratic term in the currency union duration effect analysis.

In addition to the dissolution of currency union membership, we examine whether the *duration* of sharing a common currency matters for trade. We use a novel approach that introduces (sequentially) to Equation (1) a variable that measures the number of years that a given trading partner has shared a common currency, and the variable's quadratic term to capture potential nonlinear effects of currency union duration. Results presented in Table 5 show that the duration of currency union membership is important in all cases. According to the fixed-effects estimates, one additional year of membership increases trade by about 1.4 and 2.5 percent for the world and Africa, respectively. Holding all else constant, after 10 years of currency union membership, trade increases by about 15 percent for the world and 28 percent for Africa. This implies that currency union participation would double the level of trade in about 67 years for the world and in 36 years for Africa. Interestingly, the estimated coefficient for the quadratic term is negative and statistically significant (albeit small) in all specifications. This result points at diminishing returns of currency union membership.

Comovements of prices and output

In order to investigate the effects of currency unions on the comovements of prices and output we follow the approach of Alesina and Tenreyro (2002), and Tenreyro and Barro (2003). Specifically, we test whether their results hold in a bilateral approach to currency unions, especially in the context of Africa. Hence, we estimate Equation (1), sequentially substituting the dependent variable with VP_{ij} and VY_{ij} as calculated in Equations (4) and (6).[27] The results of the estimated impact on comovements of output and prices are presented in Table 6. We find that currency unions tend to increase the comovement of prices in both the world and Africa, but that they are not systematically related to the comovement of outputs. These findings support the results of Alesina and Tenreyro (2002). However, we also find that the marginal effect of currency unions on the price comovements is higher in Africa. This is an important observation because price comovements tend to be lower in Africa (see summary statistics in Tables A2–A4). The insignificant effect of currency unions on the comovement of outputs is not surprising considering that the theoretical link between the two is ambiguous, and largely depends on the extent to which trade is intra- or

[27]The equations are estimated using the average values of the period 1948–2003 and pooled OLS. Because we cannot use panel estimation here, we account for country fixed effects as in Anderson and Van Wincoop (2003).

Table 5. Duration of Currency Union Membership

Sample:	World	World	Africa	Africa	World	World	Africa	Africa	World	World	Africa	Africa
Estimation:	OLS	OLS	OLS	OLS	Fixed effects	Fixed effects	Fixed effects	Fixed effects	PML	PML	PML	PML
Specification:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Years in currency union	0.03***	0.07***	0.03***	0.06***	0.01***	0.04***	0.02***	0.04***	0.01*	0.05***	0.03***	0.06***
	(7.47)	(7.49)	(5.99)	(5.58)	(8.05)	(13.36)	(8.54)	(7.84)	(1.73)	(3.80)	(4.71)	(4.94)
(Years in currency union)		−0.001***		−0.001***		−0.001***		−0.001***		−0.002***		−0.001***
		(4.86)		(3.26)		(10.89)		(4.81)		(3.42)		(2.94)
Log distance	−1.05***	−1.05***	−1.06***	−1.06***								
	(52.02)	(52.00)	(19.88)	(19.82)								
Log product real GDP	0.93***	0.93***	1.01***	1.01***	0.38***	0.39***	0.12***	0.13***	1.01***	1.01***	1.16***	1.15***
	(105.60)	(105.64)	(61.60)	(61.70)	(24.39)	(24.81)	(3.04)	(3.38)	(16.40)	(16.59)	(12.54)	(12.54)
Log product real GDP/capita	0.44***	0.44***	0.34***	0.34***	0.42***	0.41***	0.48***	0.46***	0.01	−0.01	−0.16**	−0.18**
	(33.88)	(33.87)	(14.77)	(14.72)	(27.60)	(26.88)	(12.57)	(11.95)	(0.17)	(0.08)	(2.11)	(2.35)
Common language	0.35***	0.35***	0.29***	0.29***								
	(9.26)	(9.22)	(4.74)	(4.67)								
Common land border	0.41***	0.40***	1.23***	1.23***								
	(3.94)	(3.89)	(7.29)	(7.29)								
Free trade agreement (with Africa)	1.19***	1.19***	0.91***	0.92***	0.33***	0.35***	0.13***	0.14***	0.47***	0.47***	0.62***	0.62***
	(14.35)	(14.41)	(7.36)	(7.42)	(13.38)	(13.94)	(2.69)	(2.85)	(3.40)	(3.40)	(3.87)	(3.94)
Number landlocked in the pair	−0.25***	−0.25***	−0.36***	−0.36***								
	(8.82)	(8.82)	(8.90)	(8.88)								
Number islands in the pair	0.02	0.02	−0.31***	−0.31***								
	(0.67)	(0.68)	(5.00)	(5.02)								
Log product of areas	−0.08***	−0.08***	−0.18***	−0.18***								
	(11.30)	(11.38)	(14.67)	(14.74)								
Common colonizer	0.51***	0.50***	0.31***	0.30***								
	(8.24)	(8.06)	(3.53)	(3.41)								

Table 5 *(concluded)*

Current colony	1.17***	1.06***	−0.04	−0.21	0.51***	0.41***	0.09	−0.03	0.97**	0.86**	0.29	0.22
	(4.80)	(4.49)	(0.09)	(0.47)	(12.19)	(9.58)	(1.14)	(0.36)	(2.39)	(2.31)	(1.29)	(1.03)
Ever colony	1.34***	1.33***	2.15***	2.14***								
	(11.10)	(11.04)	(14.69)	(14.60)								
Same nation	−0.24	−0.22	1.62***	1.73***								
	(0.21)	(0.19)	(3.23)	(3.60)								
Observations	265,262	265,262	100,597	100,597	265,262	265,262	100,597	100,597	380,512	380,512	176,712	176,712
R^2 (within)					0.14	0.14	0.03	0.03				
R^2 (between)					0.59	0.59	0.25	0.28				
R^2 (overall)	0.68	0.68	0.52	0.52	0.53	0.54	0.21	0.22				
Hausman test (*p*-value)					0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Wald χ^2 (*p*-value)									0.00	0.00	0.00	0.00

Source: Authors' calculations.

Note: This table reports pooled ordinary least squares (OLS), fixed effects, and Poisson pseudomaximum likelihood (PML) estimation results for the trade creation effect of currency union membership duration using the world and Africa (at least one country in the pair is in Africa) samples. The dependent variable is log of real trade between trading partners ($\log(X_{ij})$) in the OLS and fixed effects regressions, and real trade between trading partners (X_{ij}) in the PML regressions. Real trade is defined as the average of exports and imports between the trading partners deflated by the U.S. consumer price index. The independent variables are years in currency union (number of years since joining the currency union) and its square; log of geographical distance between the trading partners; log of the product of real GDP of the trading partners; log of the product of real GDP per capita of the trading partners; common language (dummy variable equal to 1 if the two countries share a language, and 0 otherwise); common land border (dummy variable equal to 1 if the two countries share a border, and 0 otherwise); free trade agreement with Africa (a more comprehensive version of the variable free trade agreement that also takes into account free trade agreements with African countries); number of landlocked countries in the pair (for example, 0, 1, or 2); number of islands in the pair (for example, 0, 1, or 2); log of product of land areas of the two countries; common colonizer (dummy variable equal to 1 if the trading partners share the colonizer, and 0 otherwise); current colony (dummy variable equal to 1 if one country in the pair is colonized by the other country, and 0 otherwise); ever colony (dummy variable equal to 1 if one trading partner has ever been a colony of the other, and 0 otherwise); and same nation (dummy variable equal to 1 if both partners are part of the same nation, and 0 otherwise). Year dummies included in all regressions. The table reports robust *t*-statistics in parentheses of columns 1–8, and robust *z*-statistics in columns 9–12. *, **, *** denote significance at the 10, 5, and 1 percent levels, respectively.

Table 6. Currency Union Impact on Comovements of Outputs and Prices

(Dependent Variables: Comovement of Outputs (Vyij) for (1)–(4); and Comovement of Prices (Vpij) for (5)–(8))

Sample:	World	Africa	World	Africa	World	Africa	World	Africa
Specification:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Currency union	−0.002	0.002	0.000	0.000	0.06***	0.07***	0.03***	0.04***
	(0.61)	(0.55)	(0.12)	(0.25)	(9.07)	(9.35)	(11.19)	(11.19)
Log distance	0.002***	0.008***	−0.001***	−0.001	−0.003*	0.001	−0.004***	−0.003**
	(3.56)	(7.58)	(3.56)	(0.96)	(1.79)	(0.26)	(5.33)	(2.23)
Log product real GDP	0.005***	0.005***	0.003***	0.000	0.002***	−0.003***	0.007***	−0.008***
	(22.10)	(14.19)	(5.21)	(0.30)	(4.18)	(3.96)	(3.34)	(3.57)
Log product real GDP/capita	0.005***	0.001	−0.004***	0.001	0.012***	0.016***	−0.004	0.011***
	(16.98)	(1.63)	(4.39)	(0.56)	(14.76)	(16.61)	(1.17)	(3.47)
Common language	0.001	−0.002	0.001**	0.000	0.001	0.000	0.002	0.001
	(0.65)	(1.54)	(2.05)	(0.13)	(0.42)	(0.09)	(1.43)	(1.15)
Common land border	0.006**	0.005	0.003***	0.001	−0.005	−0.007	0.004	0.004
	(2.52)	(1.44)	(3.65)	(0.82)	(0.66)	(0.77)	(1.46)	(1.24)
Free trade agreement (with Africa)	0.012***	0.010***	0.003**	−0.001	0.018***	0.015*	0.007**	−0.007**
	(4.59)	(2.65)	(2.35)	(0.60)	(2.84)	(1.89)	(2.50)	(2.21)
Number landlocked in the pair	0.005***	0.007***			0.007***	0.014***		
	(8.19)	(8.65)			(3.88)	(7.30)		
Number islands in the pair	0.005***	0.010***			0.012***	0.012***		
	(7.14)	(8.08)			(6.77)	(5.45)		
Log product of areas	0.000***	0.000			−0.004***	−0.001*		
	(3.07)	(0.66)			(9.62)	(1.94)		
Common colonizer	0.004***	0.009***	0.002***	0.001	0.010***	0.002	0.005***	0.006***
	(3.14)	(5.45)	(3.49)	(0.98)	(3.35)	(0.50)	(4.06)	(4.48)
Current colony	−0.018*	−0.044	−0.007**	−0.002	0.049*	−0.040	0.025**	−0.020
	(1.70)	(1.47)	(2.03)	(0.23)	(1.90)	(0.68)	(2.15)	(0.70)

Table 6 *(concluded)*

Ever colony	0.004*	0.018**	0.000	0.001	0.007	0.034*	−0.006	0.005
	(1.70)	(2.01)	(0.03)	(0.55)	(0.72)	(1.77)	(1.40)	(0.70)
Same nation								
Country fixed effects	No	No	Yes	Yes	No	No	Yes	Yes
Observations	6,992	3,793	6,992	3,793	7,000	3,796	7,000	3,796
R^2	0.25	0.15	0.92	0.92	0.12	0.13	0.87	0.91

Source: Authors' calculations.

Note: This table reports pooled ordinary least squares (OLS) and fixed effects estimation results for the effects of currency unions on the comovements of prices and output using the world and Africa (at least one country in the pair is in Africa) samples. The independent variables are currency union (dummy variable equal to 1 if money is interchangeable between the two countries at a 1:1 par for an extended period of time, and 0 otherwise); log of geographical distance between the trading partners; log of the product of real GDP of the trading partners; log of the product of real GDP per capita of the trading partners; common language (dummy variable equal to 1 if the two countries share a language, and 0 otherwise); common land border (dummy variable equal to 1 if the two countries share a border, and 0 otherwise); free trade agreement with Africa (a more comprehensive version of the variable free trade agreement (for example, 0, 1, or 2); log of product of land areas of the two countries; common colonizer (dummy variable equal to 1 if the trading partners share the colonizer, and 0 otherwise); current colony (dummy variable equal to 1 if one country in the pair is colonized by the other country, and 0 otherwise); ever colony (dummy variable equal to 1 if one trading partner has ever been a colony of the other, and 0 otherwise); and same nation (dummy variable equal to 1 if both partners are part of the same nation, and 0 otherwise). The table reports robust *t*-statistics in parentheses. *, **, *** denote significance at the 10, 5, and 1 percent levels, respectively.

Table 7. Currency Union Impact on Trade Stability

(Dependent Variable is Trade Variability, the Coefficient of Variation of $\log(X_{ij})$)

Sample:	World	Africa	World	Africa	World	Africa
Estimation:	OLS	OLS	Fixed effects	Fixed effects	PML	PML
Specification:	(1)	(2)	(3)	(4)	(5)	(6)
Currency union	−0.04**	−0.06***	−0.02	−0.01	0.08	−0.01
	(3.41)	(4.15)	(0.84)	(0.34)	(0.85)	(0.07)
Log distance	0.05***	0.07***				
	(22.17)	(11.88)				
Log product real GDP	−0.03***	−0.05***	−0.02	0.02*	−0.12***	−0.08***
	(30.50)	(19.75)	(1.29)	(1.71)	(7.94)	(3.21)
Log product real GDP/capita	−0.02***	−0.02***	−0.04***	−0.07***	−0.09***	−0.12***
	(9.93)	(3.94)	(3.61)	(4.74)	(8.11)	(7.80)
Common language	−0.03***	−0.03***				
	(6.25)	(4.25)				
Common land border	0.02**	−0.03**				
	(2.28)	(2.50)				
Free trade agreement (with Africa)	−0.06**	−0.08***	−0.02	−0.03*	−0.05	−0.03
	(5.95)	(5.50)	(1.34)	(1.80)	(0.99)	(0.56)
Number landlocked in the pair	0.01***	0.02***				
	(3.19)	(3.18)				
Number islands in the pair	0.02***	0.06***				
	(3.56)	(6.13)				
Log product of areas	0.00***	0.01***				
	(2.71)	(5.49)				
Common colonizer	−0.02***	0.00				
	(2.67)	(0.11)				
Current colony	−0.01	0.12***	0.04***	0.06***	0.70***	0.30*
	(0.66)	(3.59)	(2.71)	(2.69)	(4.21)	(1.72)

Table 7 ***(concluded)***

Ever colony	–0.02**	–0.04***				
	(2.19)	(2.75)				
Same nation	–0.02	–0.13***				
	(0.73)	(3.88)				
Observations	18,819	8,069	18,819	8,069	13,212	6,176
R^2 (within)			0.06	0.11		
R^2 (between)			0.11	0.01		
R^2 (overall)	0.17	0.16	0.11	0.01		
Hausman test (p-value)			0.01	0.00	0.00	0.00
Wald χ^2 (p-value)					0.00	0.00

Source: Authors' calculations.

Note: This table reports pooled ordinary least squares (OLS), fixed effects, and Poisson pseudomaximum likelihood (PML) estimation results for the effect of currency unions on trade stability using the world and Africa samples. The dependent variable is the coefficient of variation of log of real trade in columns 1–4 and of real trade in columns 5 and 6. The independent variables are currency union (dummy variable equal to 1 if money is interchangeable between the two countries at a 1:1 par for an extended period of time, and 0 otherwise); log of geographical distance between the trading partners; log of the product of real GDP of the trading partners; log of the product of real GDP per capita of the trading partners; common language (dummy variable equal to 1 if the two countries share a language, and 0 otherwise); common land border (dummy variable equal to 1 if the two countries share a border, and 0 otherwise); free trade agreement with Africa (a more comprehensive version of the variable free trade agreement that also takes into account free trade agreements with African countries); number of landlocked countries in the pair (for example, 0, 1, or 2); number of islands in the pair (for example, 0, 1, or 2); log of product of land areas of the two countries; common colonizer (dummy variable equal to 1 if the trading partners share the colonizer, and 0 otherwise); current colony (dummy variable equal to 1 if one country in the pair is colonized by the other country, and 0 otherwise); ever colony (dummy variable equal to 1 if one trading partner has ever been a colony of the other, and 0 otherwise); and same nation (dummy variable equal to 1 if both partners are part of the same nation, and 0 otherwise). The table reports robust t-statistics in parentheses of columns 1–4, and robust z-statistics in columns 5 and 6. *, **, *** denote significance at the 10, 5, and 1 percent levels, respectively.

interindustry.[28] The positive estimated effect of currency union on price comovement is relatively clear because countries that are members of currency unions avoid inflation and nominal exchange rate volatility that characterizes other regimes.

Last, we examine whether currency unions make trade more stable by reducing exchange rate volatility. Table 7 shows estimations of the gravity model with the coefficient of variation of log of real trade as dependent variable in columns 1–4, and of real trade in columns 5 and 6. The OLS results show that currency unions increase trade stability, with the marginal impact being slightly higher for Africa than for the world. However, this effect becomes statistically insignificant when country-pair specific effects are introduced in the model (columns 3–6). This indicates that the evidence of currency unions on trade stability is tentative and not robust.

Sensitivity analysis

Similar to the previous section, we check the robustness of the results presented in Tables 5–7, by conducting sensitivity checks of the specification. In particular, we estimate the dependent variable as the average of the logarithm of exports and imports, add quadratic terms for output and output per capita to control for possible sample nonlinearities, and add time effects. Estimates in Table 5 are robust to all the above sensitivity checks: after a country in the world (Africa) has been in a currency union for 10 years, trade increases by 14–22 percent (20–29 percent), and the diminishing returns of currency union membership are always significant. The conclusions of Tables 6 and 7 are also robust to changes in the specification.[29]

IV. Conclusions

This paper has provided some insights into several aspects of the performance of currency unions using an augmented version of the gravity model and focusing on two samples, the world and Africa. Our analysis confirms earlier results that the impact of currency unions is positive and significant on trade. We show that the trade-generating effect of currency unions is strong for Africa—and may be larger than that for the world—suggesting that Africa stands to benefit at least as much from currency union participation as the ROW. However, the magnitude of the currency union trade-generating effect seems to be sensitive to the inclusion of zero-trade observations as well as the estimation methodology employed. On average,

[28]As discussed in Alesina and Tenreyro (2002), increased interindustry trade may stimulate sectoral specialization and lead to less output comovement, but intra-industry trade is likely to increase their comovement.

[29]For trade stability, we estimate the model with various measures of stability, including the maximum absolute value, mean absolute value, and standard deviation of the residual from a conventional gravity equation of exports in levels. Results are available on request.

currency unions increase trade by factors of 1.2 and 1.4 in the world and Africa, respectively. Also, our results show that the duration of currency union membership matters: the longer the duration, the greater the benefits in terms-of-trade creation, albeit with some diminishing returns. These two effects combined suggest that although currency unions enhance trade by a factor of 1.2–1.4, the effect is not immediate: according to estimates, currency union participation doubles the level of trade in about 67 years for the world and 36 years for Africa. In addition, we find evidence that currency union participation increases price comovements among member countries, but has no significant effect on output comovements among members. Finally, the currency union effect on trade stability appears to be ambiguous.

Although our results indicate that several aspects of currency unions operate more or less the same in Africa as elsewhere, we would like to emphasize that the marginal effects and mechanisms of transmission may vary across the two samples. The methodology herein does not constitute an explicit investigation into how trade and its underlying determinants are connected, or the extent to which currency unions can promote growth and reduce poverty. Identifying the similarities and some of the differences across the samples is only the first step in investigating the dynamics of currency unions in Africa and may raise more questions than it answers.

Future research on the issue may explore the omitted factors that induce countries to join a currency union and to trade more, as suggested by Rose and van Wincoop (2001). Joining a currency union and opening up to the world are very often political decisions, so that such omitted factors may lie in a set of political and institutional variables. This is not a new realization: in recent papers on "deeper" determinants of economic growth, both institutions and trade openness determine a country's performance in the long run (see, for example, Acemoglu, Johnson, and Robinson, 2001; and Rodrik, Subramanian, and Trebbi, 2002). In the complicated web of relationships describing income, its determinants, and the linkages between the determinants, the interdependence of trade, and institutions is a recurring theme that is difficult to handle, not least because of the issues of causality and construction of proper instruments.

APPENDIX I

See Tables A1–A4 and Figure A1.

Table A1. Sample Data: Variable Definitions and Sources

Variable	Description	Source
Dependent variable		
X_{ijt}	The average value of real bilateral trade between i and j at time t	IMF's *Direction of Trade Statistics (DoT)*; Average of available values for export from a to b, export from b to a, import into a from b, import into b from a. Deflated by U.S. consumer price index
Explanatory variables		
Y_i	Real GDP	World Bank's *World Development Indicators*
Pop_i	Population	World Bank's *World Development Indicators*, *Penn World Tables 6.1*; Glick and Rose (2002) data used for 1948–50
D_{ij}	The distance between i and j	CIA's *World Factbook*; Great Circle Method used to calculate distance
$Lang_{ij}$	A binary variable that is unity if i and j have a common language	CIA's *World Factbook*
$Cont_{ij}$	A binary variable that is unity if i and j share a land border	CIA's *World Factbook*
Landl	The number of landlocked countries in the country pair (0, 1, or 2)	CIA's *World Factbook*
Island	The number of island nations in the country pair (0, 1, or 2)	CIA's *World Factbook*
Area	Land area of country n, $n = i, j$	CIA's *World Factbook*

Table A1 *(concluded)*

$ComCol_{ij}$	Binary variable that is unity if i and j were colonies after 1945 with the same colonizer	CIA's *World Factbook*
$CurCol_{ij}$	Binary variable that is unity if i and j are colonies at time t	CIA's *World Factbook*
$Colony_{ij}$	Binary variable that is unity if i colonized j or vice versa	CIA's *World Factbook*
$ComNat_{ij}$	Binary variable that is unity if i and j remained part of the same nation during the sample	CIA's *World Factbook*
FTA_{ij}	Binary variable that is unity if i and j belong to the same regional trade agreement	WTO publication [www.wto.org/english/tratop_e/region_e/region_e.htm]
CU_{ij}	Binary variable that is unity if i and j use the same currency at time t	Glick and Rose (2002). For 1998–2002: IMF's *Annual Report on Exchange Arrangements and Exchange Restrictions*
P_i	PPP of GDP	Penn World Tables 6.1; price level of GDP (P) is the purchasing power parity over GDP divided by the exchange rate and multiplied by 100
Tot_i	Terms of trade	IMF's *World Economic Outlook* 2004

Table A2. Summary Statistics

	World Sample (265,262 observations)		Africa Sample (100,597 observations)	
Variable	Mean	SD	Mean	SD
Log of real trade	10.020	3.246	8.813	2.951
Currency union	0.014	0.119	0.027	0.161
Log distance	8.167	0.809	8.175	0.661
Log product real GDP	47.974	2.662	47.075	2.193
Log product real GDP/capita	16.106	1.470	15.180	1.259
Common language	0.216	0.411	0.266	0.442
Common land border	0.030	0.170	0.029	0.169
Free trade agreement(with Africa)	0.034	0.180	0.045	0.207
Number landlocked in the pair	0.262	0.479	0.402	0.553
Number islands in the pair	0.338	0.537	0.250	0.470
Log product of areas	24.139	3.299	24.670	2.904
Common colonizer	0.099	0.298	0.155	0.362
Current colony	0.002	0.041	0.002	0.043
Ever colony	0.020	0.139	0.017	0.129
Same nation	0.000	0.016	0.000	0.017
Years in currency union	0.359	3.481	0.705	4.840
Comovements of outputs[1]	−0.075	0.030	−0.087	0.031
Comovements of prices[1]	−0.155	0.080	−0.163	0.070
Trade volatility[2]	0.181	0.258	0.224	0.294

[1]One value for country pair (6,992/7,000 observations for the world for comovements of outputs/prices and 3,793/3,796 for Africa samples, respectively).

[2]Two values for country pair (18,156 observations for the world and 7,845 for the Africa sample).

Table A3. Free Trade Agreements in the Sample

Regional Agreements	Members (date of entry)
Africa:	
ECOWAS: Economic Community of West African States	Benin (1975), Ghana (1975), Niger (1975), Burkina Faso (1975), Guinea (1975), Nigeria (1975), Cape Verde (1975), Liberia (1975), Senegal (1975), Côte d'Ivoire (1975), Mali (1975), Sierra Leone (1975), Gambia, The (1975), Mauritania (1975), Togo (1975)
COMESA: Common Market for Eastern and Southern Africa	Angola (1981), Malawi (1981), Tanzania (1981), Burundi (1981), Mauritius (1981), Uganda (1981), Comoros (1981), Mozambique (1981), Zambia (1981), Djibouti (1981), Rwanda (1981), Eritrea (1993), Namibia (1993), Zimbabwe (1981), Ethiopia (1981), Somalia (1981), Madagascar (1993), Kenya (1981), Sudan (1981), Seychelles (1993), Lesotho (1981), Swaziland (1981)

Table A3 *(concluded)*

Regional Agreements	Members (date of entry)
SADC: Southern African Development Community	Angola (1980), Lesotho (1980), Zimbabwe (1980), Botswana (1980), Swaziland (1980), Namibia (1990), Malawi (1980), Tanzania (1980), Mauritius (1995), Mozambique (1980), Zambia (1980), South Africa (1995)
CEMAC: Central African Economic and Monetary Community	Congo, Republic of (1962), Central African Republic (1962), Gabon (1962), Chad (1962–68, 1984), Cameroon (1962), Equatorial Guinea (1983)
WAEMU: West African Economic and Monetary Union	Benin (1984), Burkina Faso (1962), Côte d'Ivoire (1962), Mali (1962), Niger (1962), Senegal (1962), Mauritania (1962–94), Togo (1994)
Rest of the world:	
ASEAN	Brunei (1984), Cambodia (1999), Indonesia (1967), Lao PDR (1997), Malaysia (1967), Myanmar (1997), Philippines (1967), Singapore (1967), Thailand (1967), Vietnam (1995)
EU	Austria (1995), Belgium (1958), Denmark (1973), Finland (1995), France (1958), Germany (1958), Greece (1981), Luxembourg (1958), Ireland (1973), Italy (1958), Netherlands (1958), Portugal (1986), Spain (1986), Sweden (1995), United Kingdom (1973)
United States-Israel	United States (1985), Israel (1985)
NAFTA	Canada (1989), United States (1989), Mexico (1994)
CARICOM	Antigua and Barbuda (1974), Bahamas (1983), Barbados (1973), Belize (1974), Dominica (1974), Guyana (1973), Grenada (1974), Jamaica (1973), Montserrat (1974), Saint Kitts and Nevis (1974), Saint Lucia (1974), Saint Vincent and the Grenadines (1974), Suriname (1995), Trinidad and Tobago (1973), Haiti (2002)
PATCRA	Australia (1977), Papua New Guinea (1977)
ANZCERTA	Australia (1983), New Zealand (1983)
CACM	Costa Rica (1963), El Salvador (1961), Guatemala (1961), Honduras (1961), Nicaragua (1961)
SPARTECA	Australia (1981), New Zealand (1981), Cook Islands (1981), Micronesia, Federated States of (1986), Fiji (1981), Kiribati (1981), Marshall Islands (1981), Nauru (1981), Niue (1981), Papua New Guinea (1981), Solomon Islands (1981), Tonga (1981), Tuvalu (1981), Vanuatu (1981), Samoa (1981)
MERCOSUR	Argentina (1991), Brazil (1991), Paraguay (1991), Uruguay (1991), Chile (1996), Bolivia (1996)

Table A4. Currency Unions in the Sample

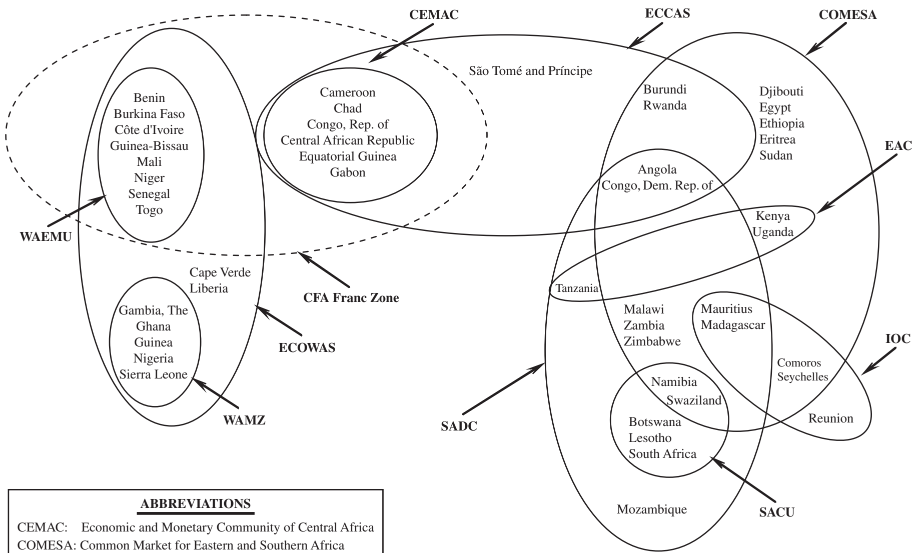
Currency Union Members		End	Currency Union Members		End
Antigua and Barbuda	Barbados	1975	Central African Republic	Congo, Republic of	Ongoing
Antigua and Barbuda	Dominica	Ongoing	Central African Republic	Côte d'Ivoire	Ongoing
Antigua and Barbuda	Grenada	Ongoing	Central African Republic	Equatorial Guinea	Ongoing
Antigua and Barbuda	Guyana	1971	Central African Republic	Gabon	Ongoing
Antigua and Barbuda	Montserrat	Ongoing	Central African Republic	Guinea	1969
Antigua and Barbuda	Saint Kitts and Nevis	Ongoing	Central African Republic	Guinea-Bissau	Ongoing
Antigua and Barbuda	Saint Lucia	Ongoing	Central African Republic	Madagascar	1982
Antigua and Barbuda	Saint Vincent and the Grenadines	Ongoing	Central African Republic	Mali	Ongoing
Antigua and Barbuda	Trinidad and Tobago	1976	Central African Republic	Mauritania	1974
Aruba	Netherlands Antilles	Ongoing	Central African Republic	Niger	Ongoing
Aruba	Suriname	1994	Central African Republic	Reunion	1976
Australia	Kiribati	Ongoing	Central African Republic	Senegal	Ongoing
Australia	Nauru	Ongoing	Central African Republic	Togo	Ongoing
Australia	Solomon Islands	1979	Chad	Benin	Ongoing
Australia	Tonga	1991	Chad	Burkina Faso	Ongoing
Australia	Tuvalu	Ongoing	Chad	Comoros	1994
Bangladesh	India	1974	Chad	Congo, Republic of	Ongoing
Barbados	Dominica	1975	Chad	Côte d'Ivoire	Ongoing
Barbados	Grenada	1975	Chad	Equatorial Guinea	Ongoing
Barbados	Guyana	1971	Chad	Gabon	Ongoing
Barbados	Montserrat	1975	Chad	Guinea	1969
Barbados	Saint Kitts and Nevis	1975	Chad	Guinea-Bissau	Ongoing
Barbados	Saint Lucia	1975	Chad	Madagascar	1982
Barbados	Saint Vincent and the Grenadines	1975	Chad	Mali	Ongoing
Barbados	Trinidad and Tobago	1975	Chad	Mauritania	1974
Belgium	Burundi	1964	Chad	Niger	Ongoing
Belgium	Congo, Democratic Republic of	1961	Chad	Reunion	1976

Table A4 *(concluded)*

Belgium	Rwanda	1966	Chad	Senegal	Ongoing
Belgium-Luxembourg	Burundi	1964	Chad	Togo	Ongoing
Belgium-Luxembourg	Congo, Democratic Republic of	1961	Comoros	Benin	1994
Belgium-Luxembourg	Rwanda	1966	Comoros	Burkina Faso	1994
Benin	Burkina Faso	Ongoing	Comoros	Congo, Republic of	1994
Benin	Côte d'Ivoire	Ongoing	Comoros	Côte d'Ivoire	1994
Benin	Equatorial Guinea	Ongoing	Comoros	Equatorial Guinea	1994
Benin	Gabon	Ongoing	Comoros	Gabon	1994
Benin	Guinea	1969	Comoros	Guinea	1969
Benin	Guinea-Bissau	Ongoing	Comoros	Madagascar	1982
Benin	Madagascar	1982	Comoros	Mali	1994
Benin	Mali	Ongoing	Comoros	Mauritania	1974
Benin	Mauritania	1974	Comoros	Niger	1994
Benin	Niger	Ongoing	Comoros	Reunion	1976
Benin	Reunion	1976	Comoros	Senegal	1994
Benin	Senegal	Ongoing	Comoros	Togo	1994
Benin	Togo	Ongoing	Congo, Republic of	Benin	Ongoing
Bhutan	India	Ongoing	Congo, Republic of	Burkina Faso	Ongoing
Bhutan	Pakistan	1966	Congo, Republic of	Côte d'Ivoire	Ongoing
Botswana	Lesotho	1977	Congo, Republic of	Equatorial Guinea	Ongoing
Botswana	Swaziland	1977	Congo, Republic of	Gabon	Ongoing
Brunei	Malaysia	1971	Congo, Republic of	Guinea	1969
Brunei	Singapore	Ongoing	Congo, Republic of	Guinea-Bissau	Ongoing
Myanmar	India	1966	Congo, Republic of	Madagascar	1982
Myanmar	Pakistan	1971	Congo, Republic of	Mali	Ongoing
Cameroon	Benin	Ongoing	Congo, Republic of	Mauritania	1974
Cameroon	Burkina Faso	Ongoing	Congo, Republic of	Niger	Ongoing
Cameroon	Central African Republic	Ongoing	Congo, Republic of	Reunion	1976
Cameroon	Chad	Ongoing	Congo, Republic of	Senegal	Ongoing

Cameroon	Comoros	1994	Congo, Republic of	Togo	Ongoing
Cameroon	Congo, Republic of	Ongoing	Côte d'Ivoire	Burkina Faso	Ongoing
Cameroon	Côte d'Ivoire	Ongoing	Côte d'Ivoire	Madagascar	1982
Cameroon	Equatorial Guinea	Ongoing	Côte d'Ivoire	Mali	Ongoing
Cameroon	Gabon	Ongoing	Côte d'Ivoire	Mauritania	1974
Cameroon	Guinea	1969	Côte d'Ivoire	Niger	Ongoing
Cameroon	Guinea-Bissau	Ongoing	Côte d'Ivoire	Reunion	1976
Cameroon	Madagascar	1982	Côte d'Ivoire	Senegal	Ongoing
Cameroon	Mali	Ongoing	Côte d'Ivoire	Togo	Ongoing
Cameroon	Mauritania	1974	Denmark	Faroe Islands	Ongoing
Cameroon	Niger	Ongoing	Denmark	Greenland	Ongoing
Cameroon	Reunion	1976	Djibouti	Benin	1949
Cameroon	Senegal	Ongoing	Djibouti	Burkina Faso	1949
Cameroon	Togo	Ongoing	Djibouti	Cameroon	1949
Central African Republic	Benin	Ongoing	Djibouti	Central African Republic	1949
Central African Republic	Burkina Faso	Ongoing	Djibouti	Chad	1949
Central African Republic	Chad	Ongoing	Djibouti	Comoros	1949
Central African Republic	Comoros	1994	Djibouti	Congo, Republic of	1949

Figure A1. Main African Regional and Subregional Economic Integration Arrangements (as of August 2005)



ABBREVIATIONS

CEMAC: Economic and Monetary Community of Central Africa
COMESA: Common Market for Eastern and Southern Africa
EAC : East African Community
ECCAS: Economic Community of Central African States
ECOWAS: Economic Community of West African States
IOC: Indian Ocean Commission
SACU: Southern African Customs Union
SADC: Southern African Development Community
WAEMU: West African Economic and Monetary Union
WAMZ: West African Monetary Zone

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International Competitiveness of the Mediterranean Quartet: A Heterogeneous-Product Approach

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The real effective exchange rate (REER) is the most commonly used measure for assessing international competitiveness. This paper develops a methodology to estimate the REER that incorporates two distinctive elements that are not considered in the current literature and applies it to the Mediterranean quartet (MQ) of Greece, Italy, Portugal, and Spain, whose common pattern of real appreciation has created concern in policy and academic circles. The two elements that this paper adds to the existing literature are (1) product heterogeneity when identifying each country's international competitors and their weights, and (2) a comprehensive treatment of services exports. Our refined measure suggests a modest reduction in the observed REER gap between the MQ countries and the other euro area countries. In particular, considering product heterogeneity and services exports implies a lower real appreciation from 1998 to 2006 on the order of 2 to 3 percent for all MQ countries. These are difference-in-difference estimates relative to the results obtained for the rest of the euro area countries using the same methodology. [JEL F10, F30]
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This paper develops a methodology to estimate the real effective exchange rate (REER) that incorporates two distinctive elements not accounted for in the current literature: (1) product heterogeneity when determining international competitors and their weights, which allows us to identify countries' direct international competitors more accurately, and (2) a comprehensive treatment of services exports, which allows us to provide a complete view of international competitiveness encompassing the entire export sector.

We apply this methodology to reexamine the evolution of the REER of the Mediterranean quartet (MQ) of Greece, Italy, Portugal, and Spain, and particularly, the evolution of their REER gap with the other euro area members. This case motivates our analysis as the common pattern of real appreciation observed in the MQ countries has created concern in policy and academic circles (European Commission, 2006; Bini-Smaghi, 2007; Roubini, 2007; and Papademos, 2007). Particular attention has been given to the fact that this pattern diverges from the average real depreciation observed in the rest of the euro area (see Figure 1). It is argued that this real appreciation is associated with a loss of international competitiveness in the MQ and that it could lead to a persistent period of slow growth, which has already materialized in the cases of Portugal and Italy (Blanchard, 2006a and 2006b).[1]

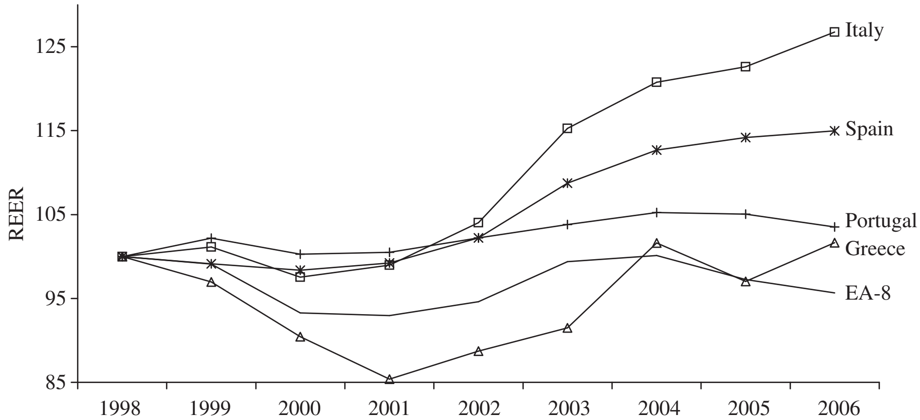
In short, the REER is an aggregated measure of cost competitiveness between countries. It tracks the evolution of cost competitiveness of a particular country with respect to a weighted average of all other countries in the world.[2] The methodologies available to calculate the REER have been constantly improving in recent decades as they have been incorporating more realistic assumptions. Table 1 summarizes the existing literature and highlights the approach taken in each study to address the key elements of the REER analysis, that is, the approach used to calculate the importance or weight of each other country and the price used to measure cost competitiveness. Bayoumi, Jayanthi, and Lee (2005, 2006) is the most comprehensive methodology currently available, which includes the latest development in the literature.

Determining the weights by identifying the degree to which countries compete in international markets, as opposed to weighting by trade partnership, is one of the most important characteristics that distinguishes the most up-to-date REER estimations. To illustrate the importance of this

[1]See, for example, "The Quest for Prosperity," Special Report, *The Economist,* March 15, 2007, which stated: "In particular, the Mediterranean quartet of Italy, Spain, Portugal and Greece has suffered a huge loss of competitiveness in a relatively short time… . This loss is reflected in colossal current-account deficits … or pitifully slow growth."

[2]See Agenor (1995); Catao (2007); Chinn (2006); Fung and Klau (2006); Marsh and Tokarick (1996); Neary (2006); and Rogoff (1996) for further references to the concept of REER.

Figure 1. Real Appreciation in the Mediterranean Quartet vs. the Rest of the Euro Area



Source: OECD.

Note: Unit-labor-cost-based real effective exchange rate index (REER) defined à la IMF (higher means more appreciated); the base year is 1998. EA-8 refers to Austria, Belgium, France, Finland, Germany, Ireland, Luxembourg, and Netherlands. The REER for the EA-8 is estimated aggregating each country's REER and weighting by their total exports of goods. All countries adopted the euro on January 1, 1999, except for Greece, which adopted the euro on January 1, 2001.

feature, consider the extreme case of two countries, A and B, that export mostly to a third country C and have nil bilateral trade between them. If the weight of country B in the calculation of country A's REER is based on trade partnership, then changes in the exchange rate of country B will not alter the REER of country A. This is not a desirable feature of an index of relative cost competitiveness, because countries A and B compete when exporting to country C and exchange rate movements in either country clearly affect the relative cost competitiveness of the other one. The interrelationship between the cost competitiveness of countries A and of country B is better captured by the REER if the weights are based on a measure of how much these two countries compete in international markets.

With respect to the method used to identify international competitors and their weights, the existing literature considers that two countries are international competitors if they both sell products in the same country, that is, in a market defined as a single aggregated sector comprising a representative product category—which we refer to as the representative-product approach (RPA). As a result, the RPA assumes implicitly that all exporters compete with each other in the destination country. In contrast, we take a more micro-based approach that considers product heterogeneity when defining markets and identifying international competitors and their weights. For each product type that we consider, we identify international competitors as competitors competing in the market for that product type in the destination country. This allows us to analyze relative cost competitiveness

Table 1. Literature Overview

	Importance of Other Countries (Definition of Markets)				
Reference	Partners/competitors	Product and market definition (goods)	Services	Local consumption of local production	Price Competitiveness Measure
Bank for International Settlements (Fung and Klau, 2006)	Competitors	Representative-product approach. All manufacturing goods (SITC Rev. 3 5–8) are treated as one identical good; nonmanufacturing goods are not considered. Markets are defined at the country level.	Services are not included in the analysis.	Total manufacturing output.	Aggregate prices. Consumer price index (CPI) and/or manufacturing unit labor cost are used to measure relative cost competitiveness.
Bank of Japan (2007)	Partners	Total exports of goods.	Services are not included in the analysis.	Does not apply.	Aggregate prices. CPI and/or manufacturing unit labor cost are used to measure relative cost competitiveness.
European Central Bank (Buldorini, Makrydakis, and Thimann, 2002)	Competitors	Representative-product approach. All manufacturing goods (SITC Rev. 3 5–8) are treated as one identical good; nonmanufacturing goods are not considered. Markets are defined at the country level.	Services are not included in the analysis.	Manufacturing output for domestic use.	Aggregate price. CPI, manufacturing unit labor cost, PPI and/or wholesale prices are used to measure relative cost competitiveness.
Federal Reserve Board (Loretan, 2005)	Average between competitors and partners	Representative-product approach. All goods are treated as one identical good (except for oil, gold, and military items, which are not considered).	Services are not included in the analysis.	Local production is not considered in the analysis.	Aggregate prices. CPI and/or manufacturing unit labor cost are used to measure relative cost competitiveness.
International Monetary Fund (Bayoumi, Jayanthi, and Lee, 2005)	Competitors	Representative-product approach. All manufacturing goods (SITC Rev. 3 5–8, excl. 68) are treated as one identical good. Commodities are disaggregated into 20 categories at 2-dig. SITC Rev.3 level. Markets	Services are considered in the analysis, but assumed to have the same trade pattern as the	Manufacturing output for domestic use.	Aggregate prices. CPI and/or manufacturing unit labor cost are used to measure relative cost competitiveness.

		are defined at the country level for manufactured goods and at the global level for commodities (global goods).	observed pattern for manufacturing goods. Tourism is treated separately only for a subset of countries.		
Organization for Economic Cooperation and Development (Durand, Simonm, and Webb, 1992; Durand, Madaschi, and Terribile, 1998)	Competitors	Representative-product approach. All manufacturing goods (SITC Rev. 3 5–9) are treated as one identical good. Markets are defined at the country level for individual OECD countries and at the level of country aggregates for six non-OECD country groups.	Services are not included in the analysis.	Total manufacturing output.	Aggregate price. CPI and/or manufacturing unit labor cost are used to measure relative cost competitiveness.
Current paper (Bennett and Zarnic)	Competitors	Heterogeneous-product approach. All goods are treated disaggregately at 4-digit ISIC Rev. 3 level. Markets are defined at the country level for all nonglobal goods and at the global level for global goods.	Services are treated disaggregately at 2-digit ISIC Rev. 3 level. Markets are defined at the country level for all services.	Total output for domestic use (at industry-level).	Aggregate and disaggregate prices. CPI, Manufacturing unit labor cost, and sectoral unit labor cost data for goods (1- and 2-digit ISIC Rev. 3) are used to measure relative cost competitiveness.

Note: The representative-product approach refers to the common approach used in the literature, which assumes that all (or most) exporting goods compete with each other in the international markets. The heterogeneous-product approach refers to the approach used in this paper, which assumes that exporting goods compete with each other within disaggregated categories of goods and of services. The following examples illustrate the difference between looking at partners, looking at competitors assuming the representative-product approach, and looking at competitors assuming the heterogeneous-product approach. First, note that two countries, A and B, could compete when exporting the same good to a third country, C, while trade between countries A and B could be either high or low. This suggests that the degree of trade between countries does not necessarily reflect the degree to which countries compete in international markets. Second, assume that country A exports textiles to country C, while country B exports cars to country C. Focusing on competitors at an aggregate levelat the manufacturing sector for example, as it is the most common case in the literature, would suggest that countries A and B compete in market C, even though exporters of cars are not necessary relevant competitors of exporters of textiles. Furthermore, the homogeneous-product approach would imply that all manufacturing goods produced in country C, are competitors of exporters to country C, regardless of the type of good that is produced in and exported to country C. See Chinn (2006) and Fung and Klau (2006) for a more detailed exposition of the different methodologies available in the literature.

at disaggregated markets according to the type of product and destination country. We aggregate these market-level REER indices to obtain a country-level REER—which we refer to as the heterogeneous-product approach (HPA). In principle, our methodology can be applied to alternative definitions of market. Based on data availability, however, we define markets at 4-digit ISIC category of goods and at 2-digit ISIC category of services.[3]

The HPA identifies more precisely a country's direct international competitors, and thus, their weights. This feature operates at two levels: first, with respect to other exporters, and second, with respect to local producers at the destination of exports. To illustrate the differences, assume that country A exports textiles to country C, but country B exports cars to country C. The RPA focuses on competitors at an aggregate level—at the manufacturing sector for example, the most common case in the literature—suggesting that countries A and B compete in market C, even though exporters of cars are not necessarily competitors of textile exporters. Furthermore, the RPA would imply that all manufacturing goods produced in country C are competitors of exporters to country C, regardless of the type of good that is produced in, and exported to, country C. In contrast, the disaggregated view of the HPA would not consider countries A and B as competitors in this example and would consider only textile producers in country C as competitors of country A.

With respect to services exports, our approach provides a comprehensive view of relative cost competitiveness by incorporating information about exports of services as well as exports of goods. Services exports have become increasingly important and represent 65 percent of total exports in Greece, 19 percent in Italy, 27 percent in Portugal, and 31 percent in Spain. As with the case of goods, we identify competitors in the destination market at disaggregated categories of services. Unfortunately, the available data on disaggregated bilateral trade in services are not as complete as the data for trade in goods, and therefore, our estimates of the REER in services are restricted to the available sample of trade flows. The average coverage of bilateral trade ranges from 89 percent of total services exports for Greece to 59 percent for Spain. The coverage for goods is above 90 percent for all MQ countries.

[3]We initially attempted to base our definition of international competitors on the degree of substitution between goods. We intended to include producers of other 4-digit level industries as competitors, weighting their importance by the degree of substitution between the corresponding goods, as measured by the cross elasticity of substitution. Available volume indices, however, present important measurement error. This affects the estimation of price indices, necessary for the estimation of the cross elasticities, and therefore, the whole methodology would have resulted in a significant increase in the variance of the REER estimates. Also, the presence of monopolistic competition at different intensities across industries and countries suggests that the estimation of cross elasticities is subject to potential identification problems. We have, therefore, assumed the simplifying assumption that the structure of international competitors within 4-digit sectors provides a good representation of the structure of competitors that exporters within those sectors face. The same is assumed for services within the 2-digit industry-level.

Our results suggest a modest reduction in the observed REER gap between the MQ countries and the other members of the euro area. Allowing for product heterogeneity (HPA) and services exports implies, compared with the standard results obtained under the RPA, a lower real appreciation from 1998 to 2006 on the order of 2 to 3 percent for all MQ countries—2 percent for Greece, 2.8 percent for Italy, 2.4 percent for Portugal, and 2.3 percent for Spain. These figures are based on difference-in-difference estimates that control for the results obtained for the rest of the euro area countries using the same methodology. As a robustness check, we also show that our results obtained under the RPA are consistent with the ones reported using the methodology presented in Bayoumi, Jayanthi, and Lee (2005 and 2006), the closest methodology to the one presented in this paper that uses RPA.

The above results are based on a single cost measure at the country-level as used in the current literature, namely the unit labor cost (ULC). To the extent that wages and productivity growth vary across exporting sectors, differentiated cost measures at the sector-level would yield a more accurate picture of international competitiveness. We explore this avenue and compute the REER using differentiated ULC measures at the 2-digit level. Unfortunately, the sample of countries with differentiated ULC series (28) is more restricted than the sample with aggregated ULC series (38), particularly regarding Asian countries. Also, the available time span is one year shorter than in the aggregate ULC data. Moreover, the ULC series at the industry-level may be more volatile because of its disaggregated nature, which could be magnifying some of the known problems of the ULC indicator as a measure of cost competitiveness (see footnote 2). Therefore, the results based on differentiated ULC should be taken with caution, given the data limitations detailed above, and should be read as an exploratory effort to determine the effect of differentiated cost measures on the REER.

The existing literature determines the weights used in the estimation of the REER using the most comprehensive bilateral trade data available. These data report the final value of the exported product rather than the value added embodied in it that is domestically produced—that is, the final value net of imported intermediate inputs. As Hummels, Jun, and Yi (2001) and Yi (2003) have pointed out, the latter is a more desirable measure of the economic importance of exports as it accounts for vertical specialization by netting imported intermediate inputs. Unfortunately, and given the nature of the available data, the estimations presented in this paper are not free of this limitation. Having said that, the HPA addresses part of the increase in the variance of the REER estimator due to the proxy used for exports—imported intermediate inputs overestimate the importance of local cost for all countries and not only for the one under analysis. To the extent that vertical specialization for each type of product is relatively similar across competitors within 4-digit level, our methodology is computing a REER that accounts for the difference in the value added of exported products across industries when selecting the relevant set of competitors. The RPA does not account for this factor given its implicit assumption that all exports products compete with each other in the destination country.

The micro-based methodology proposed in this study also allows for a quantitative assessment of each country's profile of competitors. Such evidence provides information about the exposure of each country to its key competitors around the world; for example, the exposure to emerging competitors like China, which has shown a strong pattern of productivity and trade growth, or the exposure to countries facing significant changes in their cost structure, such as the wage moderation observed recently in Germany and the depreciation of the nominal exchange rate observed in the United States during the recent years. Our findings indicate that the bulk of competition for the MQ still comes from the advanced economies, especially from the euro area—Spain and Portugal are more exposed to euro area competition, and therefore, less exposed to changes in the value of the euro. Nonetheless, there has been a change in the goods sector, as emerging economies have grown in importance since the late 1990s, particularly China in both high- and low-technology sectors.

I. Methodology

This section presents the methodology used to estimate the evolution of the REER under the HPA. The first subsection develops a generic framework to aggregate at the country-level the relative cost competitiveness dynamics observed at the country-industry-level. This framework allows us to incorporate into the analysis elements such as global goods—goods whose markets are defined at the world-level rather than at the country-level—and local consumption of local production—the competition that local producers represent at exports' destination.

A Generic Approach to Aggregate Relative Cost Competitiveness

We construct our index of relative cost competitiveness between country i and the countries in the rest of the world (ROW), denoted by $R_{i,t}$, using a geometrical Laspeyres index and the chain link methodology. The subscripts g, d, and t refer to the type of product (including both goods and services), destination (country), and time (year), respectively.

The evolution of the $R_{i,t}$ index is defined in equation (1), where T denotes the base year, and $\Delta\Omega_{i,t}$ denotes the (natural logarithmic) change of the relative cost competitiveness between country i and the ROW between period t and $t-1$.[4]

$$\ln R_{i,t} = \ln R_{i,T} + \sum_{j=T+1}^{t} \Delta\Omega_{i,j}. \tag{1}$$

The change in relative cost competitiveness at the country-level, $\Delta\Omega_{i,t}$, is constructed as the weighted average of the change in the relative cost

[4]Throughout the paper, the notation Δ denotes the natural logarithmic difference between t and $t-1$of the corresponding variable.

competitiveness between country i and the ROW in each market defined by the g,d pair (equation (2)). The change in the relative cost competitiveness in market g,d is denoted by $\Delta\theta_{i,g,d,t}$. The weights are given by the importance of each market g,d in country i's exports and are denoted by $\beta_{i,g,d,t-1}$. The term $\beta_{i,g,d,t-1}$ is computed as the share of country i's total exports represented by its exports of product g to destination d (equation (3)).

$$\Delta\Omega_{i,t} = \sum_{\forall(g,d)} \beta_{i,g,d,t-1} \cdot \Delta\theta_{i,g,d,t}, \tag{2}$$

$$\beta_{i,g,d,t-1} = \frac{S_{i,g,d,t-1}}{\sum_{\forall(g,d)} S_{i,g,d,t-1}}, \tag{3}$$

where $S_{i,g,d,t}$ represents sales by country i of product g in destination d; that is, sales by country i in market g,d.

The change in relative cost competitiveness at the market-level, $\Delta\theta_{i,g,d,t}$, is given by equation (4)—defined à la IMF, that is, a higher number means more appreciated. It is constructed as the difference between country i's cost change and a weighted average of the same cost change observed in all other countries competing in market g,d. The variables $P_{i,g,t}$ and $P_{c,g,t}$ represent the cost variable used to estimate cost competitiveness, are specific to each industry in each country, and are expressed in the local currency. The subscript c is used to denote a competitor of the country under study—which, as already mentioned, is denoted by the subscript i. The variable $E_{i,c,t}$ represents the exchange rate between country i and country c defined as units of country i's currency per unit of country c's currency.

$$\Delta\theta_{i,g,d,t} = \Delta P_{i,g,t} - \sum_{\forall c \neq i} \alpha_{i,c,g,d,t-1} \cdot (\Delta P_{c,g,t} + \Delta E_{c,i,t}). \tag{4}$$

The weight $\alpha_{i,c,g,d,t-1}$ is given by the importance of each country (c) as a competitor of country i in market g,d. We relate this weight to the market participation (share) of country c in market g,d, denoted by $\gamma_{c,g,d,t}$.[5]

$$\gamma_{c,g,d,t-1} = \frac{S_{c,g,d,t-1}}{\sum_{\forall c} S_{c,g,d,t-1}}. \tag{5}$$

Defining $\alpha_{i,c,g,d,t-1} = \gamma_{c,g,d,t-1}$, however, implies that the sum of the weights of all competitors is less than one, that is, $\Sigma_{\forall(c\neq i,g,d)}\, \beta_{i,g,d,t-1} \cdot \alpha_{i,c,g,d,t-1} < 1$, because country i is not considered as a competitor of itself in equation (4). Alternatively, one could add country i in the latter sum to make it equal to one. However, doing so would violate an important property that an estimator of the REER should have: ceteris paribus, if all competitors of

[5]The methodology presented in this paper can also be applied in the ideal case of having a database with bilateral trade of value added, that is, total export value net of imported intermediate inputs. In this case, $\beta_{i,g,d,t-1}$ and $\gamma_{c,g,d,t-1}$ would be computed based on value added figures rather than on final sale figures.

country i depreciate their currency by 10 percent, then, country i's REER appreciates by 10 percent—the 10 percent property, for short. This property is violated if country i is added to the summation because the total mass of all countries excluding country i is less than one. Another alternative would be to exclude country i when computing each country's market share $\gamma_{c,g,d,t-1}$. However, this solution creates a bias, overstating the importance of small competitors relative to the importance of big competitors. To illustrate this point, assume a foreign market with two equally large competitors plus an exporter from country i. Excluding country i from the computation of the market share would imply that the other two competitors would represent 50 percent of the market, regardless of their actual importance as competitors of country i in that market.

We propose an alternative methodology to measure the importance of each country in each market as a competitor of country i, that is, $\beta_{i,g,d,t-1} \cdot \alpha_{i,c,g,d,t-1}$. We rescale the importance of each competitor in each market based on the relative importance that this competitor has in that market vis-à-vis the importance of all other competitors in all other markets (equation (6)). Our adjusted measure satisfies the condition that the weights of all competitors sum up to one, that is $\Sigma_{\forall(c\neq i,g,d)} \beta_{i,g,d,t-1} \cdot \alpha_{i,c,g,d,t-1} = 1$, satisfies the 10 percent property, and does not over or understate the relative importance of each competitor vis-à-vis all other competitors of country i.

$$\beta_{i,g,d,t-1} \cdot \alpha_{i,c,g,d,t-1} = \frac{\beta_{i,g,d,t-1} \cdot \gamma_{c,g,d,t-1}}{\sum_{\forall(c\neq i,g,d)} \beta_{i,g,d,t-1} \cdot \gamma_{c,g,d,t-1}}. \tag{6}$$

Global Goods, Local Consumption of Local Production, Representative-Product Approach, and Aggregated Cost Measures

The more comprehensive REER estimates available in the current literature incorporate two important elements into the analysis (see Table 1 for details): global goods and local consumption of local production. The former refers to goods that can be characterized as commodities (for example, copper) and for which a more appropriate definition of market is at the world-level rather than at the country-level. Regarding consumption of local production, this refers to the competition that local producers represent at exports' destination, and it is proxied by the difference between local production of a good and the exports of that good from that destination to the rest of the world.

Our generic approach can be used to consider these two elements. First, we define an artificial additional destination d that will not correspond to a particular country but to the world. Therefore, all goods g that are considered global goods (see next section for details) are assumed to compete in the market $(g, d = world)$. Second, we incorporate local consumption of local production by defining the case $d = c$, which refers to competitor c competing in the market g, $d = c$.

We also estimate the REER under the RPA to study the marginal effect of the HPA. The methodology presented in the previous section can easily consider this case as well by redefining all goods into a single good $g=\bar{g}$. Following Bayoumi, Jayanthi, and Lee (2005, 2006), we treat global goods separately under the RPA.

Finally, the lack of data limits the extent to which differentiated cost measures by sector can be modeled. The generic approach can be adjusted to consider aggregated measures for the corresponding subsectors within the defined aggregation level by simply defining $\Delta P_{c,g,t}=\Delta P_{c,\hat{g},t}\ \forall_g$ s.t. $g\subset\hat{g}$, where $\hat{g}$ refers to an aggregated sector.

II. Data

Goods and Services

Bilateral trade data for goods are compiled from the United Nations Commodity Trade Statistics Database (COMTRADE). The data include 144 different activity classes of goods (4-digit ISIC Rev.3) across 200 countries over the period 1998–2005. We explicitly consider only the exports of domestically produced goods and exclude the exports of foreign goods to ensure that re-export goods are not considered as additional exports when assessing international competitiveness.[6]

Bilateral trade data for services are compiled from the OECD Statistics of International Trade in Services. The data include nine categories of services, according to the Extended Balance of Payments Services Classification (EBOPS), across 100 countries over the period 1999–2004. The structure of competitors in 1998 and 2005 is extrapolated from the information available for 1999 and 2004, respectively. The average coverage of bilateral trade data for the MQ ranges from 86 percent of total exports of services for Greece to 59 percent for Spain. The same figures for goods are all above 90 percent.

Disaggregated local production series are needed to estimate local consumption of local production. Obtaining consistent and complete production data at 4-digit level is a challenging endeavor because available databases present significant differences in product and time coverage across countries. We approach this difficulty by combining various databases and generating (rough) estimates where possible. Our main source is the United Nations Industrial Demand-Supply Balance database (IDSB), which contains data at the 4-digit level of ISIC Rev.3 classification, which comprises 127 manufacturing commodities and 78 countries. We extend

[6]The lack of comparable data on value added of bilateral trade flows prevent us from basing our measure of international competitiveness on value added of exports. We try to address this issue by accounting for re-exports and defining detailed markets for traded products. By doing so, we address part of the increase in the variance of the REER estimator due to the fact that the share of value added of exports may differ across products. Ideally, one would like to have a database with value of exports net of imported intermediate inputs.

the IDSB database using (1) the annual growth rates of output reported in Eurostat's Annual Enterprise Statistics database (4-digit NACE Rev.1.1 production data)[7]; (2) the observed ratio between sectoral output and aggregate manufacturing output in Eurostat's Annual Enterprise Statistics database; and (3) the observed growth rate of output production in total manufacturing to extend the series for a maximum window of three years—if output production growth in total manufacturing is not available, we use value added growth in total manufacturing. Finally, we consider only series with complete data for the period 1998 to 2005 (original or estimated) to avoid biases/changes in the REER measures due to truncation of series unrelated to changes in relative cost competitiveness.

With respect to production of services, we use the EU KLEMS Growth and Productivity Accounts database. EU KLEMS database reports data for EU25 countries, Australia, Japan, and the United States until 2005. Production data for royalties and license fees are not considered because the match between EBOPS and ISIC Rev.3 classifications has many shared codes that make it impossible to build a consistent correspondence. Time coverage differs across country and sectors, although to a much lower extent than in the case for goods. We extend the series in the same fashion as we do for goods, using production data from the OECD-STAN database and GDP growth rates. We consider only series with complete time series data (original or estimated).[8]

The coverage for local production of goods, defined as the share of exports represented by the destination markets for which we can construct data on local consumption of local production, ranges from 50 percent for Greece to 70 percent for Portugal (60 percent for Italy and 66 percent for Spain). The coverage for local production of services, defined similarly, is 85 percent for Greece, 70 percent for Italy, 83 percent for Portugal, and 94 percent for Spain. These last figures are not necessarily comparable with the figures for local production of goods because their computation is based on the available bilateral trade data, which has a lower coverage for services than for goods.

We focus on REER measures that proxy cost competitiveness using the ULC, as opposed to consumer and producer price indices (CPI and PPI, respectively). The latter variables have the advantage of being available for most countries around the world. However, the ULC measure seems to be more appropriate because it considers changes in productivity. The ULC measure allows us to incorporate, albeit not perfectly, important dynamics

[7]Eurostat's Annual Enterprise Statistics database has good coverage of production data, but includes only members of European Union. With respect to the correspondence used, we consider data that (1) corresponds to only one type of product at the 4-digit ISIC Rev.3 level and (2) does not share the image with data points that correspond to more than one type of product at the 4-digit ISIC Rev.3 level.

[8]The combined information on bilateral trade and consumption of local production accounts for more than 36 million observations.

when considering cost competitiveness, such as the Balassa-Samuelson effect and the effect of innovation and structural reforms across countries.[9]

Manufacturing ULC is used as the aggregate ULC measure at the country-level. The data are obtained directly from the OECD Analytic Database and WEO database and are available for 38 countries. Industry-level ULC data at 2-digit level is computed using the EU KLEMS database, which covers 28 countries until 2005. Table A1 in Appendix details the different samples. The industry-level ULC is computed as the ratio of the compensation of employees per hour worked to real gross value added per hour worked. The compensation of employees per hour worked is obtained from the ratio of compensation of employees to total hours worked by employees.

Most of the disaggregated ULC series are complete for the countries covered by the EU KLEMS database, although data for 2006 are not available. On average, 95 percent of manufacturing sectors and 90 percent of all sectors have complete time-series for the period 1995–2005. Disaggregated ULC series with incomplete data are not considered and are replaced by the country's ULC series for the manufacturing sector computed from the EU KLEMS database. We replace them to avoid biases/changes in the REER measures due to truncations of series unrelated to changes in relative cost competitiveness.

The annual average nominal exchange rates are obtained from the IFS database.

Global Goods

We refer to globally traded goods as those goods whose prices are quoted on organized world exchanges as defined by Rauch (1999), who classifies goods into three categories at the 4-digit SITC Rev.2. classification: commodities, reference-priced goods and differentiated goods. This classification is based on whether a good is traded and priced on organized world exchanges, listed in trade publications but not traded on organized exchanges, or does not posses a reference price, respectively.

In order to identify global goods within the 4-digit ISIC Rev.3 classification, we identify all the ISIC Rev.3 codes associated with each SITC Rev.2 in Rauch (1999) by using the UN correspondence tables. We assign a value of 1 to each good priced in organized world exchanges, a value of 2 to each good listed in trade publications but not traded on organized exchanges, and a value of 3 to each good that does not possess a reference price. We calculate the average value of the associated codes for each ISIC Rev.3 code in a similar way to Jensen (2006). We define a good as globally traded at the 4-digit ISIC Rev.3 level if the value of the calculated average

[9]See Lipschitz and McDonald (1992); Marsh and Tokarick (1996); Agenor (1995); Turner and Van t'Dack (1993); and Cerra, Soikkeli, and Saxena (2003) for more on the advantages and disadvantages of using ULC as a measure of cost competitiveness.

within each ISIC Rev.3 code lies in the interval of [1,2). Goods with values in the interval of [2,3] are not considered as globally traded goods.

Rauch (1999) presents two classifications: the "conservative" and the "liberal." The liberal version maximizes the number of globally traded goods in the cases where there was room for discretion in the sorting. Table A2 in Appendix presents the resulting group of goods identified as global goods at the 4-digit ISIC Rev.3 classification under both alternatives, conservative and liberal.

For further comparisons, we also report the goods that appear as global when applying a methodology similar to the one applied to Rauch's list to the list of global (commodity) goods defined by Bayoumi, Jayanthi, and Lee (2005, 2006) at the 2-digit SITC Rev.3 level. The resulting number of global goods from the latter source is higher than the one resulting from our methodology, most likely because in Bayoumi, Jayanthi, and Lee (2005, 2006) the list is defined at a more aggregated category of goods than in Rauch (1999) and this paper. In our analysis, we use Rauch's (1999) liberal classification following Jensen (2006), which results in a list that is closer to the one implied by the methodology used in Bayoumi, Jayanthi, and Lee (2005, 2006).

III. Results

This section presents the sensitivity analysis of the REER to the HPA and to the inclusion of services exports. Based on the methodology described in Section II, we estimate the path of the REER indices under the alternative approaches and present comparative statistics.

We first compare the estimated path of the REER under the HPA, denoted by R^G, with the equivalent REER index under the RPA, denoted by R^{1G}. The difference represents the effect of relaxing the assumption that all non-global goods are treated as identical goods, and as a result compete in the same market $\bar{g}$,d. Second, we compare R^G with the estimates obtained for the evolution of the REER that considers only exports of services (under the HPA), denoted by R^S. This allows us to have a perspective of how the REER for goods compares with the REER for services. Third, we compute the aggregated REER index for goods and services, denoted by R^{GS}, and we compare it with R^{1G}. This difference represents the sensitivity of the REER under the RPA to both the HPA and a broader coverage of exports that includes services. In addition, we perform different robostness checks.

We also study the sensitivity of the REER for goods to heterogeneous cost dynamics across sectors. We compare R^G with the estimated REER for goods under the HPA and the heterogeneous cost dynamics assumption, denoted by R^{Gd}, where d stands for differentiated cost measures. As detailed in the introduction, these results should be taken with caution because of data limitations and should be read as an exploratory effort to determine the effect of differentiated sectoral cost measures on the REER.

The contrast between different approaches, for example, the comparison between R^G and R^{1G}, is performed in two dimensions. First, we present the difference observed in the appreciation rates from 1998 to the corresponding year shown in the tables for both indices. Using 1998 as the base year for comparison is an ad hoc rule, which has no other merit than being year prior to the adoption of the euro adoption by all the euro area countries (January 1, 1999), except for Greece (January 1, 2001). This difference in levels is computed following equation (7).

$$\text{Difference in levels (Level)} = \Delta_t\%R_i^G - \Delta_t\%R_i^{1G}, \tag{7}$$

where $\Delta_t\%$ refers to the growth rate of the index observed from 1998 to year t and i refers to the country whose REER is analyzed.

Second, we study the difference observed between the two estimations of the REER considered, but constructing each estimator relative to the corresponding REER observed in the remaining 11 countries of the euro area. This difference-in-difference estimator is computed following equation (8).

$$\begin{aligned}&\text{Difference in difference (DD)}\\&\quad= \left[\Delta_t\%R_i^G - \Delta_t\%R_{EA}^G\right] - \left[\Delta_t\%R_i^{1G} - \Delta_t\%R_{EA}^{1G}\right],\end{aligned} \tag{8}$$

where $\Delta_t\%$ refers to the growth rate of the index observed from 1998 to year t, i refers to the country whose REER is analyzed, and EA refers to the remaining 11 countries of the euro area.[10]

The difference-in-difference (DD) estimator is our preferred estimator for two reasons. First, it allows us to control for methodological issues specific to each type of estimation that could be driving the results without necessarily reflecting changes in relative cost competitiveness. Second, it allows us to control for the equivalent results observed in the rest of 11 euro area countries. Therefore, the DD estimator represents the change in the relative international competitiveness position between the euro area countries. This does not mean that the DD estimator considers only direct competitors from the euro area, but that it compares among euro area countries each country's position with respect to its direct competitors across the world.

The euro area countries are all part of the same currency union, and therefore, are a natural benchmark to compare the evolution of the REER in each of the MQ countries and control for potential methodological differences particular to each type of estimation. This approach has also an important economic meaning. The exchange rates between euro area countries are fixed—although the euro is still sensitive to the international competitiveness of the euro area as a whole, in line with the standard exchange rate mechanisms associated with floating currencies. No

[10]The index R_{EA}^G is computed as a geometrical index of the individual R^G indices for each of the remaining 11 countries. We construct the weights based on their relative size of exports of the particular type of product considered (goods, services, or goods and services).

rebalancing through nominal exchange rate movements is then possible between euro area countries, but only through productivity and wage growth differentials, which tend to take longer to materialize. As a result, divergence in international competitiveness between euro area countries is an important element when assessing the medium-term economic perspective of individual euro area countries.

The Effect of the Heterogeneous-Product Approach and Services

REER for Goods under the HPA

Table 2 presents the estimation of the REER indices R^G and R^{1G} for the MQ countries, alongside the estimations for the two main euro area countries, France and Germany, for further comparison. The contrast between R^G and R^{1G} is reported in Table 3. The figures for the DD estimator imply that under the HPA Portugal's, Italy's, and Spain's REERs are less appreciated in the range of 2 percent in 2006 (1998 base). The difference is larger in the case of Greece, on the order of 7 percent. This indicates that, relative to the effect on the other 11 euro area countries, the REER under the RPA in Greece is 7 percent more appreciated than what the model assuming the HPA suggests (since 1998).

Robustness Checks

We study the robustness of our methodology and results by performing three additional contrasts. First, we compare our computation of R^{1G} with the closest measure available in the literature (based on Bayoumi, Jayanthi, and Lee, 2005, 2006); second, we modify the sample of countries considered; and third, we study how our measure of REER changes when domestic market competition is considered.

Table 4 presents the comparison between the R^{1G} and R^{IMF}, where IMF stands for the WEO estimates of the REER based on the methodology proposed by Bayoumi, Jayanthi, and Lee (2005, 2006)—the closest source to our methodology that includes the latest developments in the literature and uses the RPA. The results for the DD estimator are all within the ± 1 percent range, suggesting that our methodology under the RPA yields similar results to the existing methodologies based on the RPA.[11]

Our sample of countries is based on the available information for ULC in manufacturing in the OECD and the WEO databases (Table A1). This sample differs from the sample used to compute R^{IMF}, which considers 27 countries. An interesting aspect of the additional 11 countries used in our sample is that they constitute a sample of emerging countries not represented in the sample of 27 advanced economies, with the exception of China.[12]

[11]Our estimates as well as the IMF estimates reported are based on ULC data as of August 2007.

[12]The sample of 27 countries covers on average 70 to 85 percent of MQ's competitors, while the sample of 38 countries covers 80 to 90 percent.

Table 2. Unit-Labor-Cost-Based Real Effective Exchange Rate (REER) Indices, Goods, 1998–2006: Heterogeneous-Product Approach (G) and Representative-Product Approach (1G)

	Greece		Italy		Portugal		Spain		France		Germany	
	G	1G	G	1G	G	1G	G	1G	G	1G	G	1G
1998	100.00	100.00	100.00	100.00	100.00	100.00	100.00	100.00	100.00	100.00	100.00	100.00
1999	96.74	96.53	99.61	99.70	101.86	102.14	99.71	99.83	97.45	97.31	98.09	98.41
2000	89.76	89.27	94.75	94.61	99.03	99.28	98.73	98.67	91.40	91.31	90.76	91.28
2001	85.99	85.31	96.98	96.97	99.00	99.35	99.46	99.50	90.25	90.24	89.88	90.35
2002	89.44	90.03	102.12	102.50	101.47	102.00	101.58	102.03	92.41	92.41	92.63	93.02
2003	94.33	96.12	113.13	114.16	103.22	104.08	108.06	109.22	95.94	95.71	97.02	97.04
2004	107.10	112.30	120.88	122.26	102.30	103.27	112.41	113.84	97.50	96.92	97.52	97.09
2005	104.81	110.04	124.64	125.98	102.90	104.08	114.79	116.16	96.18	95.24	92.60	91.65
2006	107.91	113.90	127.96	129.21	103.09	104.21	116.58	117.77	97.48	96.35	90.31	88.99

Note: G and 1G refer to the unit-labor-cost-based REER for goods estimated using the heterogeneous-product approach and the representative-product approach, respectively. The base year is 1998.

Table 3. Net Appreciation Differential Since 1998: Marginal Effect of Heterogeneous-Product Approach (G vs. 1G)

	Greece		Italy		Portugal		Spain		France		Germany	
	Level	DD	Level	DD	Level	DD	Level	DD	Level	DD	Level	DD
1999	0.21	0.27	−0.09	−0.04	−0.29	−0.23	−0.12	−0.07	0.13	0.23	−0.32	−0.38
2000	0.49	0.61	0.13	0.28	−0.25	−0.14	0.06	0.18	0.08	0.23	−0.52	−0.60
2001	0.68	0.89	0.01	0.25	−0.34	−0.14	−0.04	0.18	0.00	0.25	−0.47	−0.39
2002	−0.59	−0.40	−0.38	−0.21	−0.53	−0.34	−0.45	−0.27	0.00	0.23	−0.39	−0.29
2003	−1.79	−1.64	−1.03	−0.99	−0.86	−0.71	−1.16	−1.06	0.22	0.46	−0.02	0.19
2004	−5.21	−5.30	−1.38	−1.64	−0.98	−1.06	−1.43	−1.59	0.58	0.61	0.43	0.54
2005	−5.22	−5.64	−1.34	−1.93	−1.18	−1.59	−1.37	−1.86	0.95	0.68	0.95	0.86
2006	−5.98	−6.66	−1.25	−2.12	−1.12	−1.78	−1.18	−1.93	1.12	0.58	1.31	1.04

Note: G and 1G refer to the unit-labor-cost-based real effective exchange rated for goods estimated using the heterogeneous-product approach and the representative-product approach, respectively. ''Level'' denotes the difference between the growth rates of G and 1G. DD denotes the difference-in-difference estimate between G and 1G. Results are presented in percentages.

Table 4. Net Appreciation Differential Since 1998: Comparing 1G vs. IMF

	Greece		Italy		Portugal		Spain		France		Germany	
	Level	DD	Level	DD	Level	DD	Level	DD	Level	DD	Level	DD
1999	0.06	–0.27	0.13	–0.22	0.26	–0.07	0.04	–0.31	0.15	–0.21	0.51	0.27
2000	–0.56	–0.76	–0.40	–0.69	0.14	–0.05	–0.58	–0.83	–0.27	–0.56	0.20	0.00
2001	–0.60	–0.63	–0.67	–0.80	0.08	0.05	–0.91	–1.00	–0.33	–0.43	–0.05	–0.12
2002	0.14	–0.30	0.03	–0.47	0.57	0.13	–0.51	–1.01	0.30	–0.17	0.46	0.03
2003	0.53	0.17	0.51	0.17	0.87	0.51	–0.36	–0.77	0.36	0.00	0.53	0.25
2004	1.07	0.29	1.28	0.56	1.22	0.45	–0.01	–0.84	0.71	–0.08	1.12	0.51
2005	0.35	0.15	0.50	0.33	0.81	0.62	–0.67	–0.93	0.16	–0.05	0.45	0.38
2006	0.24	0.15	0.36	0.30	0.64	0.56	–0.82	–0.96	0.06	–0.03	0.27	0.28

Note: 1G refers to the unit-labor-cost-based real effective exchange rate for goods estimated using the representative-product approach and IMF refers to the unit-labor-cost-based real effective exchange rate calculated by the IMF (based on Bayoumi, Jayanthi, and Lee, 2005). "Level" denotes the difference between the growth rates of 1G and IMF. DD denotes the difference-in-difference estimate between 1G and IMF. Results are presented in percentages.

We performed a second robustness check to study if our estimates of R^G are sensible to including the additional 11 emerging countries. We contrast R^G with the REER estimated under the HPA considering only the 27 countries, denoted by R^{G27}. The results, presented in Table 5, indicate that the REER estimated with the sample of 27 countries does not differ substantially from the REER estimated with the sample of 38 countries. The results for the DD estimator are all within the ± 1 percent range.

Finally, we consider the potential importance of domestic market competition for measuring international competitiveness. Owing to a lack of consistent data on disaggregated internal production across MQ countries, our analysis centers on the external markets where each MQ country competes. Table 6 presents the results of contrasting R^G with the REER under HPA including the available information on internal markets, denoted by R^{GIM}. It indicates that the marginal effect of domestic markets is small with differences in the range of ± 1 percent.[13]

Difference in the Structure of Competitors

We complement the results on the effect of the HPA presented in Table 3 with an aggregate view of the difference in the structure of competitors implied by the HPA and the RPA. The larger the difference, the greater the likelihood of finding a large difference between the corresponding REER measures. The actual effect, however, will depend on the interaction between the different weights and the distribution of the change in ULC across countries. In the limit, even large differences in the weights will have no effect if all countries present identical changes in their ULCs, and vice versa, even small differences in the weights can have large effects if changes in ULC are significantly different across countries.

We capture the difference in the structure of competitors implied by each approach using the formula described in equation (9). The variable $\lambda_{i,t}$ aggregates the difference observed in the weight assigned to each competitor of country i in period t.

$$\lambda_{i,t} = \sum_{\forall c \neq i} \left| \chi_{i,c,t} - \chi^{1g}_{i,c,t} \right|, \tag{9}$$

where $\chi_{i,c,t}$ is defined by equation (10) and $\beta_{i,g,d,t} \cdot \alpha_{c,g,d,t}$ refers to the relative importance of each competitor c in market (g,d) with respect to all other competitors in all other markets.[14] The variable $\chi^{1g}_{i,c,t}$ refers to the calculation under the RPA.

$$\chi_{i,c,t} = \sum_{\forall (g,d)} \beta_{i,g,d,t} \cdot \alpha_{c,g,d,t}, \quad \text{where} \sum_{\forall c \neq i} \chi_{i,c,t} = 1. \tag{10}$$

[13]We cannot perform this analysis for Greece because of insufficient data on its disaggregated structure of production.

[14]See Section I for more details.

Table 5. Net Appreciation Differential Since 1998: Marginal Effect of Extended Sample (G vs. G27)

	Greece		Italy		Portugal		Spain		France		Germany	
	Level	DD	Level	DD	Level	DD	Level	DD	Level	DD	Level	DD
1999	−0.61	−0.75	0.05	−0.10	−0.09	−0.23	0.05	−0.09	0.09	−0.07	0.28	0.21
2000	−1.25	−1.22	−0.19	−0.18	−0.46	−0.43	−0.19	−0.16	−0.06	−0.03	0.16	0.30
2001	−1.13	−0.85	−0.42	−0.16	−0.54	−0.27	−0.37	−0.09	−0.22	0.07	−0.16	0.17
2002	−0.99	−0.71	−0.42	−0.16	−0.42	−0.14	−0.29	−0.01	−0.21	0.08	−0.21	0.10
2003	−0.67	−0.57	−0.25	−0.17	−0.24	−0.14	−0.05	0.06	−0.09	0.02	−0.02	0.13
2004	−0.87	−0.77	−0.31	−0.24	−0.28	−0.18	−0.03	0.08	−0.10	0.01	0.01	0.17
2005	−1.58	−1.20	−0.83	−0.49	−0.64	−0.26	−0.42	−0.03	−0.35	0.05	−0.21	0.27
2006	−1.38	−1.05	−0.81	−0.54	−0.54	−0.21	−0.37	−0.04	−0.31	0.03	−0.13	0.30

Note: G and G27 refer to the unit-labor-cost-based real effective exchange rate for goods estimated using the heterogeneous-product approach with the full sample and with the reduced sample of 27 countries, respectively. “Level” denotes the difference between the growth rates of G and G27. DD denotes the difference-in-difference estimate between G and G27. Results are presented in percentages.

Table 6. Net Appreciation Differential Since 1998: Marginal Effect of Domestic Market Competition (G vs. GIM)

	Greece		Italy		Portugal		Spain		France		Germany	
	Level	DD	Level	DD	Level	DD	Level	DD	Level	DD	Level	DD
1999	…	…	−0.26	−0.05	−0.11	0.10	−0.22	−0.01	−0.15	0.07	−0.40	−0.27
2000	…	…	−0.82	−0.07	−0.52	0.25	−0.90	−0.15	−0.99	−0.27	−1.15	−0.58
2001	…	…	−1.00	−0.31	−0.48	0.25	−1.08	−0.37	−0.95	−0.26	−0.95	−0.33
2002	…	…	−0.76	−0.38	−0.06	0.37	−0.90	−0.50	−0.48	−0.07	−0.50	−0.10
2003	…	…	−0.26	−0.40	0.83	0.76	−0.47	−0.59	0.40	0.37	0.20	0.17
2004	…	…	0.03	−0.46	1.38	0.96	−0.20	−0.68	0.98	0.66	0.69	0.39
2005	…	…	−0.18	−0.68	1.44	1.02	−0.35	−0.83	0.99	0.67	0.77	0.52
2006	…	…	−0.26	−0.79	1.57	1.14	−0.33	−0.83	1.04	0.70	0.79	0.54

Note: G and GIM refer to the unit-labor-cost-based real effective exchange rate for goods estimated with the heterogeneous-product approach and with the heterogeneous-product approach including domestic market competition, respectively. "Level" denotes the difference between the growth rates of G and GIM. DD denotes the difference-in-difference estimate between G and GIM. Results are presented in percentages.

Table 7 presents the results for $\lambda_{i,t}$. Greece, whose sensitivity to the HPA is the highest among all countries, presents the highest level of difference. This was expected, although as mentioned above, the fact that a difference is observed in $\lambda_{i,t}$ does not necessarily imply a difference in the REER measure; it only makes it more likely. In fact, among the rest of the countries, Portugal stands out with the largest value for $\lambda_{i,t}$, but does not present the highest sensitivity to the HPA.

Table 7 also presents the difference in the structure of competitors when comparing the HPA and the partners-approach, denoted by λ_p. The partners-approach refers to considering a country's trade partners as its competitors and it is used by some sources; see Chinn (2006) for more details. The results show that considering the partners-approach yields a stronger difference, two to three times the size of $\lambda_{i,t}$. A larger difference is expected because considering competitors only on the base of trade partnerships deviates substantially from the concept of competitors used in this paper.

REER for Goods and Services under the HPA

Table 8 presents the estimation of the REER index R^S, which includes only services exports. To illustrate the evolution of the services component, we compare R^S and R^G in Table 9. These results suggest that except for the case of Greece, the services component of the REER has appreciated less than the goods component for the MQ countries. This difference ranges from −3.4 percent for Italy to −0.9 percent for Spain (DD estimator). For Greece, the difference goes in the opposite direction in the range of 7 percent (DD estimator).

Finally, Table 10 presents the estimation of the REER index R^{GS}, which includes both goods and services. We compare R^{GS} with R^{1G} in Table 11, which represents the aggregate sensitivity of the REER under the RPA to both the HPA and a broader coverage of exports that includes services.

The results suggest that these two additional factors together—HPA and services—have had a marginal effect on the REER on the order of −2 percent to −3 percent for all the MQ countries: −2 percent for Greece, −2.3 percent for Spain, −2.4 percent for Portugal, and −2.8 percent for Italy. These numbers are consistent with the previous tables, where the smaller appreciation observed in goods for Italy, Portugal, and Spain under the HPA adds to the smaller appreciation observed in services relative to goods. For the case of Greece, the strongest difference observed under the HPA shrinks significantly when combined with the larger appreciation observed in services relative to goods.

Differentiated ULC by Sector

Product heterogeneity (HPA) and services exports refine the REER as a measure of international competitiveness, but, as detailed in the previous

Table 7. The Structure of Competitors: Heterogeneous- vs. Representative-Product Approach (λ) and vs. Partners Approach (λp)

	Greece		Italy		Portugal		Spain		France		Germany	
	λ	λp	λ	λp	λ	λp	λ	λp	λ	λp	λ	λp
1999	0.134	0.309	0.054	0.170	0.105	0.196	0.070	0.186	0.051	0.203	0.063	0.219
2000	0.135	0.325	0.060	0.174	0.105	0.213	0.073	0.194	0.052	0.203	0.065	0.211
2001	0.132	0.347	0.062	0.179	0.106	0.213	0.071	0.202	0.056	0.206	0.065	0.202
2002	0.127	0.342	0.065	0.182	0.105	0.200	0.074	0.213	0.057	0.205	0.068	0.203
2003	0.125	0.319	0.065	0.190	0.102	0.200	0.074	0.206	0.058	0.210	0.067	0.204
2004	0.131	0.318	0.063	0.197	0.096	0.193	0.072	0.220	0.058	0.208	0.068	0.201
2005	0.132	0.310	0.065	0.189	0.099	0.197	0.070	0.222	0.062	0.196	0.068	0.198
Average	0.133	0.322	0.061	0.183	0.103	0.201	0.072	0.204	0.056	0.204	0.067	0.208

Note: Following equation (9), λ refers to the difference in the structure of competitors for goods when comparing the heterogeneous-product approach with the representative-product approach. Likewise, λp refers to the difference in the structure of competitors when comparing the heterogeneous-product approach with the partners-approach for goods.

Table 8. Unit-Labor-Cost-Based Real Effective Exchange Rate Indices, Services, 1998–2006: Heterogeneous-Product Approach (S)

	Greece	Italy	Portugal	Spain	France	Germany
1998	100.00	100.00	100.00	100.00	100.00	100.00
1999	95.61	98.94	101.71	99.74	96.33	97.80
2000	84.71	93.21	97.59	97.90	88.97	90.67
2001	80.27	95.88	97.59	98.73	88.14	89.97
2002	86.70	101.03	100.26	100.97	90.27	92.20
2003	94.95	112.21	102.80	109.23	94.76	96.07
2004	113.06	118.95	101.64	112.98	96.07	96.28
2005	110.78	122.49	101.83	114.58	94.15	91.28
2006	114.55	125.62	101.39	115.35	95.19	89.02

Note: S refers to the unit-labor-cost-based real effective exchange rate for services estimated using the heterogeneous-product approach. The base year is 1998.

section, these two factors do not change substantially the broad picture of international competitiveness in the MQ. In this section, we explore the sensitivity of the REER to the HPA with differentiated cost measures at the sector-level. Differentiated cost measures would yield a more accurate picture of international competitiveness to the extent that productivity and production costs vary across sectors. The set of results presented in this section, which indicate a higher sensitivity of the REER relative to the assumption of homogenous cost dynamics, should, however, be taken with caution. Given the data limitations detailed before, these results should be read as an exploratory effort to determine the effect of differentiated sectoral cost measures on the REER.

The results for the contrast between the REER estimated with an aggregated ULC measure (R^G) and the REER estimated with a differentiated ULC by sector (R^{Gd}) point to a higher sensitivity of the REER to the assumption of homogenous cost dynamics across sectors—both measures based on the sample for goods because of the excessive volatility found in the data for service. As shown in Table 12, the absolute differences range between 2 and 6 percent. For this contrast, R^G is computed using the limited data set available for the calculation of R^{Gd}.

These results are not sensitive to outliers. We recalculated the REER eliminating the 0.5 percent tails of the distribution of the annual ULC growth rates observed since 1998. No substantial differences from the results obtained in Table 12 were found. As an additional robustness check, we compared the results obtained for the contrast between R^G and R^{1G} (Table 3) with an equivalent contrast using the limited data set available for the calculation of R^{Gd}. The differences between both cases are all within the ± 1 percent range.

Table 9. Net Appreciation Differential Since 1998: Difference between Services and Goods (S vs. G)

	Greece		Italy		Portugal		Spain		France		Germany	
	Level	DD	Level	DD	Level	DD	Level	DD	Level	DD	Level	DD
1999	−1.13	−0.18	−0.67	0.36	−0.14	0.91	0.03	1.26	−1.11	−0.12	−0.29	0.94
2000	−5.05	−4.07	−1.54	−0.39	−1.44	−0.16	−0.84	0.82	−2.42	−1.42	−0.09	1.82
2001	−5.72	−5.28	−1.10	−0.33	−1.41	−0.50	−0.74	0.57	−2.11	−1.48	0.09	1.79
2002	−2.73	−1.63	−1.09	0.17	−1.21	0.18	−0.62	1.23	−2.14	−0.94	−0.43	1.63
2003	0.61	1.37	−0.92	−0.24	−0.42	0.50	1.18	2.79	−1.18	−0.34	−0.94	0.66
2004	5.97	7.01	−1.93	−1.73	−0.66	−0.01	0.57	1.93	−1.43	−0.95	−1.25	0.44
2005	5.96	6.40	−2.15	−2.66	−1.07	−0.99	−0.21	0.63	−2.03	−2.36	−1.32	0.57
2006	6.64	6.77	−2.34	−3.36	−1.70	−2.05	−1.24	−0.86	−2.29	−3.15	−1.28	0.65

Note: S refers to the unit-labor-cost-based real effective exchange rate for services estimated using the heterogeneous-product approach and G refers to the unit-labor-cost-based real effective exchange rate for goods estimated using the heterogeneous-product approach. “Level” denotes the difference between the growth rates of S and G. DD denotes the difference-in-difference estimate between S and G. Results are presented in percentages.

Table 10. Unit-Labor-Cost-Based Real Effective Exchange Rate Indices, Goods and Services, 1998–2006: Heterogeneous-Product Approach (GS)

	Greece	Italy	Portugal	Spain	France	Germany
1998	100.00	100.00	100.00	100.00	100.00	100.00
1999	96.00	99.49	101.81	99.71	97.24	98.06
2000	86.27	94.48	98.62	98.50	90.94	90.75
2001	82.01	96.79	98.60	99.25	89.86	89.89
2002	87.67	101.93	101.12	101.40	92.02	92.58
2003	95.03	112.96	103.11	108.36	95.77	96.91
2004	111.42	120.54	102.13	112.56	97.29	97.39
2005	109.16	124.26	102.62	114.75	95.85	92.47
2006	112.71	127.55	102.64	116.30	97.10	90.18

Note: GS refers to the unit-labor-cost-based real effective exchange rate for goods and services estimated using the heterogeneous-product approach. The base year is 1998.

IV. The Profile of International Competitors

The HPA proposed in our study also allows a quantitative assessment of each country's profile of competitors. Such evidence provides information about the exposure of each country to its key competitors around the world—for example, the exposure to emerging competitors like China, a country that has shown a strong pattern of productivity and trade growth, or the exposure to countries facing significant changes in their cost structure, such as the wage moderation observed recently in Germany or the depreciation of the nominal exchange rate observed in the United States during recent years. Our definition of markets also captures the potential vulnerability of each country's sectors to changing market conditions in competitors' sectors beyond the country level.

Goods

For all six countries, the bulk of competition comes from the advanced and emerging economies, representing on average 95 percent in goods (except for Greece, 92 percent) and 98 percent in services. Since the late 1990s, there has been a change in the composition with emerging economies taking greater importance: they represented in 2005 14 percent of overall exposure to competition in goods for Spain, 19 percent for Italy and Portugal, and 22 percent for Greece (Table 13). China appears as the largest emerging competitor in goods for all four countries, representing at least half of the increase in the importance of emerging economies since 1998.

Among the advanced economies, the euro area countries represent 59 percent of the competition in goods faced by Spain and Portugal, 49 percent for Italy, and 47 percent for Greece. These data indicate that Spain and Portugal are more exposed to euro area competition and therefore less

Table 11. Net Appreciation Differential Since 1998: Joint Marginal Effect of the Heterogeneous-Product Approach, Including Services (GS vs. 1G)

	Greece		Italy		Portugal		Spain		France		Germany	
	Level	DD	Level	DD	Level	DD	Level	DD	Level	DD	Level	DD
1999	−0.53	−0.30	−0.21	0.04	−0.33	−0.08	−0.12	0.15	−0.08	0.20	−0.35	−0.15
2000	−3.00	−2.69	−0.14	0.25	−0.66	−0.30	−0.17	0.27	−0.38	−0.03	−0.53	−0.17
2001	−3.29	−3.00	−0.18	0.21	−0.75	−0.37	−0.25	0.21	−0.38	−0.01	−0.46	0.00
2002	−2.36	−2.00	−0.57	−0.18	−0.88	−0.46	−0.63	−0.15	−0.39	0.02	−0.44	0.08
2003	−1.09	−0.86	−1.20	−1.07	−0.97	−0.71	−0.85	−0.53	0.06	0.38	−0.13	0.37
2004	−0.88	−0.87	−1.72	−2.00	−1.15	−1.20	−1.28	−1.29	0.36	0.38	0.30	0.68
2005	−0.88	−1.32	−1.72	−2.50	−1.46	−1.97	−1.41	−1.86	0.62	0.15	0.82	1.03
2006	−1.19	−1.96	−1.67	−2.82	−1.57	−2.44	−1.46	−2.27	0.75	−0.11	1.19	1.21

Note: GS refers to the unit-labor-cost-based real effective exchange rate for goods and services estimated using the heterogeneous-product approach and 1G refers to the unit-labor-cost-based real effective exchange rate for goods estimated using the representative-product approach. "Level" denotes the difference between the growth rates of GS and 1G. DD denotes the difference-in-difference estimate between GS and 1G. Results are presented in percentages.

Table 12. Net Appreciation Differential Since 1998: Marginal Effect of Differentiated ULC (Gd vs. G)

	Greece		Italy		Portugal		Spain		France		Germany	
	Level	DD	Level	DD	Level	DD	Level	DD	Level	DD	Level	DD
1999	0.15	−0.04	−1.29	−1.71	−2.21	−2.43	−0.09	−0.30	−0.82	−1.20	2.22	2.96
2000	1.61	0.95	−0.88	−1.79	−1.14	−1.83	−0.48	−1.23	−0.04	−0.85	4.00	4.86
2001	0.85	0.48	−0.82	−1.38	−1.17	−1.57	−1.48	−1.98	0.59	0.25	2.23	2.72
2002	2.57	2.05	−1.09	−1.86	−2.82	−3.40	−0.89	−1.52	0.42	−0.13	1.23	1.02
2003	1.83	1.24	−1.28	−2.12	−1.88	−2.51	−0.71	−1.38	1.44	1.02	0.44	−0.22
2004	1.32	−0.11	−1.92	−3.79	−0.95	−2.41	−2.90	−4.61	0.96	−0.57	2.38	1.41
2005	5.57	4.24	−3.98	−6.00	−0.64	−2.02	−3.67	−5.34	−0.30	−2.00	2.90	2.32
2006	…	…	…	…	…	…	…	…	…	…	…	…

Note: Gd and G refer to the unit-labor-cost-based real effective exchange rate for goods estimated using the heterogeneous-product approach with industry-level unit labor cost measures and the heterogeneous-product approach with a country-level unit labor cost measure, respectively. "Level" denotes the difference between the growth rates of Gd and G. DD denotes the difference-in-difference estimate between Gd and G. Results are presented in percentages.

Table 13. Structure of Competitors: Goods
(In percent)

	Greece (%)	Italy (%)	Portugal (%)	Spain (%)	France (%)	Germany (%)
Euro area, 1998	47.5	49.7	60.0	59.9	48.9	42.6
Euro area, 2005	47.0	48.6	58.6	58.5	48.6	41.1
Advanced economies, 1998	72.4	81.4	83.7	85.9	85.7	84.5
Advanced economies, 2005	69.9	75.7	77.0	81.5	80.1	77.4
Emerging economies, 1998 (1)	18.6	14.7	12.7	11.1	11.7	11.8
Emerging economies, 2005 (2)	22.0	19.2	19.0	14.4	15.9	17.7
Change in percentage points (2)-(1)	3.4	4.5	6.3	3.3	4.3	5.9
Change in percentage points due to China	3.5	3.4	3.4	1.8	2.3	2.6

Note: Percentages refer to the share of competition from each country or group of countries.

Table 14. Structure of Competitors in High- and Low-Tech Sectors in 2005
(In percent)

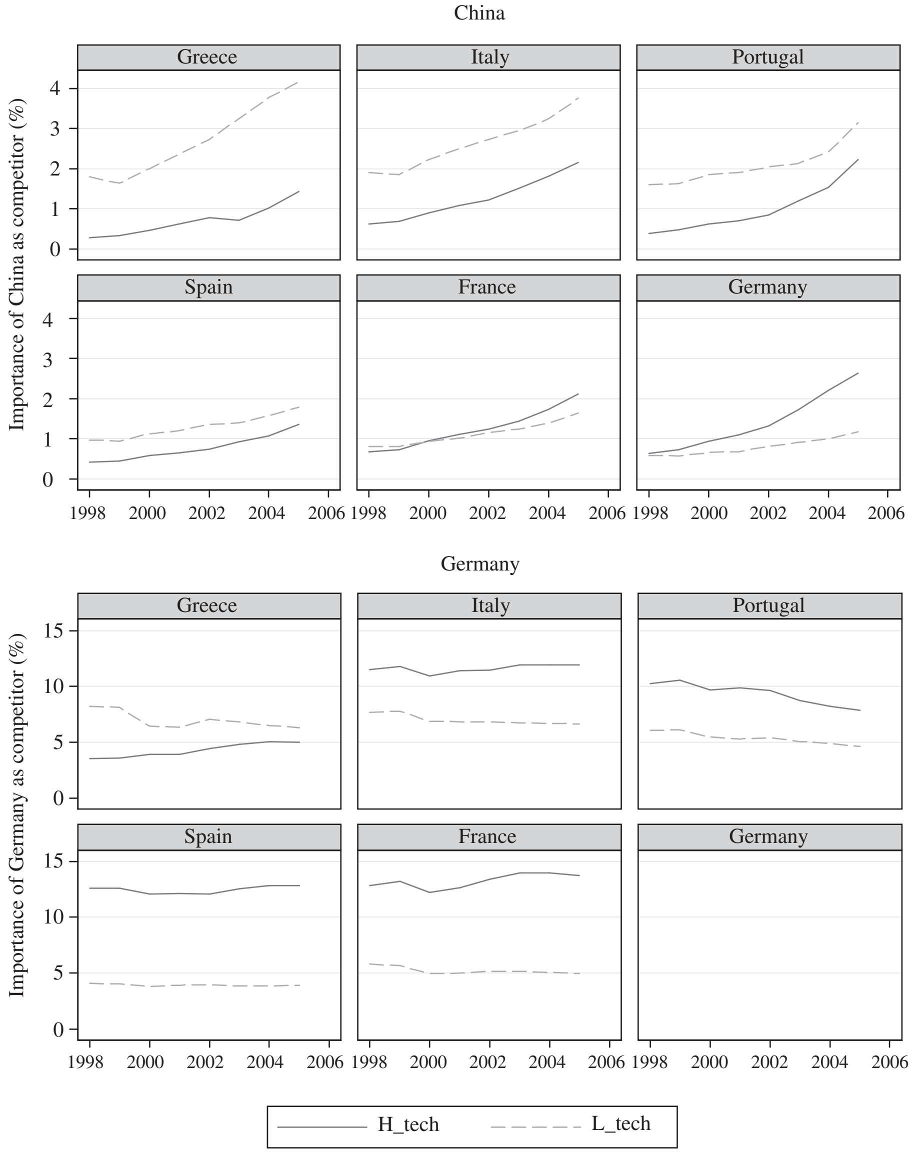
	Greece (%)	Italy (%)	Portugal (%)	Spain (%)	France (%)	Germany (%)
China as competitor						
High-tech sectors	1.43	2.16	2.22	1.36	2.12	2.64
Low-tech sectors	4.17	3.76	3.15	1.79	1.64	1.18
Germany as competitor						
High-tech sectors	5.01	11.94	7.84	12.82	13.74	…
Low-tech sectors	6.29	6.63	4.61	3.90	4.95	…

Note: Percentages refer to the share of competition from China and Germany in each sector. Technology intensity of sectors is defined according to OECD (2007) classification of industries with respect to intensity of technology used in sectors producing goods. High-tech industries comprise 2-digit industries with high and medium-high intensity of technology used and low-tech industries comprise 2-digit industries with low and medium-low intensity of technology used.

exposed to changes in the value of the euro. There has been a declining trend since 1998 in the range of 1 percentage point for all countries, which is smaller than the change observed for the aggregate of advanced economies.

From a sectoral point of view, the four MQ countries compete more in low-technology sectors with China (see Table 14). However, the importance of China in high-technology sectors is growing as well (see Figure 2). As a comparison, France and Germany compete more strongly with China in

Figure 2. Importance of China and Germany as Competitors of the Mediterranean Quartet in High- and Low-Tech Sectors in Goods



Note: Technology intensity of sectors is defined according to OECD (2007) classification of industries with respect to intensity of technology used. High-tech (that is, H_tech) industries comprise industries with high and medium-high intensity of technology used and low-tech (that is, L_tech) industries comprise industries with low and medium-low intensity of technology used. Percentages refer to the share of competition from each country.

Table 15. Structure of Competitors: Services
(In percent)

	Greece (%)	Italy (%)	Portugal (%)	Spain (%)	France (%)	Germany (%)
Euro area, 1999	37.4	49.6	60.3	54.4	42.8	46.0
Euro area, 2004	37.2	50.2	60.3	48.1	42.5	40.7
Advanced economies, 1999	95.9	94.1	97.1	95.8	95.1	93.6
Advanced economies, 2004	92.0	89.7	94.7	93.5	91.1	87.8
Emerging economies, 1999 (1)	3.5	4.9	2.4	3.5	4.2	5.0
Emerging economies, 2004 (2)	6.4	7.5	3.9	4.7	6.7	8.6
Change in % points (2)–(1)	2.9	2.6	1.5	1.2	2.5	3.6
Change in % points due to China	0.3	0.2	0.1	0.1	0.2	0.3
Change in % points due to the largest 5 EE	2.0	1.6	0.6	0.5	1.8	2.3

Note: Percentages refer to the share of competition from each country or group of countries. The largest five emerging economies competitors for Greece are South Korea, Turkey, Hungary, Czech Rep., Hong Kong; for Italy, they are Hungary, Turkey, Czech Rep., South Korea, Hong Kong; for Portugal, they are Turkey, Czech Rep., Egypt, Hungary, Mexico; for Spain, they are Turkey, Czech Rep. Egypt, Hungary, South Africa; for France, they are South Korea, Turkey, Hong Kong, Hungary, Czech Rep.; and for Germany, they are South Korea, Hong Kong, Czech Rep., Turkey and Hungary.

high-technology sectors, suggesting that China should not be seen as a potential competitor in low-tech sectors only.[15]

Figure 2 also show the importance of Germany—the main advanced-country competitor—as a competitor of the MQ. The almost flat or sometimes decreasing importance of Germany highlights the strong growth of China's importance in both high- and low-technology sectors. Nonetheless, at least until 2005, Germany was still a bigger competitor for the MQ than China in both types of sectors.

Services

In services, emerging markets represent on average about one-third of their importance in goods, showing also, although to a lesser extent, a similar increase in recent years (see Table 15). From 1999 to 2004, the composition shifted to emerging economies in the range of 3 percent for Greece and Italy, 2 percent for Portugal, and 1 percent for Spain. The data suggest that China does not appear as a strong competitor in services.

[15]We classify the 2-digit ISIC Rev.3 sectors as low-(L) or high-(H) technology sectors following the OECD (2007) revised classification, which is based on Hatzichronoglou (1997).

Table 16. Main Competitors in 2005: Goods
(In percent)

Rank	Greece	Italy	Portugal	Spain	France	Germany
1	Italy (11.84)	Germany (18.63)	Spain (15.76)	Germany (16.92)	Germany (18.82)	U.S. (13.40)
2	Germany (11.50)	France (11.60)	Germany (12.52)	France (16.46)	U.S. (12.20)	France (11.68)
3	France (7.13)	U.S. (9.72)	France (12.20)	Italy (9.98)	Italy (9.62)	Italy (9.37)
4	U.S. (6.54)	Spain (6.26)	Italy (8.52)	U.S. (7.23)	U.K. (7.76)	U.K. (7.70)
5	U.K. (6.14)	U.K. (6.25)	U.S. (5.81)	U.K. (6.49)	Spain (6.82)	Japan (6.85)
6	China (5.76)	China (5.95)	U.K. (5.63)	Belgium (4.29)	Japan (4.46)	Netherlands (4.78)
7	Spain (4.51)	Japan (4.33)	China (5.41)	Netherlands (4.06)	Belgium (4.24)	Spain (4.71)
8	Belgium (4.15)	Belgium (3.71)	Belgium (3.65)	Portugal (3.95)	Netherlands (4.10)	Belgium (4.14)
9	Netherlands (3.93)	Netherlands (3.23)	Netherlands (3.35)	Japan (3.37)	China (3.80)	China (3.83)
10	Turkey (3.59)	Austria (2.23)	Japan (2.41)	China (3.23)	Korea (1.96)	Sweden (2.73)

Note: Percentages refer to the share of competition from each country.

Among the advanced economies, the euro area countries represent 60 percent of the competition in services faced by Portugal, 50 percent for Italy, 48 percent for Spain, and 37 percent for Greece. There has been a nil trend since 1998 for all countries except for Spain, whose euro area competition has declined by 5 percentage points. These figures for services should be read with caution given the incomplete availability of the data for bilateral trade of services.

Table 16 (goods) and Table 17 (services) present a list of the top 10 competitors for each country with their corresponding weights.

V. Conclusion

We develop a complete methodology to reexamine the evolution of international competitiveness in the MQ, as measured by the REER. In addition to the elements considered in the existing literature, we (1) use a micro-based approach that considers product heterogeneity when identifying each country's international competitors and their weights and (2) include a comprehensive analysis of the services sector. Our approach enriches the

Table 17. Main Competitors in 2004: Services
(In percent)

Rank	Greece	Italy	Portugal	Spain	France	Germany
1	U.S. (29.57)	U.S. (15.85)	U.K. (17.73)	U.K. (26.40)	U.S. (16.20)	U.S. (15.97)
2	U.K. (16.20)	Germany (14.45)	Spain (17.49)	Germany (13.79)	U.K. (16.16)	France (10.38)
3	Germany (9.60)	France (13.12)	France (14.53)	France (11.74)	Italy (10.38)	Japan (10.20)
4	Italy (7.87)	U.K. (11.47)	U.S. (9.24)	U.S. (8.75)	Germany (9.68)	U.K. (9.71)
5	France (7.33)	Spain (6.04)	Germany (8.98)	Italy (7.63)	Japan (8.60)	Italy (9.38)
6	Spain (2.89)	Japan (5.21)	Italy (6.77)	Portugal (3.79)	Spain (5.88)	Netherlands (4.89)
7	Netherlands (2.67)	Austria (5.06)	Belgium (3.81)	Austria (2.67)	Belgium (4.78)	Spain (4.12)
8	Austria (2.57)	Belgium (2.92)	Netherlands (2.79)	Netherlands (2.54)	Netherlands (3.36)	Austria (3.77)
9	Japan (2.51)	Netherlands (2.76)	Austria (1.95)	Sweden (2.52)	Austria (2.65)	Belgium (3.41)
10	Belgium (2.17)	Greece (2.66)	Japan (1.71)	Belgium (2.36)	Sweden (2.23)	Denmark (3.16)

Note: Percentages refer to the share of competition from each country.

REER analysis by identifying more accurately each country's direct international competitors and providing an aggregate view of international competitiveness that encompasses the complete export sector.

Our main findings suggest that the effect of considering both the more micro-based structure of competitors and exports of services implies a modest lower real appreciation from 1998 to 2006 on the order of 2 to 3 percent for all MQ countries—2 percent for Greece, 2.8 percent for Italy, 2.4 percent for Portugal, and 2.3 percent for Spain. These estimates are based on a difference-in-difference estimator that controls for the equivalent effect observed in the rest of 11 euro area countries.

Finally, the methodology proposed in this paper also allows a detailed view of the structure of each country's competitors. Our findings indicate that the bulk of competition for the MQ still comes from the advanced economies, especially from the euro area. Nonetheless, there has been a change in the composition with emerging economies taking more importance since the late 1990s, particularly China in both the high- and low-technology sectors.

APPENDIX

Table A1. Availability of Unit Labor Cost (ULC) Data

	Country	Country-Level ULC: Sample of 38 countries	Country-Level ULC: Sample of 27 countries	Differentiated ULC: Sample of 28 countries
1	Australia	Yes	Yes	Yes
2	Austria	Yes	Yes	Yes
3	Belgium	Yes	Yes	Yes
4	Canada	Yes	Yes	No
5	China, P.R.: Hong Kong	Yes	Yes	No
6	Colombia	Yes	No	No
7	Czech Republic	Yes	No	Yes
8	Denmark	Yes	Yes	Yes
9	Finland	Yes	Yes	Yes
10	France	Yes	Yes	Yes
11	Germany	Yes	Yes	Yes
12	Greece	Yes	Yes	Yes
13	Hungary	Yes	No	Yes
14	Iceland	Yes	No	No
15	Ireland	Yes	Yes	Yes
16	Israel	Yes	Yes	No
17	Italy	Yes	Yes	Yes
18	Japan	Yes	Yes	Yes
19	Korea	Yes	Yes	No
20	Luxembourg	Yes	Yes	Yes
21	Macedonia, FYR	Yes	No	No
22	Mexico	Yes	No	No
23	Netherlands	Yes	Yes	Yes
24	New Zealand	Yes	Yes	No
25	Norway	Yes	Yes	No
26	Poland	Yes	No	Yes
27	Portugal	Yes	Yes	Yes
28	Singapore	Yes	Yes	No
29	Slovak Republic	Yes	No	Yes
30	Slovenia	Yes	No	Yes
31	South Africa	Yes	No	No
32	Spain	Yes	Yes	Yes
33	Sweden	Yes	Yes	Yes
34	Switzerland	Yes	Yes	No
35	Taiwan POC	Yes	Yes	No
36	Turkey	Yes	No	No
37	United Kingdom	Yes	Yes	Yes
38	United States	Yes	Yes	Yes
39	Cyprus	No	No	Yes
40	Estonia	No	No	Yes
41	Lithuania	No	No	Yes
42	Litva	No	No	Yes
43	Malta	No	No	Yes

Note: Country-level unit labor cost data compiled from the OECD database and IMF World Economic Outlook database. Differentiated 2-digit sector level unit labor cost data compiled from EU KLEMS.

Table A2. Classification of Global Goods

ISIC Rev.3 (4-digit)	Activity Description	Rauch Classification		IMF
		Conservative	Liberal	
0111	Growing of cereals and other crops n.e.c.	Global	Global	Global
0112	Growing of vegetables, horticultural specialties and nursery products	—	—	Global
0113	Growing of fruit, nuts, beverage and spice crops	Global	Global	Global
0121	Farming (cattle, sheep, goats, horses, asses, mules and hinnies; dairy)	Global	Global	Global
0122	Other animal farming; production of animal products n.e.c.	—	—	Global
0200	Forestry, logging and related service activities	—	—	Global
0500	Fishing operations	—	—	Global
1110	Extraction of crude petroleum and natural gas	—	Global	—
1200	Mining of uranium and thorium ores	—	Global	Global
1310	Mining of iron ores	Global	Global	Global
1320	Mining of non-ferrous metal ores, except uranium and thorium ores	—	Global	Global
1410	Quarrying of stone, sand and clay	—	—	Global
1421	Mining of chemical and fertilizer minerals	—	—	Global
1422	Extraction of salt	—	—	Global
1429	Other mining and quarrying n.e.c.	—	—	Global
1511	Production, processing and preserving of meat and meat products	Global	Global	Global
1512	Processing and preserving of fish and fish products	—	—	Global
1513	Processing and preserving of fruit and vegetables	—	—	Global
1514	Manufacture of vegetable and animal oils and fats	Global	Global	Global
1520	Manufacture of dairy products	—	Global	Global

1531	Manufacture of grain mill products	—	—	Global
1532	Manufacture of starches and starch products	—	—	Global
1533	Manufacture of prepared animal feeds	—	—	Global
1541	Manufacture of bakery products	—	—	Global
1542	Manufacture of sugar	Global	Global	Global
1543	Manufacture of cocoa, chocolate and sugar confectionery	—	—	Global
1544	Manufacture of farinaceous products (macaroni and similar)	—	—	Global
1549	Manufacture of other food products n.e.c.	—	—	Global
1551	Distilling, rectifying and blending of spirits	—	—	Global
1552	Manufacture of wines	—	—	Global
1553	Manufacture of malt liquors and malt	—	—	Global
1554	Manufacture of soft drinks; production of mineral waters	—	—	Global
1600	Manufacture of tobacco products	—	—	Global
2010	Sawmilling and planing of wood	—	—	Global
2411	Manufacture of basic chemicals, exc. fertilizers & nitrogen compounds	Global	Global	—
2412	Manufacture of basic precious and non-ferrous metals	—	Global	—
2720	Manufacture of basic precious and non-ferrous metals	Global	Global	Global
9302	Hairdressing and other beauty treatment	—	—	Global
Total		**9**	**14**	**35**

Note: The table refers to globally traded goods as goods whose prices are quoted on organized world exchanges as defined by Rauch (1999). These goods are characterized as commodities for which a more appropriate definition of market is at the world level. Rauch distinguishes between conservative and liberal classifications. We use the liberal classification that maximizes the number of global goods. For comparison, we construct an additional list of global goods, denoted by IMF, consistent with the approach followed by Bayoumi, Jayanthi, and Lee (2005).

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Dynamic Aspects of Productivity Spillovers, Terms of Trade, and the "Home Market Effect"

IPPEI FUJIWARA and NAOHISA HIRAKATA*

Recent empirical findings conclude that the terms of trade improve even after the positive productivity shock hits the economy among advanced economies. Corsetti, Martin, and Pesenti (2007), henceforth CMPs analytically show that a static two-country model with endogenous firm entry can generate improvement of the terms of trade in response to a positive technology shock in the form of lowering the entry cost. This paper evaluates the robustness of the results in CMP in a model with richer and more realistic dynamics such as nominal price and wage stickiness as in the Global Economy Model. It shows how the economic variables respond to the shocks that shift the production frontier outward, namely, productivity gains in manufacturing and efficiency gains in creating new firms. The main conclusions are that short-run responses could be different from those in CMP because of the existence of real as well as nominal rigidities, and that the persistence of shocks also alters the direction of responses via the wealth effect. These results suggest that it is of great importance for policy institutions to acknowledge the dynamic aspects of productivity spillovers by simulating a model with richer dynamics. [JEL E32, F12, J41]
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Recent empirical researches, such as Corsetti, Dedola, and Leduc (2006); Corsetti, Martin, and Pesenti (2007) (henceforth CMP); Debaere and Lee (2003); Hummels and Klenow (2005); and Kang (2005), conclude that the terms of trade improve even after the positive productivity shock hits the economy among advanced economies. This is quite contrary to the conventional wisdom because the home technological progress usually results in lower domestic prices. An increased supply of goods from an economy with high productivity growth should be absorbed by the rest of the world at falling prices, which should result in a deterioration of the terms of trade. This is what is concluded in the seminal research by Armington (1969).

CMP analytically show that a static two-country model with endogenous firm entry[1] can generate improvement of the terms of trade in response to a positive technology shock in the form of lowering the entry cost.[2] As there is no direct changes in the real marginal cost stemming from the lowered entry cost, prices set by individual firms do not change very much. Under such circumstances, the terms of trade correspond one to one with nominal exchange rates. Increase in the number of firms reflecting the lower entry cost results in tighter demand conditions for domestic labor and consequently appreciates nominal exchange rates through the relative increase in domestic wages.[3] "Hence, the Home terms of trade strengthen as the array of Home products increases" (CMP, p. 114). On the other hand, however, the welfare-based price index becomes smaller in the domestic country than in the rest of the world as the array of home products increases more. Reflecting these developments in prices, the real exchange rate, the nominal exchange rate denominated by the welfare-based price indices, namely, the international relative price in the extensive margin, depreciates for the domestic country. CMP further demonstrate directions of theoretical responses of international relative prices to the shocks that expand the production frontier, namely an increase in productivity and in labor population in addition to the efficiency gains in creating new firms. CMP show that an increase in labor population works like the efficiency gains, because it increases the number of firms as long as there is a trade cost that induces the home market effect.

[1]Reflecting the growing interests in understanding the role of firm heterogeneity or endogenous variety in the international trade as the seminal research by Melitz (2003) represents, macroeconomists also start considering the consequences of incorporating the firm dynamics in dynamic stochastic general equilibrium models. See, for example, Bergin and Corsetti (2005); Bilbiie, Ghironi, and Melitz (2005, 2006, 2007); Fujiwara (2007); and Ghironi and Melitz (2005).

[2]This point is predicted in Benigno and Thoenissen (2003) and Ghironi and Melitz (2005).

[3]CMP (p. 104) set the nominal wage as a numeraire: "the exchange rate is defined as the relative price of Foreign labor in terms of Home labor units." According to the first-order condition concerning the labor supply, namely that the marginal rate of substitution equals to the real wage, the nominal wage definitely increases as the number of firms increases. Therefore, the exchange rate appreciates and the terms of trade improve in response to the efficiency gains.

This paper evaluates the robustness of the results in CMP in a model with richer and more realistic dynamics such as one period lag in firm dynamics and nominal price as well as wage stickiness.[4] We first set up a model that incorporates firm dynamics into the Global Economy Model (GEM). Then, we demonstrate how the economic variables respond to the shocks that shift the production frontier outwards, namely productivity gains in manufacturing, and efficiency gains in creating new firms. Our main conclusions are that short-run responses could be different from those in CMP because of the existence of real as well as nominal rigidities; and that persistence of shocks also alters the direction of responses via the wealth effect. These results suggest that it is of great importance for policy institutions to acknowledge the dynamic aspects of productivity spillovers by simulating a model with richer dynamics like the GEM.

The next section of this paper briefly introduces our model. Section II analyzes the international implication of shocks, which expand the production frontier outwards, on international relative prices. We examine two shocks, namely, productivity gains in manufacturing and efficiency gains in creating new firms. We show how responses could be different in a model with richer and more realistic dynamics than those obtained analytically in a static model in CMP. Section III shows such responses in a two-country model calibrated for the Japan and the rest of the world following Laxton, N'Diaye, and Pesenti (2006).

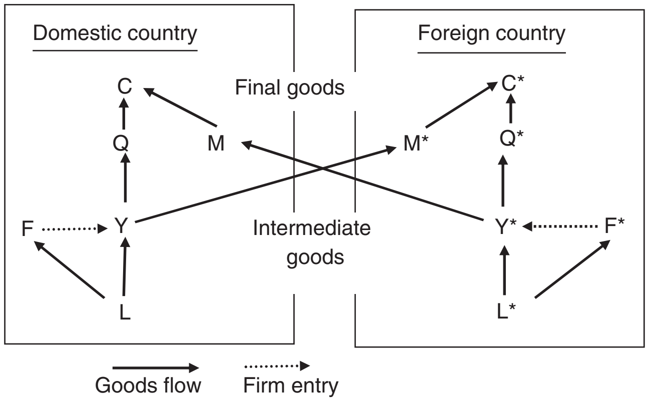
I. The Model

The model used in this paper is based on the recent literature, which combines the ‘‘New Trade Theory’’ advocated by Krugman (1980). Because the heterogeneous technology level among firms is not considered in this paper, our model can be interpreted as a dynamic extension of CMP or a two-country extension of Bilbiie, Ghironi, and Melitz (2005, 2006, 2007) and Bergin and Corsetti (2005). Contrary to these previous studies, our model includes richer and more realistic dynamics, such as nominal price and wage rigidities. Because our model is based on the GEM, it has richer dynamics and can reproduce the regularities in the observed data.

The production structure of this model is summarized in Figure 1. Intermediate goods (Y) are produced using only labor (L), which is also used for creating new firms (F). Then, they are used as intermediate inputs in final goods (Q) production[5] in either the domestic or the foreign country as

[4]For the details of the model and simulations, please see Fujiwara and Hirakata (2007, 2008). The 2007 paper shows all the details of the model and simulation, and investigates some features not shown in this paper such as the importance of the elasticity of substitution between domestic and foreign goods for the direction of responses in international relative prices. Fujiwara and Hirakata (2008) focus exclusively on this elasticity issue in the same framework as employed in CMP.

Figure 1. Production Structure



imported intermediate inputs (M). All final goods produced this way (C) are consumed. The details are shown in Fujiwara and Hirakata (2007).

II. International Implications of Expanding the Production Frontier

This section examines whether the directions of responses analytically derived by CMP remain valid in a model with richer and more realistic dynamics like ours. We examine two shocks, namely (1) productivity gains in manufacturing, and (2) efficiency gains in creating new firms. In each shock simulation, we first introduce the intuitive explanation of responses described in CMP. Then, we show and explain the impulse responses in our model.

There exist several differences between simulations in this paper and those examined in CMP. First, as has been mentioned, the model is different. Our model contains much richer and more realistic dynamics such as one period lag for new firms to start operating, and nominal price and wage rigidities. Markups fluctuate in our model. Second, the model considered in CMP does not have any intrinsic dynamics. Hence, the direction of responses obtained in CMP is considered to be that of the steady-state (long-run) responses rather than the (short-run) dynamic responses, which reflect richer dynamics.

Regarding the first point, we will also show the simulation results of the flexible price counterpart.[6] Although there is no nominal rigidity in this

[5]As shown in Bilbiie, Ghironi, and Melitz (2005), regarding empirical problems associated with increasing returns to specialization and a constant elasticity of substitution production function, it may be better to model the household consuming a basket of goods defined over a continuum of goods. Neither specification, however, makes a difference in simulations conducted in this paper.

model, the welfare-based price level still fluctuates as the number of firms existing in the economy changes. Regarding the second point, the short-run as well as the long-run—namely the steady-state—effects of these shocks will also be demonstrated in each model simulation. The long-run simulation results simply show the steady-state changes in the endogenous variables for shifts in productivity and entry cost.

The variables of interest are, as in CMP, the number of firms, the home terms of trade,[7] and the real exchange rate, because we focus our attention on the international spillovers of production-enhancing technology via international relative prices. The terms of trade are on a firm basis. Therefore, it is the international relative price in the intensive margin. On the other hand, the real exchange rate is denominated by the welfare-based CPI index, which considers the taste for variety. Hence, this is considered to be the international relative price in the extensive margin. Because the terms of trade in this paper are defined as the ratio of import prices over export prices, the positive reaction means worsening of terms of trade. Similarly, the real exchange rate is defined in a standard manner where it is appreciating when it becomes smaller. In this paper, the numeraire is assumed to be the welfare-based CPI index but it is the nominal wage in CMP.

Productivity Gains in Manufacturing

According to CMP, from a microeconomic perspective, an increase in technology results in lower marginal costs. Therefore, each firm intends to lower its price so as to expand its sales share in the market. On the other hand, from a macroeconomic perspective, this intensified competition among firms lowers the profits of each firm because of the reduced price. If the wealth effect stemming from productivity gains in manufacturing on leisure is strong, increases in consumption demand become mild. As a result, profits are not sufficient to cover the entry cost, and some firms exit from the market. If such wealth effects are not very strong, several new firms consider it profitable to enter the market because the expected profit stream is large enough to cover the entry cost. As for the effects on the terms of trade, reflecting the lower price charged by the domestic firms, the terms of trade worsen in the domestic country reflecting the productivity gains. The domestic price level becomes lower because of the reduced marginal cost as well as the increased number of varieties, causing the welfare-based price level to decrease. Reflecting these developments in prices both in intensive and extensive margins, the real exchange rate depreciates. According to CMP, directions of responses of the foreign variables, namely the number of firms and the welfare-based price level, are ambiguous.

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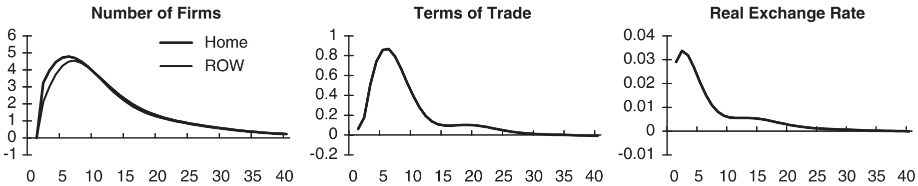
[6] We set the Rotemberg-type adjustment cost for nominal price and wage rigidities to zero.

[7] Note that the law of one price does not hold under the local currency pricing, the trade cost and sticky prices. In this paper we only show the terms of trade of the home country.

Figure 2. Sticky Price Model for Productivity Gains



Note: ROW = rest of world. Vertical axis is percentage deviation from steady state. Horizontal axis is period.

Figure 2 shows the dynamic responses for the productivity gains in manufacturing. The directions of responses are consistent with those obtained in CMP. In our model simulation, the increased technology in manufacturing causes profits to increase high enough for new firms to start business. As explained in CMP, the terms of trade deteriorate because the lowered marginal cost reduces the price of each domestic good. Because the size of the changes in the number of firms in domestic and foreign countries is similar, the real exchange rate depreciates reflecting the international relative prices in the intensive margin. The responses in a flexible price model are demonstrated in Figure 3. Because of the existence of one period lag for new firms to start operating, the number of firms increases only gradually as observed in the sticky price model. The directions of responses are consistent with those predicted in CMP. It now becomes more evident that because of the lowered marginal cost, the real exchange rate significantly depreciates for the domestic country. Figures 2 and 3 also show that increase in the number of firms is more dramatic in the sticky price model than in the flexible price model. This reflects more profit opportunities stemming from higher markups thanks to nominal rigidities.

We also checked the long-run directions of the responses for the productivity gain, namely, the changes in the steady state after the increase in technology, which are shown in Table 1. They turn out to be the same as those of the dynamic responses.[8]

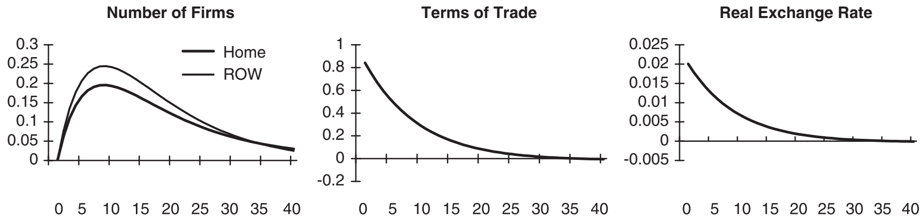
Table 1 summarizes the directions of responses in both CMP and our model. We can conclude that concerning the productivity gain, the results obtained in CMP hold in a model with much richer dynamics than the GEM.

Efficiency Gains in Creating New Firms (Goods)

Now, we will check the results obtained in CMP for the efficiency gain in creating new firms. CMP conclude that the lower fixed cost to entry

[8]Because nominal rigidities have no influence on the steady state, the long-run responses in this flexible price model are the same as the baseline model. Therefore, we will not mention the long-run responses in the flexible price model henceforth.

Figure 3. Flexible Price Model for Productivity Gains



Note: ROW = rest of world. Vertical axis is percentage deviation from steady state. Horizontal axis is period.

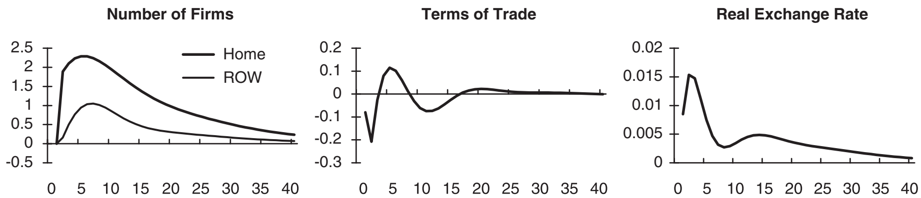
Table 1. Summary Table for Productivity Gains

	CMP	GEM Base		
		Short-run		Long-run
		Sticky	Flexible	
n	?	+	+	+
*n**	?	+	+	+
TOT	+	+	+	+
RER	+	+	+	+

Note: + = worsening of the terms of trade and depreciation of the real exchange rate. CMP = Corsetti, Martin, and Pesenti (2007); GEM = Global Economy Model; TOT = terms of trade; RER = real exchange rate.

naturally increases the number of the varieties in the domestic economy. As there is no direct changes in the real marginal cost stemming from the lowered entry cost, prices set by individual firms do not change very much. Under such circumstances, the terms of trade correspond one to one with nominal exchange rates. Increase in the number of firms reflecting the lower entry cost results in tighter demand conditions for domestic labor and consequently appreciates nominal exchange rates through the relative increase in domestic wages. "Hence, the Home terms of trade strengthen as the array of Home products increases" (CMP, p. 114). On the other hand, however, as the number of firms increases more in the domestic country, the welfare-based price index becomes smaller in the domestic country than in the rest of the world as the array of home products increases more. Reflecting these developments in prices, the real exchange rate, the nominal exchange rate denominated by the welfare-based price indices, namely, the international relative price in the extensive margin, depreciates for the domestic country.

Figure 4. Sticky Price Model for Efficiency Gains



Note: ROW = rest of world. Vertical axis is percentage deviation from steady state. Horizontal axis is period.

Figure 4 shows the responses in the sticky price model. All responses are almost consistent with the prediction by CMP. The number of firms increases, the terms of trade improve, but the real exchange rate depreciates because the welfare-based consumer price index (CPI) becomes smaller in the domestic country than in the foreign country, reflecting the dynamics of the number of varieties. The direction of the response of the terms of trade is not significant. It tends to fluctuate rather than improving as predicted by CMP. This reflects the fluctuations in the markup because of nominal rigidities.[9] As shown in Figure 5, the response of the terms of trade is exactly what is expected by CMP without such nominal rigidities. The fact that the real exchange rate initially appreciates in the flexible price model is somewhat puzzling. This implies that the appreciation in nominal exchange rates and the improvement of the terms of trade are very strong in this economy. Nominal exchange rates are appreciating thanks to the strong demand conditions in the domestic country. At the same time, a reduction in the entry cost induces investment for creating new firms, but, because of the existence of one period lag in firm dynamics, an immediate increase in the number of firms is prevented.[10] Therefore, adjustments through the extensive margin for the depreciation of the real exchange rate are hindered. As a result, the channel to increase the demand for investing in new firms becomes more significant soon after the shock hits the economy. According to this channel, a reduction in the entry fixed cost, as if it were a standard demand shock, raises the price set by each domestic firm and therefore further improves the terms of trade in the early stage of responses.[11]

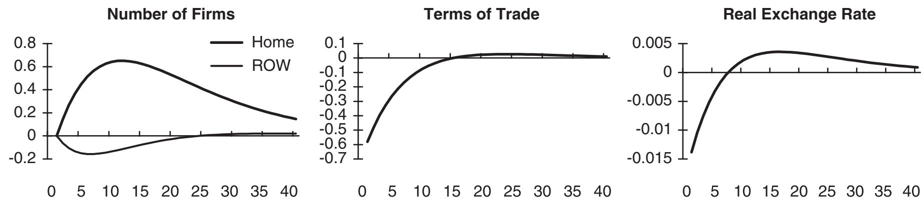
As for the long-run responses, Table 2 demonstrates that they are consistent with the prediction in CMP. However, responses of some variables

[9]As will be shown below, the elasticity of substitution between home and foreign products also affects the direction of the terms of trade.

[10]In particular, in the period when the shock hits the economy, the number of firms cannot increase.

[11]This implies that the one period lag in firm dynamics may also affect the dynamic responses for expanding the production frontier.

Figure 5. Flexible Price Model for Efficiency Gains



Note: ROW = rest of world. Vertical axis is percentage deviation from steady state. Horizontal axis is period.

Table 2. Summary Table for Efficiency Gains

	CMP	GEM Base		
		Short-run		Long-run
		Sticky	Flexible	
n	+	+	+	+
*n**	?	+	−	+
TOT	-	− to +	−	−
RER	+	+	− to +	+

Note: + = worsening of the terms of trade and depreciation of the real exchange rate. CMP = Corsetti, Martin, and Pesenti (2007); GEM = Global Economy Model; TOT = terms of trade; RER = real exchange rate.

in the long run differ from those in the short run because the wealth effect becomes more prevalent with the permanent shock.

Directions of both short-run and long-run responses are summarized in Table 2. The differences in directions between short-run and long-run responses and between our model and CMP suggest the importance of the degree of the wealth effects, and the dynamics embedded in our model such as one period lag in firm dynamics and the nominal rigidities.

This means that for a proper conduct of monetary policy, we need to check simulations not only from the theoretical model but also from a model with richer dynamics like the GEM.

III. Two-Country Example: Japan and the Rest of the World

In order to understand the realistic size of the international spillovers with more empirical relevance than in the previous analyses as above, this section first recalibrates our two-country model so that we can depict the relationship between Japan and the rest of the world. All parameters except for firm dynamics are taken from Laxton, N'Diaye, and Pesenti (2006).[12]

We will again show the responses for productivity gains in manufacturing and efficiency gains in creating new firms. All shocks occur in Japan not in the rest of the world.[13]

Figure 6 demonstrates the dynamic responses for the productivity gains in manufacturing when we give one percentage level shock to technology. As the theory or CMP predicts, qualitatively similar responses to those in Figure 2 are obtained. The increased technology in manufacturing causes profits to increase high enough for new firms to start business. The terms of trade deteriorate because the lowered marginal cost reduces the price of each domestic good. Because the welfare-based price level in the domestic country is also decreased mainly due to the reduced marginal cost, the real exchange rate also depreciates. For a 1 percentage point level shock to technology, terms of trade deteriorates by about 1 percent and the real exchange rate also depreciates by the same amount at their peaks. Because of the assumption about the country size, the responses in the rest of the world are not very large. Yet, the directions of the responses are the same both in Figures 2 and 6.

Figure 7 shows the responses for the efficiency gains in creating new firms. For a reduction in the fixed entry cost of 1 percentage point in creating new firms in Japan, the number of firms increases. This seems to result in the smaller welfare-based CPI in the domestic country than in the foreign country. These movements are mostly consistent with the theory that is explained in CMP. However, the direction of the terms of trade is ambiguous and the real exchange rate initially appreciates, in the model with more empirical relevance examined in this section. The terms-of-trade responses mostly reflect markup fluctuations. This again implies the importance of the nominal rigidities to understand the realistic responses of the international relative prices to shock. Owing to the nominal rigidities, namely the increase in the markup in this case, the number of firms increases massively in the domestic country reflecting more profit opportunities.

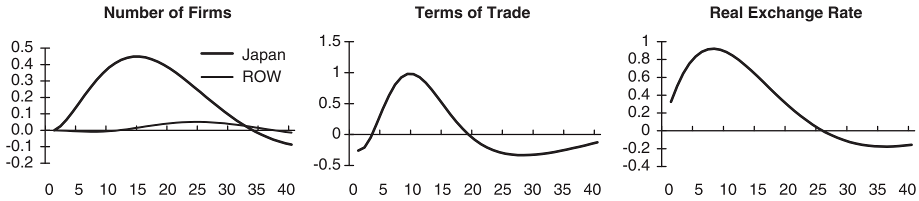
On the other hand, the reason for the appreciation of the real exchange rate is as follows. Although Japan is a small county compared with the rest of the world, the elasticity of substitution between domestic and foreign goods necessitates the increase in imports and therefore import-goods inflation reflecting tighter demand conditions, which also result in the ambiguous movements in the terms of trade.[14] Consequently, the real exchange rate denominated by the welfare-based CPI appreciates despite the increase in the number of firms in Japan. Overall, experiments with a model calibrated for the relationship between Japan and the rest of the world again show the

[12]For details, see our accompanying paper Fujiwara and Hirakata (2007).

[13]Responses for the shocks in the rest of the world are almost the mirror image of those for the shocks in Japan.

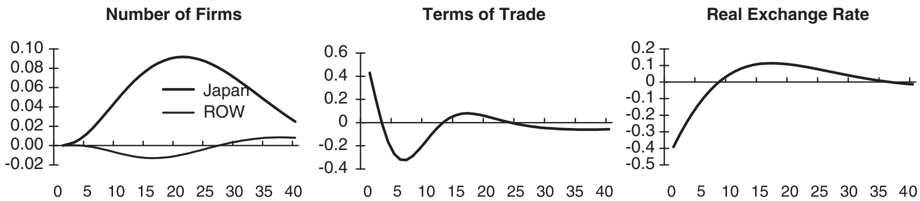
[14]We assume the local currency pricing in this model. Therefore, tighter demand conditions in the domestic country do not directly feed into the price setting in the foreign country.

Figure 6. Responses for Productivity Gains



Note: ROW = rest of world. Vertical axis is percentage deviation from steady state. Horizontal axis is period.

Figure 7. Responses for Efficiency Gains



Note: ROW = rest of world. Vertical axis is percentage deviation from steady state. Horizontal axis is period.

importance of nominal rigidities, and the persistence of shocks on the directions of responses of the international relative prices.

IV. Conclusion

This paper has shown how international relative prices respond to shocks that shift the production frontier outwards, namely, productivity gains in manufacturing and efficiency gains in creating new firms in a two-country model. For this purpose, contrary to the theoretical model used in CMP, we set up a model that contains richer and more realistic dynamics embedded in the GEM such as nominal price and wage stickiness. Our main conclusions are that real as well as nominal rigidities also alter the short-run responses by changing the responses of the number of firms and markups; and that persistence in shocks also matters for the determination of the direction of responses because it changes the size of the wealth effect. The nominal rigidities have significant effects on the price setting and therefore the markup determination, but the magnitude of the wealth effects is dependent on the persistence of the shocks. As a result, these factors are naturally considered to be very important determinants of the impulse responses. These results suggest that for a proper conduct of monetary policy, we need to check simulations not only from a theoretical model but also from a model with richer and more realistic dynamics like the GEM.

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