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Trade Openness and Growth: Pursuing Empirical Glasnost

ANDREAS BILLMEIER and TOMMASO NANNICINI*

Studies of the impact of trade openness on growth are based either on cross-country analysis—which lacks transparency—or case studies—which lack statistical rigor. This paper applies a transparent econometric method drawn from the treatment evaluation literature (matching estimators) to make the comparison between treated (that is, open) and control (that is, closed) countries explicit while remaining within a statistical framework. Matching estimators highlight that common cross-country evidence is based on rather far-fetched country comparisons, which stem from the lack of common support of treated and control countries in the covariate space. The paper therefore advocates paying more attention to appropriate sample restriction in cross-country macro research. [JEL C21, C23, F43, O57]


The relationship between trade openness or economic liberalization on the one hand, and income or growth on the other, is one of the main

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conundrums in the economics profession, especially when it comes to combining theoretical and policy-related findings with empirical findings. The theoretical advantages of trade for growth are known at least since Ricardo: international trade enables a country to specialize using its comparative advantage and benefit both statically and dynamically from the international exchange of goods. 1 From a policy perspective, the continuing efforts to liberalize international trade on a multilateral basis—first under the leadership of the General Agreement on Tariffs and Trade (GATT), now the World Trade Organization (WTO)—have contributed to better market access and rates of growth of international current account transactions much above worldwide economic growth. From an empirical point of view, however, the trade-growth link is still under discussion, both from a methodological angle and regarding the size and significance of the estimated effects.

Testing the empirical relevance of theoretical predictions in macroeconomics, growth theory, and political economics builds on cross-country evidence. In the attempt to detect correlations or causal relationships between aggregate variables, within-country variation is usually not sufficiently large to estimate the parameters of interest in a significant way, or it is so peculiar to the countries under consideration that the estimates do not hold more generally. At the end of the day, one must use some degree of cross-sectional variation to make inference on macro variables.

There is, however, widespread skepticism regarding the possibility of making sound inferences based on cross-country data. The empirical debate over the trade-growth nexus is a paradigmatic case. As Bhagwati and Srinivasan (2001) point out, both globalization supporters and foes rely on cross-country estimates, which dramatically suffer from specification problems, endogeneity, and measurement errors. According to them, cross-country regression estimates are completely unreliable, and robust evidence on the relationship between trade openness and growth “can come only from careful case studies of policy regimes of individual entries” (p. 19). Case studies, however, also suffer from apparent weaknesses as they lack statistical rigor and are exposed to arbitrary case selection. Instead of throwing out the baby (that is, cross-country statistical analysis) with the bath water, we propose to use case-study considerations as a sensitivity analysis of conventional cross-country estimates.

This paper evaluates the impact of a binary treatment—trade openness or economic liberalization—on the outcome—changes in per capita income. We

1Some theoretical models developed in the literature imply negative (or at least not necessarily positive) growth effects from trade; see the short discussion in Rodriguez and Rodrik (2001). By and large, however, macro theory has identified international exchange as a potential source of growth, theoretical exceptions being often associated with market failures that should be corrected by national policies different from protectionism; see, for example, the discussion in Bhagwati (2002). The December 2004 issue of the Journal of International Trade and Economic Development provides a recent review of the debate.
use microeconometric matching estimators from the treatment evaluation literature that are based on the same identifying assumption as ordinary least squares (OLS)—conditional independence; that is, the selection into treatment is fully determined by observable characteristics—to make the estimation procedure more transparent; in other words, to bring glasnost to muddied waters. In doing so, we are able to identify an additional weakness of cross-country estimates: we show that the country comparisons that lie behind simple cross-sectional results are often more than far-fetched.

Based on the analysis of these data-driven country comparisons, we argue that it is important to control for continent or macro-region dummies to make the matches more sensible. We also show, however, that this remedy may not be enough if open and closed countries are not evenly distributed across regions—that is, they lack common support. For example, for a prominent openness indicator from the literature (the Sachs-Warner-Wacziarg-Welch openness dummy, SWWW), developed countries should not be used to investigate the trade-growth link after 1980, as all countries in this group are open and do not provide the necessary within-group variation to estimate the counterfactual outcome in the case of no treatment.²

Cleaning the sample of countries outside the area of common support with respect to geographic areas and other important covariates, we confirm a positive and significant association between openness and growth within selected regions and after 1970. Using an alternative measure of trade barriers, we find instead inconclusive evidence, broadly confirming results in the literature that the SWWW is biased toward finding a positive effect of trade liberalization.

Matching estimates are of course subject to the same endogeneity issues of OLS, but this does not undermine the basic message of our results. Even estimators that deal more credibly with endogeneity—such as diff-in-diff panel strategies—suffer from the same common-support issue that we have identified as long as they rely to some degree on cross-sectional variation. This is the reason why, also in panel setups, more transparent estimators that control for a distributional overlap of treated and control countries in the covariate space should be preferred.

In the end, we wind up with the usual statistical trade-off between internal and external validity: although dropping countries outside the common support produces more sound statistical inference, these results cannot be extrapolated to make general statements that go beyond the sample effectively used in the estimation. In other words, it is unlikely that the effect of openness on growth can be robustly estimated for a worldwide sample of countries, casting doubt on much of the cross-country growth literature that strives to cover an ever-increasing set of countries.

²See Section II for a description of the SWWW indicator.
I. Literature Review

Empirical Studies on Trade Openness and Growth

Providing conclusive empirical evidence on the intuitively positive causal effect of trade on growth has been a challenging endeavor, complicated by a multiplicity of factors; see, for example, Winters (2004) for an overview. Most of the literature has used cross-country evidence that suffers from numerous shortcomings, related to both the measurement of openness and econometric modeling.

Following Barro’s (1991) seminal paper on growth regressions, several prominent cross-country studies established a positive link between trade openness and growth; these studies include Dollar (1992), Sachs and Warner (1995), and Edwards (1992, 1998). Similarly, Vamvakidis (2002) finds, in a historical context, evidence that trade is associated positively with growth after 1970, but not before. In a Stern Review of the cross-sectional literature on trade and growth, Rodriguez and Rodrik (2001) criticize the choices of openness measure and weak econometric strategies. They find little evidence that open trade policies as measured in the aforementioned contributions are significantly associated with economic growth once they correct for the weaknesses they point out.3 Harrison (1996) shows that most of the explanatory power of the composite openness dummy assembled in Sachs and Warner (1995) comes from the nontrade components of this measure.

DeJong and Ripoll (2006) take up one of the suggestions voiced in Rodriguez and Rodrik (2001) and construct an alternative measure of direct trade barriers—ad valorem tariff rates—that is arguably more immune to the Rodriguez-Rodrik critique. They find that the relationship between trade barriers and income is nonlinear for a panel of 60 countries. In particular, the correlation between trade barriers and income is negative for rich countries but positive (albeit statistically weaker) in poorer countries. Salinas and Aksoy (2006), however, criticize this indicator on the grounds that it is not a rigorous representation of the tariff structure and that it does not capture other trade barriers, such as nontariff barriers—in fact they find a low correlation between cross-country measures of unweighted average tariffs and the frequency of nontariff barriers.

From a methodological perspective, deep skepticism has been brought to bear against cross-country evidence on the trade-growth issue. In addition to the citation in the previous section, Bhagwati and Srinivasan (2002, p. 181) point out that “cross-country regressions are a poor way to approach this question” and that “the choice of period, of the sample, and of the proxies, will often imply many degrees of freedom where one might almost get what one wants if one only tries hard enough!” Levine and Renelt (1992) and Temple (2000) apply extreme-bounds analysis to show that the results of

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cross-country growth regressions are not robust to even small changes in the conditioning information set (that is, right-hand side variables).

Focusing on identification issues, cross-country studies suffer from two major weaknesses: reverse causality (that is, liberalized trade causes higher economic growth versus more trade being the result of economic growth) and endogeneity (for example, country-specific omitted characteristics affecting both openness and growth). Dealing with endogeneity has triggered a substantial amount of interest in the use of instrumental variables (IV). This family of models suggests using regressors that have an impact on openness, but are uncorrelated with income. Using gravity models, Frankel and Romer (1999) and Irwin and Tervio (2002) find a positive effect running from trade to growth by isolating geographical components of openness that are assumed independent of economic growth, including population, land area, borders, and distances. But even these presumably exogenous instruments could have indirect effects on growth, thereby biasing the estimates. Dollar and Kraay (2003) suggest estimating the regressions in differences and using lagged openness as instrument. However, the simultaneity bias in the trade-growth context could extend over time—trade today may depend on growth tomorrow via imports for investment purposes—and using lagged variables as instruments is unlikely to fully correct for the bias.

As an alternative approach to classic IV, Lee, Ricci, and Rigobon (2004) use identification through heteroskedasticity in a panel framework, and find that openness has a small, positive, but not particularly robust effect on growth. They have to rely, however, on the nontestable assumption that the structural shocks in the system of simultaneous equations are uncorrelated. Using the same technology, Rigobon and Rodrik (2005) find that trade openness (defined as the trade share in GDP) has a significant negative effect on income.

Another strand in the trade and growth literature seeks to improve upon cross-country regressions by employing panel techniques, geared at controlling for (time-invariant) unobservable country effects. An early example is Harrison (1996), who uses fixed-effect estimators and finds a stronger impact of various openness indicators in a panel setup compared with standard cross-country regressions. Wacziarg and Welch (2003) further the discussion in the literature in three directions: they update, expand, and correct the trade openness indicator in Sachs and Warner (1995); they show that the Sachs and Warner (1995) results of a positive effect of trade on growth break down if extended to the 1990s in a cross-sectional setup; and they provide evidence in a panel context that, even in the 1990s, there is a positive effect of trade on growth when the analysis is limited to

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4In fact, Irwin and Tervio (2002) and Rodriguez and Rodrik (2001) find that geographical latitude has a significant effect on growth, casting doubt on the identifying assumption used by Frankel and Romer (1999). Furthermore, these instruments relate primarily to trade volumes, not trade policies, as discussed by Rodriguez and Rodrik.
within-country effects. Slaughter (2001) uses a diff-in-diff approach to infer the effect of four very specific trade liberalization events on income growth dispersion and finds no systematic link between trade liberalization and per capita income convergence. Giavazzi and Tabellini (2005) also apply a diff-in-diff approach to study the interactions between economic and political liberalizations. They find a positive and significant effect of economic liberalization on growth, but they claim that this effect cannot be entirely attributed to international trade, as liberalizations tend to be accompanied by other policy improvements.

Empirical Studies Applying Matching Estimators to Macro Data

A limited, but growing, strand of aggregate empirical literature—particularly in political economics—applies microeconometric estimators developed in the treatment evaluation literature to cross-country data to overcome the weaknesses of OLS in cross-sectional setups. Persson and Tabellini (2003) use propensity-score matching methods to estimate the effects of political institutions (proportional against majoritarian electoral rule; presidential against parliamentary regime) on a set of relevant economic variables. Edwards and Magendzo (2003) apply matching estimators to analyze the macroeconomic record of dollarized economies. Atoyan and Conway (2006) use matching estimators to evaluate the impact of IMF programs.

All these studies point to the fact that nonparametric (or semiparametric) matching estimators allow the OLS linearity assumption to be relaxed. This is not their only merit, however, as the linearity assumption can be also relaxed in the OLS framework by specifying a fully saturated model. The major advantage of matching techniques is that they allow the researcher to carefully check for the existence of a common support in the distributions of treated and control units across the covariate space. And this advantage can be even greater in a small sample of countries, as the “matched” treated and control units can be easily identified. This “transparency” attribute of matching estimators is described and exploited in Section III with respect to the estimated effect of trade openness on growth.

II. Data and Variables of Interest

Under the trade and growth umbrella, a whole set of relationships have been analyzed in the literature. As the dependent variable, GDP levels, changes, GDP per capita, and relative incomes (or dispersion thereof) have been used as outcome measures, mainly to distinguish between level, growth, and convergence effects. We employ the difference of (log) per capita GDP, as we are interested in the dynamic impact of trade openness over time, not only in its one-off effects on the individual income level.

5Wacziarg and Welch (2003) essentially conduct difference regressions in growth, or diff-in-diff regressions in log income.
For trade and openness, two major groups of indicators have emerged in the literature, addressing somewhat different questions. On the one hand, there are simple measures of trade volumes that are particularly subject to endogeneity problems (especially if normalized by GDP), and have in fact been used within an IV framework (for example, Frankel and Romer, 1999; Rigobon and Rodrik, 2005). On the other hand, there have been repeated efforts to identify the impact of trade policy and lower trade barriers on economic growth. To this end, a variety of indicators have been constructed, the most notable among them being the binary indicator by Sachs and Warner (1995), extended, updated, and revised by Wacziarg and Welch (2003); short SWWW. According to this indicator, a country is considered closed to international trade in any given year if at least one of the following conditions is satisfied: (i) average tariffs exceed 40 percent; (ii) nontariff barriers cover more than 40 percent of its imports; (iii) it has a socialist economic system; (iv) the black market premium on the exchange rate exceeds 20 percent; and (v) much of its exports are controlled by a state monopoly. A country is open if none of these conditions applies. As our binary indicator of openness—or economic liberalization in the language of Giavazzi and Tabellini (2005)—we use the SWWW trade openness policy dummy. Our main treatment indicator thus intends to capture policy changes that reduce the constraints on market operations below a critical threshold along these five dimensions.

In the third part of Section III, we repeat the analysis to the extent feasible with an alternative trade barrier indicator suggested by Rodriguez and Rodrik (2001) and applied in DeJong and Ripoll (2006). This annual indicator is available for the period 1975 until 2000 and represents, in the words of DeJong and Ripoll (2006, p. 630), “ad-valorem tariffs, measured using import duties as a percentage of imports, as reported by the World Bank.” This indicator is, hence, essentially a subindicator of SWWW—corresponding to (i) above—but with an additional degree of freedom: to obtain a binary treatment consistent with our treatment evaluation framework, we need to split the country sample according to the degree of protection by choosing an appropriate tariff threshold to indicate whether a country is open or closed.

To anchor our results in the existing literature, we draw on two data sets used recently in a related context. Vamvakidis (2002) presents historical evidence of the connection between openness and growth over the period 1870–1990; we focus on the post-1950 part of his data set. The data set

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6For a comparison of various indicators, see Harrison (1996).
7The SWWW dummy captures, in fact, more than just openness to trade, for example, also the socialist origin. Nevertheless, we base our initial analysis on this dummy, given the prominence it has achieved in the literature. Sachs and Warner (1995, p. 25) note that the socialism indicator serves as a proxy for central planning, which could be viewed as a substitute for overt trade policies such as tariffs.
8We are grateful to an anonymous referee for this suggestion.
9See Vamvakidis (2002) for a detailed description of the data sources.
consists of repeated country cross-sections for the intervals 1950–70, 1970–90, and 1990–98. Besides the average GDP per capita growth and the openness dummy, the data set contains information on the initial GDP, investment share, population growth, secondary school enrollment, inflation, and black market premium.

The other data set we use has been analyzed in Giavazzi and Tabellini (2005) and Persson and Tabellini (2006). Of this very rich panel data set covering about 180 countries over the period 1960–2000, we use decade averages (1961–70, 1971–80, 1981–90, and 1991–2000) only for a few variables that are related to the question at hand: the updated SWWWW, the log change in per capita GDP, and the same control variables mentioned above (with the only two exceptions that inflation is not reported, while a democracy dummy is present).

III. Matching Estimators and Cross-Country Analysis

Methodology

The common aim of most of the empirical studies reviewed in Section I is to assess whether a pro-openness trade policy has a causal effect on either the level or the growth rate of GDP. This problem of inference involves “what if” statements and thus counterfactual outcomes. Hence, it can be translated into a treatment-control situation and analyzed within Rubin’s (1974) potential-outcome framework for causal inference. The essential feature of this approach is to define the causal effect of interest as the comparison of the potential outcomes for the same unit measured at the same time: \( Y(0) = (\text{the value of GDP growth } Y \text{ if the country is exposed to treatment } T = 0, \text{ that is, if it is closed to trade}), \) and \( Y(1) = (\text{the value of GDP growth if the same country is exposed to treatment } T = 1, \text{ that is, it is open to trade}). \) Only one of these two potential outcomes can be observed—specifically, the one corresponding to the treatment the country received—but the causal effect is defined by their comparison, that is, \( Y(1) - Y(0). \) This highlights that estimating the causal relationship between \( T \) and \( Y \) is hampered by a problem of missing data—the counterfactual outcomes \( Y(0) \) for open countries and \( Y(1) \) for closed countries.

In this setting, the aim of statistical analysis is usually that of estimating some features of the distribution of \( Y(1) - Y(0), \) like

\[
E[Y(1) - Y(0)],
\]

which is called the average treatment effect (ATE). Alternatively, one can be interested in the ATE for the subpopulation of the treated observations:

\[
E[Y(1) - Y(0)|T = 1],
\]

which is called the average effect of treatment on the treated (ATT). In the present context, the ATE corresponds to the counterfactual question: what would have been the growth rate of the countries in our sample had they...
decided to switch their trade regime? On the contrary, the ATT focuses on the counterfactual question for treated units only: what would have been the growth rate of open countries had they decided to close their economies?

Problems for the identification of these ATE may arise from the existence of country-specific unobservables affecting both the two potential outcomes (or just one of them) and the treatment indicator. The fact that the treatment might be endogenous reflects the idea that the outcomes are jointly determined with the treatment, or that there are omitted confounders related to both the treatment and the outcomes. One of the assumptions that allow the identification of the ATE is the “unconfoundedness” condition, also referred to as “selection on observables” or “conditional independence assumption,” which is the rationale behind common estimation strategies such as regression modeling and matching.\(^{10}\) This assumption considers the conditioning set of all relevant pretreatment variables \(X\) and assumes that

\[
Y(1), Y(0) \perp T | X
\]

\(0 < \Pr(T = 1 | X) < 1.\) \(^{(4)}\)

That is, conditioning on observed covariates \(X\), the treatment assignment is independent of potential outcomes.\(^{11}\) Unconfoundedness says that treatment assignment is independent of potential outcomes after accounting for a set of observable characteristics \(X\). In other words, exposure to treatment is random within cells defined by the variables \(X\).

Under unconfoundedness, one can identify the ATE within subpopulations defined by \(X\):

\[
E[Y(1) - Y(0) | X] = E[Y(1) | T = 1, X] - E[Y(0) | T = 0, X],
\]

and also the ATT as

\[
E[Y(1) - Y(0) | T = 1, X] = E[E[Y(1) | T = 1, X] -
E[Y(0) | T = 0, X] | T = 1],
\]

where the outer expectation is over the distribution of \(X\) in the subpopulation of treated units. In other words, thanks to unconfoundedness, one can use the observed outcome of treated (control) units, conditional on \(X\), to estimate the counterfactual outcome of control (treated) units.

An implication of the above results is that, if we could divide the sample into cells determined by the exact values of the variables \(X\), then we could just

\(^{10}\)See Imbens (2004) for a review of nonparametric estimation methods under this assumption.

\(^{11}\)To identify the ATT, a weaker version of these conditions suffices: \(Y(0) \perp T, X\) and \(\Pr(T = 1 | X) < 1.\)
take the average of the within-cell estimates of the ATE. Often the variables \( X \) are continuous, so that smoothing techniques are needed; under unconfoundedness several estimation strategies can serve this purpose. Regression modeling and matching are viable alternatives, which rely on the same identification condition. The main advantage of matching with respect to linear regression is that the latter obscures information on the distribution of covariates in the two treatment groups. In principle, one would like to compare countries that have the same values of all covariates; but unless there is a substantial overlap between the two covariates distributions, a regression model relies heavily on model specification—that is, on extrapolation—for the estimation of (treatment) effects. It is thus crucial to check how much the distributions of the treated and control units overlap across covariates, and which is the region of common support for the two distributions.

In contrast to other studies that apply the propensity-score version of matching to macro data—see Persson and Tabellini (2003)—we implement the above strategy by using the “nearest neighbor” algorithm for covariate matching.\(^\text{12}\) Matching estimators impute the country’s missing counterfactual outcome by using average outcomes for countries with “similar” values of the covariates. The nearest neighbor algorithm uses the following simple approach to estimate the pair of potential outcomes. The potential outcome associated to the treatment that country \( A \) received is simply equal to the observed outcome of \( A \). The potential outcome associated to the treatment that country \( A \) did not receive is equal to the outcome of the nearest country that received the opposite treatment (country \( B \)), where “nearest” means that the vector of covariates of \( B \) shows the smallest distance from the vector of covariates of \( A \) according to some predetermined distance measure.

Formally, define \( \|x\|_V = (x'Vx)^{1/2} \) as the vector norm with positive definite weight matrix \( V \), and let \( \|x-z\|_V \) be the distance between vectors \( x \) and \( z \).\(^\text{13}\) Let \( d(i) \) be the smallest distance from the covariates of country \( i \), \( X_i \), with respect to the covariates of all other countries with the opposite treatment. Allowing for the possibility of ties, define \( J(i) \) as the set of indices for the countries that are at least as close to country \( i \) as its nearest neighbor:

\[
J(i) = \{ k = 1, ..., N | T_k = 1 - T_i, \|X_k - X_i\|_V = d(i) \}. \quad (7)
\]

The pair of potential outcomes for country \( i \) are estimated as

\[
\hat{Y}_i(l) = Y_i \quad \text{if} \quad T_i = l \quad (8)
\]

\(^{12}\)See Abadie and others (2004) for a description of this algorithm and the program that implements it in Stata.

\(^{13}\)Following Abadie and others (2004), we let \( V \) be the diagonal matrix with the inverses of the variances of the covariates on the main diagonal. All the estimates presented in this section are robust to the utilization of a different distance metric—the Mahalanobis distance suggested by Rubin (1980).
\[ \hat{Y}_i(l) = \frac{1}{\# J(i)} \sum_{k \in J(i)} Y_k \quad \text{if} \ T_i = 1 - l, \]  

(9)

where \( \# J(i) \) is the numerosity of the set \( J(i) \). The ATE and ATT are thus estimated as

\[ \tau_{ATE} = \frac{1}{I} \sum_{i=1}^{I} [\hat{Y}_i(1) - \hat{Y}_i(0)] \]  

(10)

\[ \tau_{ATT} = \frac{1}{I_T} \sum_{i=1}^{I_T} [Y_i - \hat{Y}_i(0)], \]  

(11)

where \( I \) and \( I_T \) are the sample size and the number of treated countries, respectively. These nearest-neighbor matching estimators both allow for the identification of the ATE and ATT under unconfoundedness and are fully transparent, as the list of country matches underlying, the results can be displayed in small samples (see the next two subsections and the online appendix).

Summing up, applying matching estimators to (small) cross-country samples comes with a disadvantage and an advantage. The disadvantage is that unconfoundedness is unlikely to hold, as it is often implausible to assume that country-specific unobservable characteristics do not play any role in treatment assignment. The advantage is that they allow us to transparently check for the existence of common support. Consequently, matching estimators are not used in this section as a magic bullet able to produce more reliable estimates than regression modeling, as both estimation strategies rest on the same identification condition and are therefore subject to the same specification problems. They are used, instead, to highlight the country comparisons that are behind cross-sectional results, to assess their plausibility, and to check whether the distributions of treated and control countries display sufficient overlap in the covariate space. After these steps, the cross-sectional results are improved by restricting the estimates to the region of common support. Even though these refined results must also rely on the conditional independence assumption, their plausibility can be further assessed by a careful inspection of the new country matches produced by the nearest neighbor algorithm. In other words, as the estimation process is no longer a black box but based on a transparent match of countries, case-study considerations along the lines of Bhagwati and Srinivasan (2001) can be introduced to assess the robustness of the results.

**The Unbearable Lightness of Cross-Country Estimates**

We now turn to the data sets introduced in Section II and apply matching estimators to shed light on the country comparisons underlying cross-sectional estimates.
Table 1 presents results for the Vamvakidis (2002) data set. We confirm his results that openness—as represented by the Sachs and Warner (1995) dummy—has a significant effect on growth after 1970, but not before. The coefficients indicate that an open country grows, on average, by 1.5 to 2 percentage points per year faster than a closed economy. The results for both types of matching estimates, ATE and ATT, are qualitatively and quantitatively similar to the standard OLS results. The estimates are robust to the introduction of regional dummies among control variables. Unfortunately, the data set comes with several drawbacks: (i) the data are pooled for 20-year intervals; (ii) the information stops in 1998, too early to meaningfully capture the countries of the former Soviet Union territory; and (iii) the sample size is very small in the 1950s and 1960s—mainly developed and Latin American countries.

In Table 2, we repeat the exercise switching to the Persson and Tabellini (PT) data set. This data set contains more countries and extends until 2000, using the Wacziarg and Welch (2003) update of the Sachs-Warner dummy. We produce pooled estimates by decades for the whole data set. Again, the matching results are very similar to the OLS results, and we find a significant effect of trade on growth for the 1990s and 1970s. In these decades, open countries grew on average by 1.5 to 2 percentage points faster than closed countries, whether we control for regional dummies or not. The growth effect of openness is not significantly different from zero in the 1980s and 1960s.

So far, matching does not add anything to OLS results, as the estimates are very similar and based on the same identification assumption. We now turn to the transparency advantage of the nearest-neighbor matching estimator to reveal the country comparisons underlying the estimates from the PT data set. Tables 3 and 4 display the full list of treated (that is, open to trade for more than half of the decade) and control (that is, closed) countries and their nearest neighbors in the ATE estimation for the 1990s. The online appendix contains the full lists of treated and control countries and their nearest neighbors for the 1990s, 1980s, 1970s, and 1960s. In all of these tables, the first column (Country) indicates the country under consideration; the second column (Baseline) shows the nearest neighbor used to estimate the counterfactual outcome of the country in the first column for the ATE estimation without area dummies; the third column (Area) shows the nearest neighbor used to estimate the counterfactual outcome of the country in the first column for the ATE estimation with area dummies. For example, the Baseline matches in Tables 3 and 4 for the 1990s are the country comparisons underlying the 1.505 coefficient in Table 2 (that is, the effect of openness on growth without controlling for area dummies), whereas the Area matches in

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14Online appendix available at http://www.tommasoannicini.eu/Portals/0/trade_matching_only_appendix.pdf. Tables 1 through 8 display the matches for the PT data set.
the same tables lie behind the 1.318 coefficient in Table 2 (that is, the effect of openness after controlling for area dummies).

Tables 3 and 4 indicate that a few Baseline matches appear to work reasonably “well”—for example, in the 1990s for Bulgaria and Egypt, which are matched with Ukraine and Algeria, respectively. Arguably, this intuitive appreciation is based on the implicit assumption that there are region-specific unobservable effects, for example, a common language, colonization, level of development, geographic proximity, or legal origin. For others—for example,
Albania and Sri Lanka, which are matched with Central African Republic and Algeria, respectively—the matches are somewhat less meaningful. In particular, all (treated) developed countries give rise to very poor matches (for example, Italy and the United Kingdom with Russia, or the United States and Canada with China). In other words, most of the baseline matches do not appear robust to area-specific unobservables.

Therefore, we construct country groups that may capture some of these area-specific unobservables. We divide the world into six groups: Africa, Asia, Latin America, Middle East, developed economies, and transition economies.

### Table 2. Openness and Growth, Cross-Country Evidence (II), 1961–2000

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Source: Authors’ calculations based on data in Persson and Tabellini (2006).

Note: Dependent variable ($Y$): real GDP per capita growth. Treatment indicator ($T$): trade openness dummy (Sachs and Warner, 1995; Wacziarg and Welch, 2003). Control variables ($X$) include initial GDP per capita, secondary school enrollment, population growth, and investment share. Area dummies refer to Africa, Asia, Latin America, Middle East, developed countries, and transition economies. ATE and ATT stand for average treatment effect and average treatment effect on the treated, respectively, and are estimated by nonparametric nearest-neighbor matching. *corresponds to 5 percent significance level; **corresponds to 1 percent significance level.
Table 3. Cross-Country Matches, Treated Countries, 1991–2000

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Source: Authors' calculations based on data in Persson and Tabellini (2006).

Note: Baseline and Area refer to the nearest-neighbor match without and with area dummies, respectively, in the ATE estimation (see Table 2). Refined refers to the nearest-neighbor match without Latin America and developed countries in the ATE estimation (see Table 5). C.A.R. and P.N.G. stand for Central African Republic and Papua New Guinea, respectively.
economies (where being a developed country takes precedence over geographic region). The matches underlying the estimation with these area dummies are reported in the Area column. For Albania and Sri Lanka in the 1990s, this step appears to work reasonably well, as they are now matched with Belarus and Pakistan, which are certainly perceived as more similar than the baseline-nearest neighbors. There are, however, certain surprising findings: for example, all developed countries are matched with Iceland! A similar result obtains for Latin America, where Venezuela is the only control that is picked to be a match. This is due to the fact that Iceland and Venezuela are, according to the SWWWW classification, the only closed economies in the group of developed countries and Latin America in the 1990s. In other words, there is no common support between treated and control countries in those two regions. Introducing area dummies is not dropped

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<td>Iran</td>
<td>Argentina</td>
<td>Tunisia</td>
<td>Tunisia</td>
<td>Zimbabwe</td>
<td>Ecuador</td>
<td>Zambia</td>
<td>Ghana</td>
</tr>
</tbody>
</table>

Source: Authors’ calculations based on data in Persson and Tabellini (2006).

Note: Baseline and Area refer to the nearest-neighbor match without and with area dummies, respectively, in the ATE estimation (see Table 2). Refined refers to the nearest-neighbor match without Latin America and developed countries in the ATE estimation (see Table 5). C.A.R. and P.N.G. stand for Central African Republic and Papua New Guinea, respectively.

15 We label as developed all countries that joined the Organization for Economic Cooperation and Development (OECD) between its foundation in 1961 and 1973—the end of the initial participation wave that concluded with the accession of New Zealand. In addition, we add Cyprus and Israel for lack of better options. Countries that joined the OECD more recently (starting with Mexico in 1994) are allocated to their geographic region. The label transition is used for all countries in central and eastern Europe that are contained in the sample, including for the period before 1990.

16 See the online appendix mentioned in footnote 14 for a precise list of treated and untreated countries across regions.
enough to control for area-specific unobservables, unless there is a sufficient overlap of treated and untreated countries in all areas.

Summing up, the matches listed in the Baseline and Area columns of Tables 3 and 4 for the 1990s (and in the online appendix for the other decades) show that country comparisons underlying cross-country analysis are often more than far-fetched. This unbearable lightness of cross-country analysis extends from matching to other cross-sectional estimators that rely on the unconfoundedness assumption, such as plain regression modeling. This is due to the fact that OLS estimates are based either on the same implicit but far-fetched country comparisons or—even worse—on parametric extrapolation beyond the region of common support. In fact, if treated and control countries are very different from each other with respect to covariates, the OLS estimate of the counterfactual outcome of the treated is constructed by linearly extrapolating the observed outcome of control units, and vice versa.

**Refined Evidence in Selected Samples**

The above discussion shows that—as long as we want to control for area-specific unobservable characteristics—we should restrict the analysis of the trade-growth nexus to regions with enough treatment variation. In other words, to improve the quality of the country matches underlying the results, we should drop regions with no common support between treated and control units.

In Table 5, we reestimate the pooled specification eliminating countries that lack common support with respect to regional affiliation. As a criterion, we establish that the ratio between treated and control countries (or vice versa) should be below 10; that is, for every 10 treated countries, we require more than 1 potential control country (or vice versa). As shown in the online appendix, this requirement eliminates the group of developed economies and Latin America in the 1990s because (almost) all of them are open. The same holds for other regions and other decades, however: in the 1980s, almost all developed economies are already open. In the 1970s and 1960s, almost all African economies (except Mauritius in the 1970s) are closed according to the SWWW dummy. Moreover, transition economies are excluded prior to the 1990s, as none of them is in fact in transition (that is, open).

Table 5 reports matching estimates—of both the ATE and ATT—restricted to countries that meet the common-support condition for geographic areas. Comparing these estimates to the previous ones for the unrestricted sample (Table 2), the coefficients appear to be slightly more significant and also larger in magnitude in the 1990s and 1970s. Moreover, we now find stronger evidence of a marginally significant positive effect of openness on growth in the 1980s, especially for the countries that were open to trade (ATT estimate). All the estimated effects lie in the 1.5 to 2.5 percentage point range. For the 1960s, again, the coefficients are never
significantly different from zero. The *Refined* column in Tables 3 and 4 reports the country matches underlying these results. Counterintuitive matches are now considerably reduced.

We conclude from this exercise that it is important to check for the existence of common support. In fact, in small samples of countries, the advantage of matching estimators lies in the guidance for appropriately restricting the analysis to specific subsamples. Unlike the estimates in Table 2—which replicate common results from growth regressions in the literature—the estimates presented in Table 5 fully control for area-specific unobservables and are based on more plausible country comparisons. There is no free lunch, however, as the external validity of the estimates is now

<table>
<thead>
<tr>
<th>Table 5. Openness and Growth, Refined Evidence (I), 1961–2000</th>
</tr>
</thead>
<tbody>
<tr>
<td>ATE: $E[Y(1) - Y(0)</td>
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</thead>
<tbody>
<tr>
<td><strong>Matching without area dummies</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Estimate (ATE)</td>
<td>1.921**</td>
<td>0.949</td>
<td>1.841**</td>
<td>0.084</td>
</tr>
<tr>
<td>SE</td>
<td>(0.647)</td>
<td>(0.706)</td>
<td>(0.634)</td>
<td>(0.666)</td>
</tr>
<tr>
<td>Estimate (ATT)</td>
<td>1.760**</td>
<td>2.358*</td>
<td>2.229**</td>
<td>0.613</td>
</tr>
<tr>
<td>SE</td>
<td>(0.633)</td>
<td>(0.959)</td>
<td>(0.625)</td>
<td>(0.915)</td>
</tr>
<tr>
<td><strong>Matching with area dummies</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Estimate (ATE)</td>
<td>1.389*</td>
<td>0.809</td>
<td>2.240**</td>
<td>−0.143</td>
</tr>
<tr>
<td>SE</td>
<td>(0.708)</td>
<td>(0.526)</td>
<td>(0.684)</td>
<td>(0.680)</td>
</tr>
<tr>
<td>Estimate (ATT)</td>
<td>0.849</td>
<td>1.252*</td>
<td>1.986*</td>
<td>0.544</td>
</tr>
<tr>
<td>SE</td>
<td>(0.738)</td>
<td>(0.524)</td>
<td>(0.770)</td>
<td>(0.924)</td>
</tr>
</tbody>
</table>

|                      |           |         |         |         |
| Africa               | Yes (19–14) | Yes (5–30) | No (1–33) | No (0–35) |
| Asia                 | Yes (9–4)  | Yes (6–8)  | Yes (6–8) | Yes (4–10) |
| Latin America        | No (19–1)  | Yes (6–17) | Yes (4–19) | Yes (5–18) |
| Middle East          | Yes (6–3)  | Yes (2–6)  | Yes (1–7) | Yes (1–7) |
| Developed countries  | No (25–1)  | No (24–2)  | Yes (21–4) | Yes (21–4) |
| Transition economies | Yes (9–3)  | No (0–3)   | No (0–3)  | No (0–1)  |
| Treated              | 43        | 19       | 32       | 31       |
| Controls             | 24        | 61       | 38       | 39       |
| Observations         | 67        | 80       | 70       | 70       |

Source: Authors’ calculations based on data in Persson and Tabellini (2006).

Note: Dependent variable ($Y$): real GDP per capita growth. Treatment indicator ($T$): trade openness dummy (Sachs and Warner, 1995; Wacziarg and Welch, 2003). Control variables ($X$) include initial GDP per capita, secondary school enrollment, population growth, investment share, and area dummies (as indicated). The two numbers in parentheses after each area refer to the number of treated and control countries, respectively. Samples restricted to certain areas to meet the common-support condition for area dummies. ATE and ATT stand for average treatment effect and average treatment effect on the treated, respectively, and are estimated by nonparametric nearest-neighbor matching. *corresponds to 5 percent significance level; **corresponds to 1 percent significance level.
reduced. The results recommend refraining from commenting on the effect of trade openness on growth in developed countries after 1980, in Africa before 1980, and in Latin America after 1990.

The estimates in Table 5 control for the existence of common support with respect to a set of covariates that we deem important to capture unobservable regional characteristics associated to geography, level of development, culture, or legal origins—that is, area dummies for Africa, Middle East, Asia, Latin America, transition economies, and developed economies. The common support, however, should also be checked for other covariates. In principle, we would like to match countries that are very similar with respect to all covariates, but this is impossible if treated and control units are not evenly distributed across all the ranges of variation of covariates. Figures 1–8 show that, for example, this condition is not often met for investment share and secondary school enrollment. These figures report the kernel density of treated and control countries over the ranges of variation of these two variables. For instance, Figure 5 shows that the common support for investment share in the 1970s ranges from 0.11 to 0.39, with 27 (control) countries below this region and 1 (treated) country above. To meet the common-support condition, these 28 countries should be dropped from the estimation sample.

Table 6 reports matching estimates for samples restricted to the regions of common support identified in Figures 1–8. This evidence is consistent with the one described in Table 5. When carefully matching only countries that lie in the common support, cross-sectional estimates using the SWWW dummy

![Figure 1. Common Support for Investment Share, 1991–2000](image_url)

Source: Authors’ calculations based on data in Persson and Tabellini (2006).
Note: Kernel density of investment share in the 1990s for both treated and control observations. Treated countries: 87. Control countries: 26. All countries in common support.
continue to detect a positive and significant association between openness and growth in the 1970s, 1980s, and 1990s, but not in the 1960s.

Finally, we are aware that—even though our refined estimates improve the internal validity of cross-country results by checking for common
support—our results still suffer from the fact that country-specific unobservables (that is, endogenous selection into treatment) might violate the conditional independence assumption. By the same token, if conditional independence is not met, matching estimates should not be interpreted as...
causal effects, because the direction of causality is unclear. The main message of this paper is somewhat different, however: matching estimators clarify the importance of controlling for common support in the covariate space, especially in macro samples, as they tend to be comparatively small.
This point extends to state-of-the-art estimation techniques. Panel methods, for example, can overcome some of the OLS weaknesses by using within-country (that is, time-series) information to control for unobservable time-invariant country characteristics. However, as long as they still use some cross-sectional variation—as does, for example, the diff-in-diff estimator—they still suffer from insufficient transparency that could hide a lack of common support.

More Evidence Using an Alternative Indicator of Trade Restrictions

The trade restriction measure compiled by DeJong and Ripoll (2006) consists of annual observations from 1975 to 2000 of ad valorem tariffs—import duties as a share of imports—for 74 countries. To match the structure of our 10-year pooled data set, we discard the beginning of the sample (which also displays a number of missing observations) and use the averages for 1981–90 and 1991–2000, consistent with the framework above. Contrary to the SWWW dummy, the trade barrier measure is not binary, and a threshold needs to be introduced to distinguish “open” (that is, treated) from “closed” (that is, control) countries. In the baseline specification, we use the median tariff rate as the threshold; that is, countries with a lower average protection over the decade are regarded as open; countries with a higher import duties ratio as closed. We also discuss results that emerge if we set the threshold at the first and third quartile of the tariff level distribution. After merging the PT and DeJong-Ripoll data sets, we are left with 68 countries for the 1990s.
and 71 for the 1980s, significantly less than for the SWWW measure (which contains 113 and 109 observations in the 1990s and 1980s, respectively).

Although not key to our argument, we briefly discuss the estimation results based on this alternative indicator of trade liberalization. Table 7 shows the cross-country evidence for the DeJong-Ripoll measure, corresponding to Table 2 for the SWWW indicator. For tariff 1, the baseline (median) tariff threshold, the results, although statistically insignificant, would indicate that in a treated country—that is, a country with a tariff-to-imports ratio below the median—the annual growth rate of real per capita GDP in the 1990s is up to 1.2 percent lower than in the control country. Compared with the results in Table 2, there is a somewhat larger discrepancy between OLS and matching estimates. In the 1980s, the link between openness and trade flips to become mostly positive, but again largely insignificant, except for the ATT estimator without area dummies.

<table>
<thead>
<tr>
<th>Table 6. Openness and Growth, Refined Evidence (II), 1961–2000</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>ATE:</strong> $E[Y(1) - Y(0)</td>
</tr>
</tbody>
</table>

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<tr>
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</thead>
<tbody>
<tr>
<td><strong>Matching with common support A</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Estimate (ATE)</td>
<td>1.318*</td>
<td>0.884</td>
<td>1.916**</td>
<td>−0.273</td>
</tr>
<tr>
<td>SE</td>
<td>(0.672)</td>
<td>(0.510)</td>
<td>(0.743)</td>
<td>(0.705)</td>
</tr>
<tr>
<td>Estimate (ATT)</td>
<td>1.130</td>
<td>0.813</td>
<td>1.645*</td>
<td>0.513</td>
</tr>
<tr>
<td>SE</td>
<td>(0.742)</td>
<td>(0.627)</td>
<td>(0.686)</td>
<td>(1.049)</td>
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<tr>
<td><strong>Matching with common support B</strong></td>
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<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Estimate (ATE)</td>
<td>1.644*</td>
<td>1.139*</td>
<td>2.007**</td>
<td>−0.360</td>
</tr>
<tr>
<td>SE</td>
<td>(0.731)</td>
<td>(0.524)</td>
<td>(0.717)</td>
<td>(0.641)</td>
</tr>
<tr>
<td>Estimate (ATT)</td>
<td>1.407</td>
<td>2.019**</td>
<td>1.936**</td>
<td>0.034</td>
</tr>
<tr>
<td>SE</td>
<td>(0.819)</td>
<td>(0.747)</td>
<td>(0.690)</td>
<td>(0.833)</td>
</tr>
<tr>
<td><strong>Matching with common support C</strong></td>
<td></td>
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<td></td>
<td></td>
</tr>
<tr>
<td>Estimate (ATE)</td>
<td>1.644*</td>
<td>1.141*</td>
<td>1.774*</td>
<td>−0.203</td>
</tr>
<tr>
<td>SE</td>
<td>(0.731)</td>
<td>(0.531)</td>
<td>(0.757)</td>
<td>(0.678)</td>
</tr>
<tr>
<td>Estimate (ATT)</td>
<td>1.407</td>
<td>1.691*</td>
<td>1.619*</td>
<td>0.304</td>
</tr>
<tr>
<td>SE</td>
<td>(0.819)</td>
<td>(0.725)</td>
<td>(0.651)</td>
<td>(0.843)</td>
</tr>
</tbody>
</table>

Source: Authors’ calculations based on data in Persson and Tabellini (2006).
Note: Dependent variable ($Y$): real GDP per capita growth. Treatment indicator ($T$): trade openness dummy (Sachs and Warner, 1995; Wacziarg and Welch, 2003). Control variables ($X$) include initial GDP per capita, secondary school enrollment, population growth, investment share, and area dummies. Restricted samples to meet the common-support condition for investment share (A), secondary school enrollment (B), or both (C). See Figures 1–8 for the numbers of treated and control countries dropped because outside of common supports A, B, and C. ATE and ATT stand for average treatment effect and average treatment effect on the treated, respectively, and are estimated by nonparametric nearest-neighbor matching. *corresponds to 5 percent significance level; **corresponds to 1 percent significance level.
The other threshold levels also offer conflicting and mainly weak signals: for tariff 2—only countries with a tariff-to-imports ratio below the 25th percentile of the distribution are defined as open—the results broadly indicate a positive effect of a lower tariff level on growth in the 1990s and are mixed for the 1980s. For tariff 3—only countries with a tariff-to-imports ratio below the 75th percentile of the distribution are defined as open—the
results indicate a significant negative effect of trade liberalization on growth in the 1990s and mixed effects in the 1980s.

More importantly for our purposes, the country pairings stemming from the matching exercise are similarly far-fetched in the case of the DeJong-Ripoll tariff measures as in the case of the SWWW trade liberalization.

<table>
<thead>
<tr>
<th>Country Pairings</th>
<th>DeJong-Ripoll Tariffs</th>
<th>SWWW Trade Liberalization</th>
</tr>
</thead>
<tbody>
<tr>
<td>Africa</td>
<td>No (1–11)</td>
<td>No (0–12)</td>
</tr>
<tr>
<td>Asia</td>
<td>Yes (4–7)</td>
<td>Yes (4–8)</td>
</tr>
<tr>
<td>Latin America</td>
<td>Yes (5–9)</td>
<td>Yes (4–13)</td>
</tr>
<tr>
<td>Middle East</td>
<td>No (0–6)</td>
<td>Yes (1–5)</td>
</tr>
<tr>
<td>Developed</td>
<td>No (23–0)</td>
<td>No (23–0)</td>
</tr>
<tr>
<td>Transition</td>
<td>Yes (1–1)</td>
<td>No (0–2)</td>
</tr>
<tr>
<td>Economies</td>
<td></td>
<td>No (2–0)</td>
</tr>
</tbody>
</table>


Note: Dependent variable (Y): real GDP per capita growth. Treatment indicators (T): tariff1 equal to 1 if import duties as a percentage of imports lower than the sample median; tariff2 equal to 1 if import duties as a percentage of imports lower than the sample 25th percentile; tariff3 equal to 1 if import duties as a percentage of imports lower than the sample 75th percentile. Control variables (X) include: initial GDP per capita, secondary school enrollment, population growth, investment share, and area dummies (as indicated). The two numbers in parentheses after each area refer to the number of treated and control countries, respectively. Samples restricted to certain areas to meet the common-support condition for area dummies. ATE and ATT stand for average treatment effect and average treatment effect on the treated, respectively, and are estimated by nonparametric nearest-neighbor matching. *corresponds to 5 percent significance level; **corresponds to 1 percent significance level.
In Table 8, we proceed to eliminate regions that lack common support with respect to regional affiliation; that is, where there are not enough treated compared with control countries (or vice versa) according to the criterion established above. Due to the smaller sample size of the DeJong-Ripoll data set, the number of observations drops further, and amounts now to less than half of the observations for the SWWW dummy (see Table 5). For example, the group of developed countries is excluded from the estimation for the median and 75th percentile threshold specification, whereas it is included for the 25th percentile threshold of the tariff-to-GDP ratio, implying that developed economies on average have a rather low tariff level and can only be included in the restricted sample if countries with a tariff ratio above the 25th percentile are coded closed. All estimates are insignificant and there is no clear directional effect: in precisely 50 percent of the cases (11), the impact is positive, in the rest negative. Using tariff 2—that is, estimating a country sample based exclusively on the group of developed countries in the 1990s and almost exclusively in the 1980s—the impact of trade liberalization is unambiguously positive (but insignificant), qualitatively confirming a result in DeJong and Ripoll (2006), who find evidence of a negative relationship between tariffs and growth only among the world’s rich countries.

To sum up, for the DeJong-Ripoll measure of tariff barriers, and after controlling for common support, the sample size drops drastically and makes sound statistical inference difficult. Although the unrestricted sample shows a more pronounced negative effect of trade openness (in the sense of a low tariff level) on growth, the appropriately restricted sample does not convey any strong message and is even consistent with the opposite affirmation.

IV. Conclusions

In this paper, we take another look at the openness-growth nexus in international macroeconomics. To add empirical glasnost to the results obtained in the literature, we examine classic pooled cross-country regressions and show the pitfalls related to the underlying country comparisons. Employing matching estimators from the treatment evaluation literature, we show that the country matches behind the estimates are often far-fetched—the unbearable lightness of cross-country estimates. We explain this problem as a lack of overlap between open and closed countries in the covariate space.

We show that restricting the sample to treated and control countries that share a common support is not always feasible due to data restrictions related to the openness measure or the data set employed. When this restriction can be applied, as is the case of the SWWW openness indicator, we confirm a positive correlation between trade openness and growth in selected regions after 1970. As the conditional independence assumption is likely not to hold

17See Tables 9 through 12 in the online appendix.
also in samples restricted to meet the common-support condition, however, we cannot interpret this correlation as a causal effect.

The main argument of the paper, however, goes far beyond the evidence presented. The lack of glasnost identified above is not limited to OLS-type estimates but extends to any econometric framework that uses at least some cross-sectional variation. Controlling for common support should, hence, be part of any empirical strategy in macro cross-country investigations.

REFERENCES


Public Debt, Money Supply, and Inflation:
A Cross-Country Study

GOOHOON KWON, LAVERN MCFARLANE, and WAYNE ROBINSON*

This paper provides comprehensive empirical evidence that supports the predictions of Sargent and Wallace’s “unpleasant monetarist arithmetic” that an increase in public debt is typically inflationary in countries with large public debt. Drawing on an extensive panel data set, we find that the relationship holds strongly in indebted developing countries, weakly in other developing countries, and generally does not hold in developed economies. These results are robust to the inclusion of other variables, corrections for endogeneity biases, relaxation of common-slope restrictions, and are invariant over subsample periods. We estimate a vector autoregression to trace out the transmission channel and find the impulse responses consistent with the predictions of a forward-looking model of inflation. Wealth effects of public debt could also affect inflation, as posited by the fiscal theory of the price level, but we do not find supportive evidence. The results suggest that the risk of a debt-inflation trap is significant in highly indebted countries and pure money-based stabilization is unlikely to be effective over the medium term. Our findings stress the importance of institutional and structural factors in the link between fiscal policy and inflation. [JEL E31, E62, E63, C59]


*Goohoon Kwon is Executive Director of the Global Investment Research Department of Goldman Sachs. Lavern McFarlane is an economist, and Wayne Robinson is the chief economist, with the Research Department of the Bank of Jamaica. The authors thank Ruben Atoyan, Mark de Broeck, Robert Flood, Jaewoo Lee, John Robinson, and Jeromin Zettelmeyer as well as seminar participants at the IMF and the Bank of Jamaica for valuable comments and suggestions. The authors are also grateful to two anonymous referees for their helpful comments.

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The most widely accepted school of thought on inflation is that it is a monetary phenomenon and therefore the reduction of inflation is largely the purview of monetary policy, particularly in the initial stages of disinflation. This school of thought, based on the quantity theory of money, posits that inflation is determined solely by the change in the relative supply of money and goods. Against this background, disinflation policy in many countries is framed with the objective of constraining monetary growth to be in line with the expansion in nominal income. Hence, an increasing number of countries have granted their central banks autonomy in the hope that it will insulate them from having to accommodate imprudent fiscal policies.

However, given that current money demand should depend on expectations about future inflation, a purely monetary effort at reducing inflation may not be successful. Theoretically, once account is taken of forward-looking expectations, multiple equilibrium paths for inflation can coexist. Under such circumstances, money supply alone may not be sufficient to pin down the time path of inflation.

Against this background, attention has increasingly been given to the role of fiscal policy in determining inflation. The main result of the seminal paper by Sargent and Wallace (1981) is that the effectiveness of monetary policy in controlling inflation depends critically on its coordination with fiscal policy. In their model, tighter monetary policy could lead to higher inflation under certain circumstances, even when the traditional relationship between money and the price level holds. The rationale is that, with the demand for government bonds given and in the absence of changes in future fiscal policy, a part of government obligations has to be covered by seignorage at some point in the future.

A similar line of reasoning lies behind the fiscal theory of the price level (FTPL). Apart from seignorage financing, traditional analysis of the fiscal impact on inflation focuses mostly on Keynesian aggregate demand considerations, public wage spillovers to private sector wages, and taxes affecting marginal costs and private consumption (Elmendorf and Mankiw, 1999). The FTPL identifies the wealth effect of government debt as an additional channel of fiscal influence on inflation and, amid debates over the coherence of the theory (Buiter, 1999; Niepelt, 2004), has spawned an extensive literature (Sims, 1994; Woodford, 1994 and 2001; Loyo, 1999; Christiano and Fitzgerald, 2000; Canzoneri, Cumby, and Diba, 2001; Cochrane, 2001 and 2005; Gordon and Leeper, 2002). This theory posits that increased government debt adds to household wealth and, hence, to demand for goods and services, leading to price pressures.

This paper provides a comprehensive empirical examination of the link between fiscal policy and inflation identified by the forward-looking fiscal-monetary models of inflation. We draw on an extensive cross-country data set for 71 countries spanning up to 43 years. We think that this helps overcome, or mitigate substantially, potential biases arising from the
selection of sample countries and sample periods. In addition, given the importance of policy regimes in the forward-looking models, we rely on flexible econometric techniques allowing for cross-country heterogeneity, which is often neglected in empirical studies for the sake of stronger testing power. Our approach also differs from much of the existing empirical literature on fiscal policy and inflation (Evans, 1987a and 1987b; Elmendorf, 1993; Ardagna, Caselli, and Lane, 2004; Catao and Terrones, 2005) in that we focus on the role of public debt—instead of the budget deficit—in determining inflation and inflation expectations. In so doing, we account for the nontraditional channels of fiscal influence on inflation—namely, monetization expectations and wealth effects of public debt, which do not necessarily move in lockstep with the size of the budget deficit due to a host of technical factors including nondebt financing, debt indexation, exchange rate movements, as well as the government’s assumption of accumulated quasi-fiscal liabilities (IMF, 2003; Singh and others, 2005).

The closest paper to our study is that of Castro, De Resende, and Ruge-Murcia (2003), who by estimating the degree of interdependence between monetary and fiscal policy for developed countries draw inferences about the relation between debt and the price level. Our empirical results are consistent with their conclusion that debt plays only a minor role in determining the price level in developed countries. Our contribution is to estimate the link between public debt and inflation in a generalized framework and to extend the empirical analysis to developing countries.

I. Conceptual Framework

There is a rich literature on forward-looking models of inflation. Aiyagari and Gertler (1985) introduced an overlapping generation model, deriving a simple, direct link between public debt and the price level. Calvo (1988) developed an alternative model based on a loss function of the authorities, which establishes a similar link between prices and public debt. Bohn (1988) also created a rational expectation model of a similar nature. The key common ingredients of these models are rational expectations, Cagan-type money demand, and a non-Ricardian regime that takes government bonds as net wealth.

In one variant of such models, a functional relationship can be derived for price, on the one hand, and money, debt, and output, on the other, as follows (see Appendix I for the derivation):

\[
P_t = \frac{(M_t + \delta B_t)}{\gamma(i)w}, \quad \gamma(i) = \beta \left( \frac{1 + i_t}{i_t} \right) + \alpha \delta,
\]

where \( P_t \), \( M_t \), \( B_t \), and \( w \) denote price, money, debt, and real income (or wealth in this simple model). \( \alpha \) and \( \beta \) are functions of the structural parameters of the household’s optimization problem, \( i_t \) is the yield on debt, and \( \delta \) is a portion of government debt that is not backed by the government’s current and future primary surpluses.
Equation (1) nests the quantity theory of money and the unpleasant monetary arithmetic of Sargent and Wallace (1981). The price level is proportional to the broadly defined monetary aggregate, \( M_t + \delta B_t \), which is the sum of high-powered money demanded by the household for transactions and by the government for monetizing the debt, with \( \delta \) reflecting the extent of monetary accommodation of the budget deficit and, more broadly, the nature of coordination between monetary and fiscal policy. Suppose the government pursues a policy of no monetization of its debt and runs a balanced budget over the long term. Then, \( \delta \), which can be taken as a monetization factor, reduces to 0 and the equation simplifies to the conventional quantity theory of money. More broadly, if fiscal policy is undertaken flexibly, for example, in ways to keep the debt-to-GDP ratio fixed all the time, then the monetization factor will remain 0 with no effect of public debt on prices. Alternatively, if the policy arrangement is full monetization of all public debt, \( \delta \) becomes 1, implying that an increase in public debt would influence inflation as strongly as monetary expansion does. In reality, the parameter is likely to vary between 0 and 1, with the exact scale depending on the capacity and willingness of the government to service public debt, as often construed from the debt size, policy credibility, and institutional and political constraints.

Equation (1) is also consistent with the predictions of the FTPL. Although we do not explicitly allow for the wealth effect of government debt as advanced by some versions of the FTPL (Leeper and Yun, 2006), the equation is fully consistent with the general implications of the theory. Therefore, the establishment of a positive relationship between public debt and prices does not necessarily tell whether the link is from monetization concerns as stressed by Sargent and Wallace (1981) or the wealth effects as advanced by the FTPL. We will discuss theoretical and empirical implications of the two competing models of inflation and attempt to test them by distinguishing the residency of public debt holders in the next section.

Drawing on Equation (1), we consider the following generalized price function.

\[
Pt = f(X_t) = f\left(M_t, B_t, w_t\right), \quad \text{where } f_1 > 0, f_2 > 0, \text{ and } f_3 < 0. \tag{2}
\]

Equation (2) can be log-linearized around equilibrium values \( X^* \) to obtain a more tractable specification as follows:

\[
\log Pt = f(X^*_t) + X^*_t f'(X^*_t) \hat{x}_t, \quad \text{where } \hat{x}_t = \log X_t - \log X^*_t, \\
\text{then, } \hat{p}_t = f(X^*_t) - \log P^*_t + X^*_t f'(X^*_t) \hat{x}_t, \quad \text{where } \hat{p}_t = \log P_t - \log P^*_t. \tag{3}
\]

This transformation establishes a simple linear relationship between inflation and increases in money supply, public debt, and output. Equation (3) could be modified to the following dynamic form, which embodies a partial adjustment process that allows gradual restoration to the equilibrium

\[
\begin{align*}
\Delta P_t & = f\left(M_t, B_t, P_t\right) - f\left(M_t, B_t, P^*_t\right) \\
\Delta B_t & = f\left(M_t, B_t, P_t\right) - f\left(M_t, B_t, P^*_t\right)
\end{align*}
\]
(Hendry, Pagan, and Sargan, 1984):

\[
p_t = \alpha \hat{p}_{t-1} + \beta_1 \hat{m}_t + \beta_2 \hat{b}_t - \beta_3 \hat{w}_t,
\]

where \( \hat{p} \), \( \hat{m} \), \( \hat{b} \), and \( \hat{w} \) denote deviations from equilibrium values in logarithms of price, money, debt, and real income, respectively.

II. Empirical Findings of the Cross-Country Study

Basic Stylized Facts

Our main data set is a panel dataset spanning 71 countries over up to 42 years (1963–2004). Data definitions and sources, as well as country grouping criteria are in Appendix II. Table 1 provides selected descriptive statistics of the main data set. It shows that, during the sample period, the average growth rate of money exceeded average inflation by about 4 percentage points per annum. Money supply grew at about the same pace as nominal GDP, implying a virtually unchanged level of money velocity over the long term. In contrast, public debt grew faster than both nominal GDP and money by about ½ percentage point per annum—a small but significant difference if extended over the long term. This could reflect financial deepening, which tends to expand nonmonetary financial instruments faster than monetary aggregates.

There is considerable variation across countries in the data, indicating potentially large gains from using panel data. Table 2 shows a summary of regional variations of selected macroeconomic indicators averaged over the sample period. Average inflation (geometric) is the highest in European developing countries, reflecting hyperinflation in many of these countries that transitioned to market economies in the early 1990s. With regard to public debt and inflation, public debt rose nearly twice as fast as inflation in low-inflation regions but not quite as fast in high-inflation regions. This could imply that nominal debt issuance, if excessive, is eroded quickly by inflation, pointing to the existence of a natural limit to real debt growth. A similar observation could be made with respect to the relationship between money growth and inflation but the extent is less prominent.

Our preferred form of data for regressions is first differences. Panel cointegration tests are not conclusive, as is often the case with medium-sized panels. The tests for stationarity, based on Pedroni (1999), reject the null of cointegration among the four main variables (CPI, money, public debt, and real output) in both the pooled and group mean \( t \)-tests at the 5 percent level but not always in the panel and group \( r \)-tests (Table 3).\(^1\) In light of these mixed outcomes, we proceed mainly with their first-difference terms, which are stationary, as we are keen to avoid the risk of spurious regressions arising from partially nonstationary or highly persistent data. Figure 1 shows the means of cross-country data in the first-difference logarithmic terms over the

\(^1\)We thank Peter Pedroni for sharing his computer programs.
full sample period. It also shows the means of public debt ratios, which increased sharply over about decade and half following the end of the Breton Woods era. Similar patterns are also observed in their median values.

Limitations of Long-Term Average Data

We first undertake a simple long-term cross-country regression as a quick way of examining relations among the variables. The results, presented in Table 4, confirm the findings of other empirical studies that long-term inflation is strongly positively associated with long-term money growth and negatively with long-term output growth. This is in line with the quantity theory of money and consistent with many empirical studies on this subject (Schwartz, 1973; Vogel, 1974; Lucas, 1980; Duck, 1993; Favero and Spinelli, 1999). In addition, the regression results show that flexible exchange regimes tend to be associated with higher inflation, although the causality is by no means established in this simple regression. With regard to the role of public debt, there is a positive linear relationship between inflation and public debt growth and a weak association between inflation and the size of public debt (Figure 2). However, both the levels and changes in public debt lose their explanatory power for inflation when money growth is controlled for.

It is, however, difficult to make direct inferences about the link between public debt and inflation from these long-term data. Although these results appear to reconfirm the dominant influence of money supply on long-term inflation, they do not necessarily reject the possibility that large and increasing public debt could push up inflation at some points in time. This can occur because the accumulation of public debt will either eventually increase primary surpluses or lead to monetization and inflation, depending on the policy regimes in place (Sargent, 1982). In the former case, there will
Table 2. Selected Macroeconomic Indicators, 1963–2004
(Average annual percentage changes, unless otherwise indicated)

<table>
<thead>
<tr>
<th></th>
<th>Real GDP Growth</th>
<th>Inflation</th>
<th>Money Growth¹</th>
<th>Public Debt Growth</th>
<th>Debt-GDP (ratio)</th>
<th>M-GDP (ratio)</th>
<th>Nominal GDP Growth</th>
<th>Seignorage (% GDP)</th>
<th>Fx Deprec</th>
<th>Fx Regime²</th>
<th>Years Covered³</th>
<th>Start Year</th>
<th>End Year</th>
</tr>
</thead>
<tbody>
<tr>
<td>Unweighted averages</td>
<td>3.6</td>
<td>14.2</td>
<td>18.6</td>
<td>21.9</td>
<td>51.8</td>
<td>18.6</td>
<td>18.3</td>
<td>2.8</td>
<td>9.9</td>
<td>2.3</td>
<td>20</td>
<td>1973</td>
<td>2002</td>
</tr>
<tr>
<td>Major advanced economies</td>
<td>2.9</td>
<td>5.7</td>
<td>9.1</td>
<td>12.4</td>
<td>54.2</td>
<td>39.7</td>
<td>8.9</td>
<td>4.2</td>
<td>0.1</td>
<td>0.0</td>
<td>32</td>
<td>1968</td>
<td>1999</td>
</tr>
<tr>
<td>Other advanced economies</td>
<td>3.6</td>
<td>11.2</td>
<td>13.9</td>
<td>17.2</td>
<td>46.5</td>
<td>11.4</td>
<td>15.7</td>
<td>3.4</td>
<td>5.6</td>
<td>1.9</td>
<td>36</td>
<td>1967</td>
<td>2003</td>
</tr>
<tr>
<td>Developing countries (48)</td>
<td>3.7</td>
<td>17.4</td>
<td>22.4</td>
<td>25.8</td>
<td>52.2</td>
<td>14.2</td>
<td>21.6</td>
<td>2.3</td>
<td>13.9</td>
<td>2.3</td>
<td>28</td>
<td>1975</td>
<td>2003</td>
</tr>
<tr>
<td>Latin America and Caribbean (20)</td>
<td>3.0</td>
<td>21.7</td>
<td>26.6</td>
<td>30.9</td>
<td>51.7</td>
<td>11.9</td>
<td>25.2</td>
<td>2.1</td>
<td>18.5</td>
<td>1.9</td>
<td>28</td>
<td>1975</td>
<td>2003</td>
</tr>
<tr>
<td>Latin America (13)</td>
<td>3.3</td>
<td>29.1</td>
<td>34.9</td>
<td>39.4</td>
<td>36.5</td>
<td>11.2</td>
<td>33.1</td>
<td>3.0</td>
<td>26.0</td>
<td>2.4</td>
<td>28</td>
<td>1976</td>
<td>2003</td>
</tr>
<tr>
<td>Caribbean (7)</td>
<td>2.4</td>
<td>8.0</td>
<td>11.3</td>
<td>15.0</td>
<td>79.9</td>
<td>13.0</td>
<td>10.6</td>
<td>0.5</td>
<td>4.6</td>
<td>1.4</td>
<td>28</td>
<td>1975</td>
<td>2003</td>
</tr>
<tr>
<td>Asia (9)</td>
<td>4.9</td>
<td>8.2</td>
<td>13.9</td>
<td>15.8</td>
<td>48.9</td>
<td>12.1</td>
<td>14.0</td>
<td>1.8</td>
<td>6.0</td>
<td>2.0</td>
<td>31</td>
<td>1971</td>
<td>2002</td>
</tr>
<tr>
<td>Middle East (6)</td>
<td>5.1</td>
<td>7.1</td>
<td>16.0</td>
<td>19.2</td>
<td>63.7</td>
<td>22.0</td>
<td>13.9</td>
<td>2.5</td>
<td>3.0</td>
<td>1.9</td>
<td>31</td>
<td>1971</td>
<td>2002</td>
</tr>
<tr>
<td>Europe (5)</td>
<td>3.7</td>
<td>32.6</td>
<td>37.4</td>
<td>37.6</td>
<td>50.3</td>
<td>24.5</td>
<td>36.2</td>
<td>4.4</td>
<td>25.2</td>
<td>3.7</td>
<td>17</td>
<td>1986</td>
<td>2002</td>
</tr>
<tr>
<td>Africa (8)</td>
<td>3.4</td>
<td>13.0</td>
<td>15.3</td>
<td>20.2</td>
<td>49.9</td>
<td>11.2</td>
<td>16.6</td>
<td>1.9</td>
<td>9.9</td>
<td>2.5</td>
<td>29</td>
<td>1974</td>
<td>2003</td>
</tr>
</tbody>
</table>

Sources: IMF, International Financial Statistics (IFS); IMF, World Economic Outlook (WEO); Organization for Economic Cooperation and Development; IMF Western Hemisphere Department databases; and Reinhart and Rogoff (2004).

Note: Country groupings are based on the IMF’s WEO classifications as of September 2005. Details are in Appendix I.

¹Narrowest definitions of money available from IFS, WEO, and OECD databases.
²Based on de facto exchange regimes (scaled from 1 to 5) of Reinhart and Rogoff (2004). The higher the indices, the more flexible the exchange rate regime.
³Average number of years. For each country, the coverage is adjusted for the longest time period for which data are available.
be no relationship between debt growth and inflation and, in the latter, the correlation between money growth and inflation will likely be the dominant factor.

Main Results of Panel Data Regressions

Given the limitations of long-term average data, our main modeling strategy is to use panel data, which allows for variability of individual countries while preserving the dynamics of adjustment within countries. We consider the following dynamic model, drawing on Equation (4):

\[
\begin{align*}
\Delta \log cpi_{it} &= \alpha \Delta \log cpi_{it-1} + \beta_1 \Delta \log money_{it} + \beta_2 \Delta \log pdebt_{it} \\
&+ \beta_3 \Delta \log rgdp_{it} + \eta_i + t + \nu_{it},
\end{align*}
\]

for \( i = 1, \ldots, N, \) and \( t = 2, \ldots, T, \) where \( \eta_i \) and \( \nu_{it} \) have the standard error component structure

\[
E[\eta_i] = E[\nu_{it}] = E[\eta_i \nu_{it}] = 0,
\]

Table 3. Panel Cointegration Tests: CPI, Money, Debt, GDP

<table>
<thead>
<tr>
<th>Weighted by long-term variances</th>
<th>Panel v</th>
<th>Panel rho</th>
<th>Panel t</th>
<th>Panel adf</th>
<th>Group rho</th>
<th>Group t</th>
<th>Group adf</th>
</tr>
</thead>
<tbody>
<tr>
<td>Cross-section common-time effects subtracted</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Homogenous time trends</td>
<td>7.34</td>
<td>-4.62</td>
<td>-8.08</td>
<td>-2.64</td>
<td>-4.14</td>
<td>-12.40</td>
<td>-3.66</td>
</tr>
<tr>
<td>Heterogeneous time trends</td>
<td>8.47</td>
<td>-1.09</td>
<td>-6.23</td>
<td>-0.41</td>
<td>1.18</td>
<td>-7.73</td>
<td>0.07</td>
</tr>
<tr>
<td>No cross-section common-time effects</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Homogenous time trends</td>
<td>4.96</td>
<td>-1.23</td>
<td>-4.32</td>
<td>-0.64</td>
<td>-0.92</td>
<td>-6.93</td>
<td>-1.65</td>
</tr>
<tr>
<td>Heterogeneous time trends</td>
<td>3.13</td>
<td>0.22</td>
<td>-4.32</td>
<td>-0.19</td>
<td>1.23</td>
<td>-6.48</td>
<td>-1.19</td>
</tr>
<tr>
<td>Nonweighted</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Cross-section common-time effects subtracted</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Homogenous time trends</td>
<td>6.60</td>
<td>-3.76</td>
<td>-6.32</td>
<td>-2.73</td>
<td>-4.14</td>
<td>-12.40</td>
<td>-3.66</td>
</tr>
<tr>
<td>Heterogeneous time trends</td>
<td>7.72</td>
<td>0.60</td>
<td>-3.05</td>
<td>-1.41</td>
<td>1.18</td>
<td>-7.73</td>
<td>0.07</td>
</tr>
<tr>
<td>No cross-section common-time effects</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Homogenous time trends</td>
<td>2.59</td>
<td>0.45</td>
<td>-2.37</td>
<td>0.86</td>
<td>-0.92</td>
<td>-6.93</td>
<td>-1.65</td>
</tr>
<tr>
<td>Heterogeneous time trends</td>
<td>1.24</td>
<td>0.86</td>
<td>-3.00</td>
<td>0.20</td>
<td>1.23</td>
<td>-6.48</td>
<td>-1.19</td>
</tr>
</tbody>
</table>

Source: Authors’ calculation.

Note: Based on unbalanced panel cointegration tests of Pedroni (1999) for price, money, public debt, and output. Statistics in bold note the rejection of the null of no-cointegration at the 5 percent confidence level.
and the transient errors are serially uncorrelated:

$$E[V_t V_s ] = 0 \text{ for } s \neq t \text{ for } i = 1, \ldots, N, \text{ and } t = 2, \ldots, T.$$ (7)

$d \log cpi$ refers to inflation, and $d \log money$, $d \log pdebt$, and $d \log rgdp$ refer to changes in money, public debt, and real GDP, respectively, all in first-difference logarithms. $t_r$ is a set of time dummies to control for unobserved global inflation pressures shocks such as the oil crises in the 1970s. $\eta_i$ represents unobserved country-specific effects, which are meant to capture heterogeneity in the debt-inflation nexus across countries as implied by our conceptual framework. This model is designed to reflect potentially complex
dynamics of public debt, inflation, and other macroeconomic variables within the constraints of a medium-sized panel. The existence of the fixed country effects, as opposed to the random effects, is supported by the results of the Breusch-Pagan (1980) Lagrange multiplier test. The poolability of the panel data is easily rejected by the standard Chow test.

The conceptual framework as reflected in Equations (1) and (2) suggests that the coefficients for debt and money should be positive and negative for output. In most specifications, we assume that coefficients in the vector $\beta$ are constant for each country group although we relax this homogenous slope assumption in robustness tests. No other restrictions are imposed on the coefficients of the explanatory variables as the conceptual framework is ambivalent about the speed of adjustment to the equilibrium.

In addition, we assume for now that all explanatory variables in $X_{it-s}$ (d log $cpi_{it-s-1}$, d log $money_{it-s}$, d log $pdebt_{it-s}$, d log $rgdp_{it-s}$) are predetermined such that

$$E[X_{it-s}v_{it}] = 0 \text{ for } s \geq 0.$$  \hspace{1cm} (8)

This standard set of assumptions yields the following moment conditions

$$E[X_{it-s}\Delta v_{it}] = 0 \text{ for } s \geq 1.$$  \hspace{1cm} (9)

These two moment conditions allow the use of suitably lagged levels of the variables as instruments, after the equation has been first differenced to eliminate the country-specific effects (Arellano and Bond, 1991). Given that our regression variables are not persistent as shown in our panel unit root test (Table 3), we believe that the first-differenced instruments are adequate and do not suffer from a weak instrument problem.\(^2\)

\(^2\)If autocorrelation is very high in endogenous right-hand variables, their lagged variables suffer from weak instrument problems. Blundell and Bond (2000), for example, consider problems of highly persistent variables in the estimation of production functions and, more broadly, Stock, Wright, and Yogo (2002) discuss a weak instruments problem in general terms.
The moment condition (8), based on the assumption of predetermined explanatory variables, might not necessarily hold in many countries. Macroeconomic policies, as embodied in money and debt variables in our

![Figure 2. Scatter Plots of Selected Macroeconomic Indicators and Public Debt Growth](image)


Note: See Appendix II for country lists, data sources, and definition as well as sample periods.
model, could be undertaken without any significant delays from the occurrence of shocks. Monetary policy, in particular, could be tightened at early signs of inflation but fiscal policy could be relaxed quickly in response to natural disasters. Separately, some variables, in particular interest rates and exchange rates, could change in sync with price shocks, affecting our explanatory variables—in particular the nominal value of foreign currency denominated debt—thereby invalidating the assumption of predetermination of these explanatory variables.

We will hence allow for contemporaneous feedback effects from money, debt, and output to prices, and use their lagged variables as explanatory variables as follows:

\[
\log cpi_{it} = \alpha d \log cpi_{it-1} + \beta_1 d \log money_{it-1} + \beta_2 d \log pdebt_{it-1} + \beta_3 d \log rgdp_{it-1} + \eta_i + t_i + \nu_{it}. \tag{10}
\]

We continue to assume that the transient errors are serially uncorrelated as before but that variables in \( X_{it} \) are endogenous with respect to the serially uncorrelated \( \nu_{it} \) shocks and hence only a subset of the moment conditions (8) and (9)

\[
E[X_{it-s}v_{it}] = 0 \quad \text{for } s \geq 1 \quad \text{and}
\]

\[
E[X_{it-s}\Delta v_{it}] = 0 \quad \text{for } s \geq 2,
\]

remains valid.

We are interested in consistent estimation of the parameters. A dynamic fixed-effect estimator and a first-difference GMM estimator are used for the purpose. For the GMM estimator, we prefer the one-step estimator to the two-step estimator, as the latter is prone to small sample biases, which could be considerable, in particular, in subperiod regressions. The possible existence of serial correlation of errors is handled through the use of a robust version of each estimator.

Regressions are run separately for different groups of countries in order to address a potential problem of slope heterogeneity without sacrificing efficiency gains from panel data. In line with the conceptual framework, countries are grouped according to the degree of economic development and, among subgroups, by the extent of sovereign indebtedness as classified by the most recent IMF’s World Economic Outlook (IMF, 2005): 13 major advanced economies, 10 other advanced economies, and 48 developing economies, among which 42 countries are classified as net debtors based on balance of payment data over 1972–2005. Although there could be different groupings,

---

3A dynamic pooled model is likely to bias the coefficient of a lagged dependent variable upward due to its correlation with time-invariant country effects (Bond, 2002). In contrast, estimates from a dynamic fixed-effect model are likely to be biased downward due to the demeaning process of the fixed-effect model (Kiviet, 1995). However, the biases are not a serious concern in our main specification as \( T \) is higher than 30 for the fixed-effect estimator (Judson and Owen, 1999) and \( N \) substantially higher than 20 for the GMM estimator.
we believe that our grouping serves the purpose of our study well because the criteria are objective and broadly correspond to institutional strength and policy credibility of sample countries. Appendix II provides a detailed country list and the grouping criteria.

Below is a summary of the main findings. Our regression results show a strong and stable positive effect of debt growth on inflation in developing economies but not in major advanced economies (Tables 5 and 6). In both specifications of Equations (5) and (10), the coefficient for public debt is 0.2 in the GMM estimator for developing countries but it is insignificant in developed countries. This implies that a 1 percent increase in public debt leads to a 0.2 percentage point increase in inflation. The coefficients are lower than those of money growth but are significant at the 5 percent level and rise to 0.3 for a subset of 42 indebted developing countries. There are signs of serial correlations in the regressions of specification (5) as evidenced by the presence of second-order serial correlation in the lagged first-differenced residual (Arellano and Bond, 1991). However, in the alternative specification (10) that uses lagged variables as explanatory variables, the serial correlations disappear and the coefficients for public debt remain largely unaffected.

Our findings are consistent with other empirical studies on inflation. Many studies report the existence of a positive relationship between budget deficits and inflation mostly in developing countries but not in developed economies (Feldstein, 1986; Orr, Edey, and Kennedy, 1995; Fischer, Sahay, and Vegh, 2002; Engen and Hubbard, 2004; Catao and Terrones, 2005). In the case of developed economies, many studies find virtually no linkage even between money and inflation (Dwyer, 1982; Christiano and Fitzgerald, 2003). Supportive empirical evidence is found also in studies of fiscal reaction functions, which conclude that fiscal policies in developed economies limit the increase in the debt-to-GDP ratio (Bohn, 1998) but those in developing countries do not (IMF, 2003).

We further explore the relationship between inflation and debt in order to examine possible determinants of the strength of such a relationship. We first test whether the debt-inflation linkage is affected by the level of debt. For this exercise, we divide countries into high- and low-debt economies, with the threshold determined by the median level of debt-to-GDP ratios in each country group. We then run the dynamic panel regressions over lagged variables (Equation (10)) for each group. We also run the same regressions with debt and money expressed in percent of GDP rather than in nominal terms, in order to address possible endogeneity of the nominal explanatory variables to inflation. The results show that highly indebted developing countries tend to have significant effects of debt growth on inflation but the relationship becomes insignificant or less robust if the sample is extended to all countries (Table 7). A similar outcome is observed for high inflation countries.

We also test whether exchange rate regimes matter in the link between debt growth and inflation. For the test, we define exchange rate regime dummies, drawing on Reinhart and Rogoff (2004). Our floating rate regime
<table>
<thead>
<tr>
<th></th>
<th>All Countries</th>
<th>Major Advanced</th>
<th>Other Countries</th>
<th>Developing Countries</th>
<th>Of which: Debtors&lt;sup&gt;1&lt;/sup&gt;</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Fixed</td>
<td>GMM&lt;sup&gt;2&lt;/sup&gt;</td>
<td>Fixed</td>
<td>GMM</td>
<td>Fixed</td>
</tr>
<tr>
<td>Lagged inflation</td>
<td>0.26</td>
<td>0.21</td>
<td>1.11</td>
<td>0.68</td>
<td>0.25</td>
</tr>
<tr>
<td></td>
<td>0.14</td>
<td>0.10</td>
<td>0.45</td>
<td>0.24</td>
<td>0.13</td>
</tr>
<tr>
<td>Money growth</td>
<td>0.37</td>
<td>0.27</td>
<td>0.22</td>
<td>0.33</td>
<td>0.37</td>
</tr>
<tr>
<td></td>
<td>0.06</td>
<td>0.12</td>
<td>0.21</td>
<td>0.36</td>
<td>0.07</td>
</tr>
<tr>
<td>Debt growth</td>
<td>0.09</td>
<td>0.13</td>
<td>0.01</td>
<td>0.02</td>
<td>0.12</td>
</tr>
<tr>
<td></td>
<td>0.04</td>
<td>0.07</td>
<td>0.01</td>
<td>0.03</td>
<td>0.07</td>
</tr>
<tr>
<td>Real GDP growth</td>
<td>−0.07</td>
<td>−0.07</td>
<td>0.50</td>
<td>−1.37</td>
<td>−0.08</td>
</tr>
<tr>
<td></td>
<td>0.06</td>
<td>0.06</td>
<td>0.51</td>
<td>1.01</td>
<td>0.06</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.38</td>
<td>0.26</td>
<td>0.59</td>
<td>0.59</td>
<td>0.59</td>
</tr>
<tr>
<td>Within</td>
<td>0.24</td>
<td>0.26</td>
<td>0.42</td>
<td>0.42</td>
<td>0.42</td>
</tr>
<tr>
<td>Between</td>
<td>0.92</td>
<td>0.34</td>
<td>0.94</td>
<td>0.94</td>
<td>0.94</td>
</tr>
<tr>
<td>Arellano-Bond AR (1)</td>
<td>−1.53</td>
<td>−1.43</td>
<td>−1.52</td>
<td>−1.46</td>
<td></td>
</tr>
<tr>
<td>Arellano-Bond AR (2)</td>
<td>1.96</td>
<td>−0.08</td>
<td>2.39</td>
<td>3.09</td>
<td></td>
</tr>
<tr>
<td>Number of countries</td>
<td>71</td>
<td>71</td>
<td>13</td>
<td>13</td>
<td>58</td>
</tr>
<tr>
<td>Number of observations</td>
<td>2,149</td>
<td>2,076</td>
<td>428</td>
<td>1,706</td>
<td>1,721</td>
</tr>
</tbody>
</table>

Source: Authors' calculation.

Note: Coefficients significant at the 5 percent level are in bold. The robust standard errors are below the estimated coefficients.

<sup>1</sup>Indebted developing countries, based on the balance of payments data over 1972–2005.

<sup>2</sup>GMM based on the first-difference transformation, assuming that explanatory variables are predetermined. Standard errors are adjusted for serial correlations and heteroscedasticity. The instrument set includes up to four lags of the right-hand side and explanatory variables.
### Table 6. Panel Regression Outcomes

(*Dependent variable: inflation 1963–2004*)

<table>
<thead>
<tr>
<th></th>
<th>All Countries</th>
<th>Major Advanced</th>
<th>Other Countries</th>
<th>Developing Countries</th>
<th>Of which: Debtors(^1)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Fixed</td>
<td>GMM(^2)</td>
<td>Fixed</td>
<td>GMM</td>
<td>Fixed</td>
</tr>
<tr>
<td>Lagged inflation</td>
<td>0.23</td>
<td><strong>0.10</strong></td>
<td>1.20</td>
<td>0.44</td>
<td>0.21</td>
</tr>
<tr>
<td></td>
<td>0.17</td>
<td>0.04</td>
<td>0.50</td>
<td>0.27</td>
<td>0.17</td>
</tr>
<tr>
<td>Lagged money growth</td>
<td><strong>0.30</strong></td>
<td>0.61</td>
<td>−0.03</td>
<td>0.65</td>
<td><strong>0.31</strong></td>
</tr>
<tr>
<td></td>
<td>0.10</td>
<td>0.10</td>
<td>0.18</td>
<td>0.56</td>
<td>0.09</td>
</tr>
<tr>
<td>Lagged debt growth</td>
<td>0.05</td>
<td><strong>0.18</strong></td>
<td>0.01</td>
<td>0.01</td>
<td>0.09</td>
</tr>
<tr>
<td></td>
<td>0.04</td>
<td>0.07</td>
<td>0.01</td>
<td>0.02</td>
<td>0.05</td>
</tr>
<tr>
<td>Lagged real GDP growth</td>
<td>−0.03</td>
<td>−0.33</td>
<td>0.84</td>
<td>1.14</td>
<td>−0.04</td>
</tr>
<tr>
<td></td>
<td>0.04</td>
<td>0.08</td>
<td>0.67</td>
<td>1.07</td>
<td>0.05</td>
</tr>
<tr>
<td>(R^2)</td>
<td>0.32</td>
<td>0.27</td>
<td>0.51</td>
<td>0.49</td>
<td>0.52</td>
</tr>
<tr>
<td>Within</td>
<td>0.18</td>
<td>0.40</td>
<td>0.32</td>
<td>0.31</td>
<td>0.31</td>
</tr>
<tr>
<td>Between</td>
<td>0.92</td>
<td>0.27</td>
<td>0.94</td>
<td>0.93</td>
<td>0.90</td>
</tr>
<tr>
<td>Arellano-Bond AR (1)</td>
<td>−1.87</td>
<td>−1.37</td>
<td>−1.77</td>
<td>−1.77</td>
<td>−1.51</td>
</tr>
<tr>
<td>Arellano-Bond AR (2)</td>
<td>1.62</td>
<td>−0.35</td>
<td>1.98</td>
<td>1.98</td>
<td>0.96</td>
</tr>
<tr>
<td>Number of countries</td>
<td>71</td>
<td>71</td>
<td>13</td>
<td>13</td>
<td>58</td>
</tr>
<tr>
<td>Number of observations</td>
<td>2,131</td>
<td>2,076</td>
<td>430</td>
<td>415</td>
<td>1,701</td>
</tr>
</tbody>
</table>

Source: Authors’ calculation.

Note: Coefficients significant at the 5 percent level are in bold. The robust standard errors are below the estimated coefficients.

\(^1\)Indebted developing countries, based on the balance of payments data over 1972–2005.

\(^2\)GMM based on the first difference transformation, assuming that explanatory variables are predetermined. Standard errors are adjusted for serial correlations and heteroscedasticity. The instrument set includes up to four lags of the right-hand side and explanatory variables.
dummy, which varies by time and across countries, represents an independently floating exchange rate regime as defined by Reinhart and Rogoff (2004). Our fixed rate regime dummy covers all other exchange rate regimes (those with pegs, limited flexibility, and managed floats). We then multiply the exchange rate regime dummies by the debt growth variable to isolate the impact of the exchange rate regimes on the debt-inflation nexus. The regressions show that the sensitivity of inflation to debt is higher and significant under a floating rate regime but the sensitivity is low and often insignificant under a fixed rate regime (Table 8). The results do not change substantially if we exclude managed floats from the fixed rate regime dummy. This finding could reflect, as stressed by Ghosh and others (1997), the generally favorable commitment effect of the fixed exchange rate regime on inflation and inflation expectations.

**Robustness of the Results**

We test the robustness of the results from three angles. First, we test whether the inclusion of other variables affect the regression outcomes. We are particularly concerned about the possibility that some explanatory variables, in particular changes in debt values, could be affected by current or future

<table>
<thead>
<tr>
<th></th>
<th>Debtors¹</th>
<th>All Countries</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>High debt</td>
<td>High inflation</td>
</tr>
<tr>
<td>Thresholds (mean)</td>
<td>54.0</td>
<td>14.4</td>
</tr>
<tr>
<td>Thresholds (median)</td>
<td>46.5</td>
<td>8.7</td>
</tr>
<tr>
<td>Debt-inflation coefficient²</td>
<td>Fixed-effect</td>
<td>0.06</td>
</tr>
<tr>
<td></td>
<td>First-difference GMM</td>
<td>0.18</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th></th>
<th>Debtors¹</th>
<th>All Countries</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>High debt</td>
<td>High inflation</td>
</tr>
<tr>
<td>Number of countries</td>
<td>21</td>
<td>21</td>
</tr>
<tr>
<td>Number of observations</td>
<td>589</td>
<td>559</td>
</tr>
</tbody>
</table>

Source: Authors’ calculation.
Note: Coefficients significant at the 5 percent level are in bold. Dynamic panel regressions on lagged variables (Equation (10)).
¹Indebted developing countries (see Appendix II for definition).
²Debt and money in nominal terms as in Equation (10).
³Debt and money expressed in percent of GDP.
Table 8. Panel Regression Outcomes
(Dependent variable: inflation 1963–2004)

<table>
<thead>
<tr>
<th></th>
<th>All Countries</th>
<th>Major Advanced</th>
<th>Other Countries</th>
<th>Debtors$^3$</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Fixed</td>
<td>GMM$^1$</td>
<td>GMM$^2$</td>
<td>Fixed</td>
</tr>
<tr>
<td>Lagged inflation</td>
<td>0.12</td>
<td>0.10</td>
<td>0.11</td>
<td>1.45</td>
</tr>
<tr>
<td></td>
<td>0.14</td>
<td>0.05</td>
<td>0.07</td>
<td>0.07</td>
</tr>
<tr>
<td>Lagged money growth</td>
<td>0.21</td>
<td>0.47</td>
<td>0.39</td>
<td>−0.13</td>
</tr>
<tr>
<td></td>
<td>0.06</td>
<td>0.11</td>
<td>0.09</td>
<td>0.19</td>
</tr>
<tr>
<td>Lagged debt growth</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Fixed rate regime</td>
<td>−0.00</td>
<td>−0.01</td>
<td>0.02</td>
<td>0.01</td>
</tr>
<tr>
<td></td>
<td>0.01</td>
<td>0.08</td>
<td>0.02</td>
<td>0.01</td>
</tr>
<tr>
<td>Floating rate regime</td>
<td>0.30</td>
<td>0.32</td>
<td>0.32</td>
<td>0.79</td>
</tr>
<tr>
<td></td>
<td>0.09</td>
<td>0.08</td>
<td>0.09</td>
<td>0.73</td>
</tr>
<tr>
<td>Lagged real GDP growth</td>
<td>−0.02</td>
<td>−0.43</td>
<td>−0.13</td>
<td>0.79</td>
</tr>
<tr>
<td></td>
<td>0.04</td>
<td>0.18</td>
<td>0.10</td>
<td>0.65</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.34</td>
<td></td>
<td></td>
<td>0.26</td>
</tr>
<tr>
<td>Within</td>
<td>0.20</td>
<td></td>
<td></td>
<td>0.27</td>
</tr>
<tr>
<td>Between</td>
<td>0.85</td>
<td></td>
<td></td>
<td>0.09</td>
</tr>
<tr>
<td>Arellano-Bond AR (2)</td>
<td></td>
<td>2.31</td>
<td>1.83</td>
<td>−0.92</td>
</tr>
</tbody>
</table>

Number of countries 71  71  71  13  13  58  58  42  42
Number of observations 2,131 2,076 2,076 430 415 1,701 1,661 1,181 1,157

Source: Authors’ calculation.
Note: Coefficients significant at the 5 percent level are in bold. The robust standard errors are below the estimated coefficients.
$^1$GMM based on the first difference transformation, assuming that explanatory variables are predetermined. Standard errors are adjusted for serial correlations and heteroscedasticity.
$^2$Three lags are used to remove autocorrelations of errors.
$^3$Indebted developing countries, based on the balance of payments data over 1972–2005.
inflationary shocks through a valuation channel. For example, foreign-currency-denominated debt, which is often a sizable portion of public debt, could rise or fall depending on the exchange rate, which is in turn affected by current inflation or the inflation outlook. We hence control for the valuation channel by including exchange rate movements as another explanatory variable and test whether its inclusion affects the regression coefficient for public debt. We also include the output gap, estimated from the Hodrick-Prescott (H-P) filter, to control for cyclical factors. The results are largely unchanged with the inclusion of these additional variables.

Second, we run rolling regressions for subsample periods in order to address a potential problem of parameter instability. The main results described above are largely maintained in regressions over each rolling 20-year period of 1963–83, 1972–93, and 1983–2003 (Table 9). The sensitivity of inflation to debt growth in indebted developing countries remains significant and similar to its sensitivity to money growth. It is notable that the coefficients are larger in the later period than in the earlier period, possibly reflecting the relative dominance of flexible exchange rate regimes during the post–Bretton Woods era.

Third, we relax the common slope assumption. Pesaran and Smith (1995) illustrates that, in the case of dynamic panel data with heterogeneous slopes, pooling and aggregating produce inconsistent and potentially highly misleading estimates of the coefficients. Hence we relax the common slope assumption and calculate the mean group estimator (Pesaran and Smith, 1995) and the panel fully modified OLS estimator (Pedroni, 2000). Mean group estimates show that debt growth, both contemporaneous and lagged ones, affect inflation positively and its degree is stronger in indebted developing countries (Table 10). Similar patterns are observed in fully modified OLS estimates (FMOLS), although the levels of the coefficients are not directly comparable to those from our main regressions in log difference form (Table 11). The panel FMOLS estimator is one of the least restrictive estimators for panel data, which is adjusted for endogeneity and short-run cross-country heterogeneity while exploiting long-run information contained in the panel.

Transmission Channels

We undertake a vector autoregression (VAR) to trace out the transmission channels of the fiscal influence on inflation and to cross-check the validity of the panel regressions above, particularly, with respect to a potential endogeneity problem. Our panel VAR consists of inflation and growth of public debt, money, and real GDP. Country fixed effects and time effects are controlled for through country dummies and time dummies. Impulse responses are based on the Cholesky decomposition of the structural

\[4\] In a similar vein, we also undertake an exclusion sensitivity analysis by removing each country sequentially from the regressions. The results remain largely unchanged.
Table 9. Panel Regression Outcomes  
(Dependent variable: inflation)

<table>
<thead>
<tr>
<th></th>
<th>All Countries</th>
<th>Major Advanced</th>
<th>Other Countries</th>
<th>Of which: Debtors¹</th>
</tr>
</thead>
<tbody>
<tr>
<td>Lagged inflation</td>
<td>0.59</td>
<td>0.48</td>
<td>0.18</td>
<td>0.58</td>
</tr>
<tr>
<td></td>
<td>0.09</td>
<td>0.15</td>
<td>0.17</td>
<td>0.07</td>
</tr>
<tr>
<td>Lagged money growth</td>
<td>0.12</td>
<td>0.16</td>
<td>0.34</td>
<td>0.13</td>
</tr>
<tr>
<td></td>
<td>0.03</td>
<td>0.04</td>
<td>0.11</td>
<td>0.04</td>
</tr>
<tr>
<td>Lagged debt growth</td>
<td>0.02</td>
<td>0.12</td>
<td>0.05</td>
<td>–0.02</td>
</tr>
<tr>
<td></td>
<td>0.01</td>
<td>0.04</td>
<td>0.04</td>
<td>0.02</td>
</tr>
<tr>
<td>Lagged real GDP growth</td>
<td>–0.04</td>
<td>0.01</td>
<td>–0.01</td>
<td>0.13</td>
</tr>
<tr>
<td></td>
<td>0.03</td>
<td>0.02</td>
<td>0.05</td>
<td>0.10</td>
</tr>
<tr>
<td>R²</td>
<td>0.78</td>
<td>0.82</td>
<td>0.30</td>
<td>0.80</td>
</tr>
<tr>
<td>Within</td>
<td>0.54</td>
<td>0.53</td>
<td>0.16</td>
<td>0.76</td>
</tr>
<tr>
<td>Between</td>
<td>0.96</td>
<td>0.92</td>
<td>0.87</td>
<td>0.94</td>
</tr>
<tr>
<td>Number of countries</td>
<td>60</td>
<td>67</td>
<td>71</td>
<td>13</td>
</tr>
<tr>
<td>Number of observations</td>
<td>819</td>
<td>1,187</td>
<td>1,312</td>
<td>202</td>
</tr>
</tbody>
</table>

Source: Authors’ calculation.
Note: Coefficients significant at the 5 percent level are in bold. Based on a dynamic fixed-effects model. Below the estimated coefficients are robust standard errors.

¹Indebted developing countries, based on the balance of payments data over 1972–2005.
Table 10. Mean Group Estimates
(Dependent variable: inflation 1963–2003)

<table>
<thead>
<tr>
<th>Countries Other Than:</th>
<th>Whole sample</th>
<th>Major advanced economies</th>
<th>Of which: debtor countries¹</th>
</tr>
</thead>
<tbody>
<tr>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
<td>(1)</td>
</tr>
<tr>
<td>Lagged inflation</td>
<td>0.50</td>
<td>0.55</td>
<td>0.50</td>
</tr>
<tr>
<td></td>
<td>0.04</td>
<td>0.03</td>
<td>0.05</td>
</tr>
<tr>
<td>(Lagged) money growth</td>
<td>0.07</td>
<td>0.08</td>
<td>0.08</td>
</tr>
<tr>
<td></td>
<td>0.02</td>
<td>0.02</td>
<td>0.02</td>
</tr>
<tr>
<td>(Lagged) debt growth</td>
<td>0.08</td>
<td>0.08</td>
<td>0.11</td>
</tr>
<tr>
<td></td>
<td>0.02</td>
<td>0.02</td>
<td>0.03</td>
</tr>
<tr>
<td>(Lagged) real GDP growth</td>
<td>-0.24</td>
<td>0.08</td>
<td>0.07</td>
</tr>
<tr>
<td></td>
<td>0.10</td>
<td>0.09</td>
<td>0.12</td>
</tr>
<tr>
<td>GDP gap</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
</tr>
<tr>
<td>Number of countries</td>
<td>71</td>
<td>71</td>
<td>71</td>
</tr>
</tbody>
</table>

Source: Authors' calculation.
Note: Coefficients significant at the 5 percent level are in bold. Based on country-by-country dynamic ordinary least squares (OLS) regressions. The standard errors are below the estimated coefficients.

¹Indebted developing countries, based on the balance of payments data over 1972–2005.
(1), (2): Mean of OLS regression coefficients for each country (over contemporaneous explanatory variables).
(3): Mean of OLS regression coefficients for each country (over one-year lag explanatory variables).
<table>
<thead>
<tr>
<th></th>
<th>Whole Sample</th>
<th>Advanced Economies</th>
<th>Developing Countries(^1)</th>
<th>Of which: Debtor Countries(^1)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Coefficient</td>
<td>Coefficient</td>
<td>Coefficient</td>
<td>Coefficient(^2)</td>
</tr>
<tr>
<td><strong>Money</strong></td>
<td>0.58</td>
<td>0.26</td>
<td>0.56</td>
<td>0.59</td>
</tr>
<tr>
<td></td>
<td>78.39</td>
<td>17.31</td>
<td>87.13</td>
<td>60.83</td>
</tr>
<tr>
<td><strong>Public debt</strong></td>
<td>0.13</td>
<td>0.21</td>
<td>0.05</td>
<td>0.25</td>
</tr>
<tr>
<td></td>
<td>11.95</td>
<td>11.86</td>
<td>4.14</td>
<td>25.32</td>
</tr>
<tr>
<td><strong>Real GDP</strong></td>
<td>-0.25</td>
<td>-0.09</td>
<td>-0.31</td>
<td>-0.32</td>
</tr>
<tr>
<td></td>
<td>-11.38</td>
<td>-1.32</td>
<td>-19.72</td>
<td>-1.51</td>
</tr>
<tr>
<td><strong>Number of countries</strong></td>
<td>71</td>
<td>23</td>
<td>48</td>
<td>25</td>
</tr>
</tbody>
</table>

Source: Authors’ calculation.
Note: Coefficients significant at the 5 percent level are in bold. Based on FMOLS regressions over the variables in the level (Pedroni, 2000).
\(^1\)Indebted developing countries, based on the balance of payments data over 1972–2005.
\(^2\)Common time dummies not included. Common time dummies included in all other regressions.
shocks in the order of output \((R)\), prices \((I)\), money \((M)\), and public debt \((D)\). This ordering allows possible contemporaneous feedback from inflation and money growth on debt growth but not from debt growth on inflation or money growth. This is a very conservative assumption, given that the focus of our exercise is to estimate the impact of debt growth on inflation. In the choice of the lag length, we use the Schwarz criterion which imposes a larger penalty for additional coefficients than the Akaike information criterion.

The results are largely comparable with those from our panel regressions, rendering additional support to the prediction of the fiscal-monetary model of inflation—the debt-inflation link exists and its extent is affected by institutional and structural factors. Our panel VARs show a weak or no response of inflation to fiscal shocks in major advanced economies (Figure 3a). A similar pattern is observed in the monetary response to fiscal shocks. In contrast, impulse responses for indebted developing countries (Figure 3b) show a strong and positive response of money supply and inflation to fiscal shocks. The results are largely invariant to changes in the shock ordering (Table 12).

Our findings dovetail with those of many empirical studies that document policy responses to macroeconomic shocks in developing countries (Mélitz, 1997; Akitoby and others, 2004; Kaminsky, Reinhart, and Vegh, 2004). The phenomenon of fiscal dominance in developing countries is confirmed by our VAR evidence that fiscal relaxation tends to be accommodated by monetary easing in those countries (the third chart in the right column of Figure 3b). Our VARs also show that fiscal and monetary policies in developing countries are largely insensitive to output shocks in contrast to some evidence of countercyclical fiscal policy in advanced economies (the left bottom of Figure 3a).

**Implications of the Unpleasant Monetary Arithmetic and the FTPL**

The implications of rising public debt for inflation are observationally similar in the Sargent-Wallace framework (1981) and the FTPL. Nonetheless, there is an important theoretical distinction between the two (Leeper and Yun, 2006). Under the FTPL, an increase in government debt raises the wealth of bond holders while not reducing the wealth of others. Hence, the increase in government debt boosts aggregate demand and pushes up the price level and in turn money demand. Money supply is endogenous in this regime and, as such, increases in response to the higher money demand. In this regime, the price level is the factor that equilibrates the nominal value of future discounted primary surplus and the nominal value of public debt. In contrast, under the Sargent-Wallace framework of the so-called unpleasant monetary arithmetic, an increase in government debt, if not fully backed by future real primary surplus, will increase concerns about monetization of public debt, which will in turn raise inflation expectations and thereby increase long-term interest rates. This will in turn reduce money demand and push up the price level even without a contemporaneous increase in money supply.
Figure 3a. Impulse Responses in Major Advanced Economies

Response to Cholesky One S.D. Innovations 2/S.E.

Response of DLOGRGDP to DLOGCPI

Response of DLOGRGDP to DLOGMONEY

Response of DLOGRGDP to DLOGPDEBT

Response of DLOGCPI to DLOGRGDP

Response of DLOGCPI to DLOGMONEY

Response of DLOGCPI to DLOGPDEBT

Response of DLOGMONEY to DLOGRGDP

Response of DLOGMONEY to DLOGCPI

Response of DLOGMONEY to DLOGPDEBT

Response of DLOGPDEBT to DLOGRGDP

Response of DLOGPDEBT to DLOGCPI

Response of DLOGPDEBT to DLOGMONEY

Source: Authors’ calculation.

Note: Country groupings are based on IMF, World Economic Outlook classifications as of September 2005. See Appendix II for details.
Source: Authors’ calculation.
Note: These are countries whose cumulative current account balance over the period 1972–2005 (1992–2005 for transition economies) is negative. See Appendix II for details.
Notwithstanding the observational similarity, the differences in the transmission channel allow an empirical testing of the two models. The FTPL implies that the wealth effect of debt increases should materialize mainly from public debt held by residents. In contrast, monetization concerns, as stressed by the Sargent-Wallace framework, should be affected by the total size of public debt, regardless of the residency of the debt holders. Our panel regressions of a smaller data set spanning a subset of 30 Latin and Caribbean countries between 1997 and 2004 point to the dominance of monetization concerns as opposed to the wealth effects, although the lack of debt data by residency makes it difficult to extend the analysis to other countries and over a longer time span. The regression results show that the impact of total public debt growth on inflation is significantly positive but the inflation impact of an increase in domestic public debt (that is, public debt held by residents) is insignificant (Table 13).

### III. Application to Jamaica

Jamaica is one of the most heavily indebted countries in the world. The public debt sharply increased to nearly 140 percent of GDP over the past decade from an already high level of 80 percent of GDP earlier. The sharp increase was due mainly to the assumption of off-budget liabilities, notably the bailout of financial institutions in the late 1990s, with accumulated budget deficits accounting for only a quarter of the surge. Debt service costs have hovered about 15 percent of GDP in recent years and to help meet these...
payments, primary surpluses have been generated in excess of 10 percent of GDP over the past several years.

Motivated by the need to reduce the large public debt, the Jamaican authorities started in 2004 an ambitious program that includes as its objective the goal of reducing inflation to single digits. The ultimate goal of the government’s comprehensive program is to reduce public debt to 100 percent of GDP by 2008 through fiscal consolidation. This consolidation effort, in turn, is expected to lead to a virtuous circle of higher economic growth, lower inflation, and lower interest rates and hence reduced debt.

Inflation in Jamaica has been high and volatile, compared with neighboring countries. Unlike many other countries in similar circumstances, the Bank of Jamaica (BOJ) has traditionally adopted a conservative monetary policy stance, with seignorage financing of the budget deficit rarely exceeding 1 percent of GDP. This policy stance was possible thanks to its strong operational autonomy, notwithstanding overall low statutory independence (Jácome and Vázquez, 2005). Inflation nonetheless has remained at double digits since 2003 and fluctuated widely but most neighboring countries had much lower inflation during the same period. The BOJ’s ability to reduce inflation was hampered by frequent exogenous shocks, large government debt and instruments for open market operations (OMOs), and already high sterilization costs (1½–2½ percent of GDP per year in recent years).

We apply a VAR to Jamaica to test whether the cross-country debt-inflation relationship identified in our panel regressions holds for Jamaica. The estimation uses annual data between 1980 and 2004 for CPI, real GDP,
reserve money, and government debt, which includes OMO debt. The exchange rates are also included in the robustness test to control for possible biases from exchange rate volatility on the debt dynamics. Data for GDP and CPI are from the Statistical Institute, and government debt from the Finance Ministry. All other data are from the BOJ. All the variables are nonstationary and, as such, we test whether any stationary long-run relation exists among the variables. Both the trace and maximum eigenvalue tests based on the full information maximum likelihood method reject the null hypothesis of no cointegration but the number of cointegration vectors depend on the specification of the cointegration equations, most probably in reflection of the short time span. Hence, we run VARs both with and without the error correction terms.

The VAR outcomes confirm the significance of public debt dynamics in determining inflation in Jamaica. The impulse response functions show that the price level is positively affected by money supply and public debt but the latter has stronger and more lasting effects on inflation (Figure 4). Also, fiscal shocks have positive and persistent effects on money supply but the opposite does not hold. These results are similar to those from the panel VAR estimates for developing countries and robust to changes in the ordering of the shocks. The directions of the impulse responses remain unchanged in an alternative VAR including the exchange rate as an endogenous variable and alternative regressions based on the vector error correction model.

Caution is needed, however, in interpreting these outcomes, as it is not clear which fiscal channel is driving inflation in Jamaica. The regression results do not separate the wealth effects of public debt from its effects on monetization expectations. It could well be that the former effects are more important in Jamaica than the latter effects, given the sustained efforts of the authorities for fiscal consolidation. It should, therefore, be stressed that our VAR results do not necessarily mean that the relatively high inflation in Jamaica signals concerns about monetization of debt. Notwithstanding this caveat, our regression results confirm that the movements of public debt do matter for inflation dynamics in Jamaica.

Figure 4. Jamaica: Impulse Responses

Source: Authors’ calculation.

1Based on a one year lag VAR using annual data from 1980 to 2004.
IV. Policy Implications

Our regression results point to a number of policy implications for countries with high debt. First, there is a significant risk of a debt-inflation trap in highly indebted countries. A rise in inflation expectations will eventually push up nominal interest rates, elevating public debt unless fully countered by a primary surplus. The debt increase will in turn raise inflation expectations further. This vicious feedback effect implies that rising inflation expectations could increase budgetary costs more than proportionally. This also means that rising inflation expectations could be destabilizing to the debt dynamics more than adverse output shocks do—possibly by as much as one-third to one-half (the numerical relationship is derived in Appendix III).

More broadly, the conduct of monetary policy is extremely challenging in highly indebted developing countries. In principle, flexibility in monetary policy would be severely constrained by considerations about implications of interest and exchange rate volatilities on debt dynamics. Operationally, monetary data alone might not provide reliable indications of emerging inflation pressures, as growth in government debt in lieu of money printing could also affect inflation expectations. In this regard, sustained sterilized intervention could backfire since such interventions would limit growth in money supply but raise public debt. In sum, in countries with large debt overhangs, purely money-based stabilization is unlikely to be effective without the support of fiscal consolidation.

Second, the importance of inflation expectations in the debt-inflation dynamics implies that the budgetary costs of noncredible disinflation policy are potentially large in highly indebted countries. In Jamaica, for example, the central bank has medium-term inflation forecasts of 5 percent, which are considerably lower than current inflation. Suppose that bond holders believe that inflation would indeed fall but remain still high at 10 percent over the medium term, with correspondingly higher nominal interest rates that they demand for holding debt. In the event that inflation indeed falls to 5 percent, the ex post budgetary real interest payments would be much higher (by about 3 percent of GDP, given Jamaica’s debt profile) than in the case of 10 percent inflation. Conversely, unanticipated inflation would help reduce borrowing costs in the short term but only exacerbate the credibility problem and ratchet up borrowing costs over the medium term. This points to the need for managing inflation and inflation expectations in ways to minimize surprises.

Third, institutional and structural factors matter in the dynamics between public debt and inflation. Fiscal rules that limit the size of budget deficits or public debt could, under appropriate circumstances, be an important institutional tool for safeguarding price stability to the extent that the commitment is credible. Independence of the central bank could also help

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*A similar observation has been made in Blanchard (2004) and Favero and Giavazzi (2004), which examined the relationship between depreciation expectations and public debt in Brazil.*
reduce monetization concerns. The development of the financial sector could help promote price stability as a developed financial market tends to support the central bank’s policy autonomy (Posen, 1995). The financial sector could also reinforce fiscal discipline by providing immediate and clear signals about perceived risks of debt monetization (Rubin and Weisberg, 2003).

V. Summary and Conclusions

Our study provides comprehensive and robust evidence in support of Sargent and Wallace’s (1981) “unpleasant monetary arithmetic” that an increase in government debt is typically inflationary, in countries with large public debt. The regression results show that an increase in public debt is significantly and strongly associated with high inflation in indebted developing countries, after controlling for money growth and real output growth. In contrast, this pattern holds less strongly in other developing countries and does not hold in major advanced economies, consistent with the thesis of a forward-looking model of inflation that—unlike the implications of a static aggregate demand model—policy regimes matter in the debt-inflation nexus. These results are invariant over subsample periods and robust to the inclusion of other variables, corrections for possible endogeneity biases and relaxation of common-slope restrictions. Our regressions also show that public debt growth is more inflationary in high-debt countries than in low-debt countries and that the debt-inflation linkage is weak in fixed or managed exchange rate regimes. A panel VAR traces out the transmission mechanism that a positive innovation to debt has a positive and persistent effect both on inflation and money supply. Wealth effects of public debt could also affect inflation, as hypothesized by the FTPL, but our study does not find supporting evidence.

The findings highlight challenges for price stabilization in highly indebted countries. They point to a significant risk of a debt-inflation trap, potentially large budgetary costs of noncredible disinflation policy, and limitations of sustained sterilized interventions in stabilizing prices and exchange rates. They also stress the importance of institutional and structural factors in the debt-inflation link, such as fiscal rules, inflation targeting, and the depth and breadth of the financial sector. They also indicate that, notwithstanding an important role of monetary policy in managing inflation expectations, fiscal policy would likely be the dominant factor for inflation in highly indebted developing countries. This implies that price stability achieved mainly through the issuance of central bank open market instruments (that is, accumulation of public debt) in lieu of deficit monetization could be sustained only if supported by fiscal consolidation and structural reforms to boost monetary policy independence.

Further research could be usefully undertaken in several areas. The link between inflation and economic growth has been extensively investigated both in empirical (for example, Barro, 1996; Ghosh and Phillips, 1998) and
theoretical studies (Smith and van Egteren, 2005). Our findings of a direct link between public debt growth and inflation could shed further light on the effect of fiscal policy on economic growth. In addition, our rich empirical findings could be utilized to fine-tune debt sustainability analysis. Finally, our empirical framework could be modified to assess the impact of debt structures, in particular currency and maturity, on inflation dynamics although data limitations would be a major challenge.

APPENDIX I. RELATIONSHIP BETWEEN PRICE, MONEY, DEBT, AND OUTPUT

A simplified version of Castro, De Resende, and Ruge-Murcia (2003) can be used to derive a functional relationship between price on the one hand and money, debt, and output on the other. In this simple version, a representative household is endowed with fixed resources, y, for each period, and allocates its real wealth among real consumption (c), real domestic money (m/p), and nonindexed real government bonds (b/p) in order to maximize the following utility function:

\[ \sum_{t=0}^{\infty} \beta^t (\ln(c_t) + \gamma \ln(m_t/p_t)), \]

subject to a resource constraint of

\[ c_t + \frac{m_t}{p_t} + \frac{b_t}{p_t} = y_t - \tau_t + \frac{m_{t-1}}{p_t} + \frac{i_{t-1}b_{t-1}}{p_t}, \]

where \( \tau_t \) is the lump-sum tax and \( i_{t-1} \) is a nominal gross return of a government bond between periods \( t-1 \) and \( t \). This maximization problem yields the following standard first-order conditions for consumption and real money demand, respectively:

\[ \frac{c_{t+1}}{c_t} = \frac{\beta b_t}{\pi_{t+1}}, \]

\[ \frac{m_t}{p_t} = \frac{\gamma c_t i_t}{i_t - \pi}, \]

where \( \pi_t = p_{t+1}/p_t \). These two first-order conditions nest a Cagan-type money demand function, which is inversely related to inflation expectations.

The government is faced with the following intertemporal budget constraint:

\[ G_t + (i_{t-1} - 1) \frac{B_{t-1}}{p_t} = \tau_t + \frac{(M_t - M_{t-1})}{p_t} + \frac{(B_t - B_{t-1})}{p_t}, \]

Forward iteration on Equation (A.5) and no-Ponzi game conditions on the government imply the following long-term budget constraint of the government:

\[ i_{t-1} \frac{B_{t-1}}{p_t} = \sum_{j=0}^{\infty} \frac{\tau_{t+j}}{R_{t+j}} = \sum_{j=0}^{\infty} \frac{G_{t+j}}{R_{t+j}} + \sum_{j=0}^{\infty} \frac{M_{t+j} - M_{t+j-1}}{p_{t+j} R_{t+j}}, \]

where \( G \) is the real government spending and \( R_{t+j} \) is the compounded real discount rate, as expressed as \( R_{t+j} = \prod_{h=1}^{j} r_{t+h} \) where \( r_{t+h} \) is the exogenous real interest rate between periods \( t + h - 1 \) and \( t + h \). In the case of a fiscal policy rule of backing a part, \( (1 - \delta) \), of the debt service by future primary surpluses and monetizing the remainder \( \delta \), we obtain the
following money supply function:

$$\frac{M_t}{P_t} = \frac{i_t - 1}{i_t} \left[ \frac{\delta i_{t-1} B_{t-1}}{p_t} + \frac{M_{t-1}}{p_t} - \sum_{j=1}^{\infty} \frac{M_{t+j}}{p_{t+j}} \frac{i_{t+j} - 1}{i_{t+j}} \right]. \quad (A.7)$$

Equation (A.7) shows that the path of money supply is determined by the extent of debt monetization (the first variable in the right) and savings in the future interest payments brought about by current monetary financing of the budget deficit (the third variable).

Imposing equilibrium conditions on Equations (A.4) and (A.7) and exploiting the recursive nature of the Euler equation in (A.3), we obtain the equilibrium price as follows:

$$p_t = \frac{(1 - \beta)(M_{t-1} + \delta i_{t-1} B_{t-1})}{\gamma C_t}. \quad (A.8)$$

Given the recursive nature of the equilibrium and no arbitrage between bond and real asset returns ($r_{t+1} = i_t/(p_{t+1}/p_t)$), the equilibrium price can be rearranged to:

$$p_t = \frac{(1 - \beta)(M_t + \delta B_t)}{\gamma C_t}. \quad (A.9)$$

**APPENDIX II. DATA SOURCES AND DEFINITIONS AND COUNTRY GROUPING**

Our main data set is a panel data set spanning 71 countries over up to 43 years, collected from a variety of sources. The main data set includes annual data for CPI, money, public debt, and real GDP of each country for the maximum period of 1962–2004. Country selections were based primarily on the availability of the data and hence exclude many African countries and some small Caribbean countries. Data for inflation and real GDP—a proxy for real consumption—are mostly from the International Financial Statistics (IFS), but, in some cases, are from the World Economic Outlook data set of the IMF. Public debt data are from a variety of sources, including the IFS, World Economic Outlook and OECD databases, and, for Jamaica, Russia, and Turkey, the country authorities. Our debt data refer to both foreign- and domestic-currency-denominated debt, and cover general government debt in most developed countries and consolidated central government in most developing countries. Given that fiscal autonomy of local governments, in particular borrowing rights, is limited in developing countries, we do not think that the use of narrower definitions of public debt for developing countries affects our findings materially. Monetary data are mainly from the IFS and the World Economic Outlook database, and, in the case of the euro-zone countries, the OECD. The definition of money is reserve money, or the narrowest definition available in the databases. Appendix Table A1 shows the definitions and sources for public debt and money for each country.

In addition to the four main variables, several other data were used for alternative specifications and various robustness tests. These include exchange rate regimes (Reinhart and Rogoff, 2004), exchange rates (IFS), and output gap estimates (derived from detrended real GDP using the H-P filter).

Countries are divided into 13 major advanced economies and 58 other countries, based on the classification of the IMF’s World Economic Outlook (IMF, 2005). The other countries include 48 developing countries and 10 nonmajor advanced economies as defined in the World Economic Outlook such as Korea, Israel, and Ireland, which could be considered as developing countries in a broad sense. This classification is broadly in line with other studies on fiscal variables and inflation (Catao and Terrones, 2005), which reported some evidence of significant heterogeneity between developed and developing countries.
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<th>Public Debt (in foreign or local currencies)</th>
<th>Money</th>
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<th>Money</th>
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<td>Israel</td>
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<td>New Zealand</td>
<td>IFS (1962–99)/WEO (2000–05)</td>
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<td>Switzerland</td>
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### Other emerging market and developing economies

#### Net credit countries

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<th>Source</th>
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#### Net debtor countries

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<th>Monetary Aggregation</th>
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<td>WEO</td>
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<td>Trinidad and Tobago</td>
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<td>IFS</td>
<td>Reserve money</td>
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</tbody>
</table>

1IFS refers to government financial statistics in the IMF’s International Financial Statistics (IFS) database (line 88). WEO refers to data from IMF’s World Economic Outlook database. WHD refers to data from IMF’s Western Hemisphere Department, using the broadest definition of government debt in each country.
Further groupings have been made on the basis of the extent of each country’s external balance positions and financing sources in line with the World Economic Outlook criteria. Subgroups have also been made on the basis of inflation rates and the size of public debt.

- Developing countries—creditors vs. debtors: If the sum of current account balances over 1972–2005 (1992–2005 for transition economies) is positive, the country is a net creditor; and if the result is negative, the country is a net debtor.
- Debtor developing countries—by financing sources: If the net external borrowing from official creditors\(^6\) are on average over 2001–05 equal or larger than 66 percent of the net external financing,\(^7\) then the country’s financing source is official financing. The source is private financing if the ratio is less than 33 percent, and is diversified financing if the ratio is in between.
- High-inflation and low-inflation countries: Countries with average annual inflation higher than the group median rate are high-inflation countries. Those with below median inflation rates are low-inflation countries.
- High-debt and low-debt countries: Countries with average public debt-to-GDP ratios higher than the group median rate are high debt countries and the others low debt countries.

**APPENDIX III. DEBT-INFLATION TRAP AND DEBT SUSTAINABILITY**

A rise in inflation will eventually push up nominal interest rates, which will in turn increase public debt unless countered by a higher primary surplus. This feedback effect implies that budgetary costs of rising inflation expectations rise more than proportionally to the increase in inflation expectations. This point can be illustrated by simple debt dynamic accounting as follows:

\[
\frac{\Delta B_t}{B_t} = R_t - \frac{S_t}{B_t}, \quad (A.10)
\]

where \(B\) is public debt, \(R\) is the interest rate, and \(S\) is the primary surplus. If the interest rate is set in line with inflation expectations \(\pi_t^e\) and the primary surplus in percent of GDP is predetermined,\(^8\) the debt dynamics can be simplified as follows:

\[
\frac{\Delta B_t}{B_t} = \left(\pi_t^e + r\right) - \frac{S_t}{Y_t} / \frac{B_t}{Y_t} = \left(\pi_t^e + r\right) - C, \quad (A.11)
\]

where \(C = (S_t/Y_t)/(B_t/Y_t)\). In a steady state of no change in the debt-to-GDP ratio, \(C\) is constant. If inflation expectations \(\pi_t^e\) rise in a proportion to debt growth \(\Delta B_t/B_t\) in line with our empirical findings, \(\pi_t^e = \alpha(\Delta B_t/B_t) + \beta X + \epsilon\), then \(\Delta B_t/B_t = (\beta X + \epsilon + r - C/1 - \alpha)\).

---

\(^6\)Official debt securities + liabilities to official creditors + balance on the capital account.

\(^7\)Current account balance—direct investment abroad—reserve assets—portfolio investment assets—other investment assets—errors and omissions.

\(^8\)These are strong, simplified assumptions that hardly hold in reality in the current form since most revenues and expenditures are likely to be affected by contemporaneous inflation and inflation expectations. Persson, Persson, and Svenson (1998) presents, for example, a calibrated model where changes in inflation and inflation expectations affect government revenues and expenditures significantly due to a variety of indexation schemes in tax rules and expenditure arrangements. In their model, changes in inflation expectations do not necessarily lead to simultaneous and equal changes in interest rates.
Hence, an increase in inflation expectations (as embodied in a jump in \( \varepsilon \)) raises debt not only directly (through an immediate increase in the borrowing cost) but also indirectly (through a multiplier effect resulting from the debt-inflation nexus).

An alternative way of looking at this is to see the implications on the debt-stabilizing levels of the primary surplus \( (S_t^*) \). The levels can be represented as follows:

\[
\frac{S_t^*}{Y_t} = \left[ \frac{R_t B_t - B_t \Delta Y_t}{Y_t} \right] = B_t \left( \frac{R_t - \Delta Y_t}{Y_t} \right)
\]

\[
= B_t \left( (1 + \pi_t^e)(1 + r) - (1 + \pi_t)(1 + g_t) \right)
\]

\[
= B_t \left( (\pi_t^e - \pi_t) + (r - g_t) \right).
\]

(A.12)

Given that inflation expectations, \( (\pi_t^e) \), could rewritten as

\[
\pi_t^e = \frac{\beta X + \varepsilon + r}{1 - \alpha} - \frac{\bar{\alpha} C}{1 - \alpha}
\]

(A.13)

it follows that the debt-stabilizing primary surplus could be rearranged to the following simplified form:

\[
\frac{S_t^*}{Y_t} \approx B_t \left( \beta X + \varepsilon + r \right) + (1 - \alpha) B_t \left( -\pi_t + r - g_t \right).
\]

(A.14)

This means that rising inflation expectations (as embodied in a jump in \( \varepsilon \)) would elevate the debt-stabilizing level of the primary surplus more than the same percentage decline in real GDP growth would. Our regression results for the debt-inflation link place \( \alpha \) in the range of \( \frac{1}{4} \) (mean group estimator) to \( \frac{1}{2} \) (GMM estimator). This implies that the effect of rising inflation expectations could be larger than the effect of a decline in real GDP by as much as one-third to one-half.

REFERENCES


Net Capital Flows, Financial Integration, and International Reserve Holdings: The Recent Experience of Emerging Markets and Advanced Economies

WOON GYU CHOI, SUNIL SHARMA, and MARIA STRÖMQVIST

The paper examines the link between net capital flows and international reserves emphasizing the external financing of reserve accumulation in the context of increasing international financial integration. The paper finds that the effect of net capital flows on reserve accumulation has shifted from negative to positive for emerging markets but not for advanced countries. The empirical results suggest that in recent years emerging markets, with concerns about sudden stops in capital flows, have rapidly built up reserves through external financing with net capital inflows, whereas the advanced countries, with more secure access to international finance, have balanced reserves accumulation with investments in higher-yielding foreign assets. [JEL E50, G10]


*Woon Gyu Choi is a senior economist at the IMF Institute; Sunil Sharma is the director of the IMF–Singapore Regional Training Institute; and Maria Strömqvist, a doctoral student at the Stockholm School of Economics, was a summer intern at the IMF Institute in 2005. The authors thank David Cook, Enrica Detragiache, Michael Devereux, Robert Flood, Brenda Gonzalez-Hermosillo, Leslie Lipschitz, Enrique Mendoza, Jaihyun Nahm, Jorge Roldos, Kwanho Shin, Evan Tanner, and an anonymous referee for helpful comments. The authors are also grateful to participants at the conference on “Korea and the World Economy” held in Seoul in 2006, and to the IMF Institute’s weekly seminar for discussions. Si-Yeon Lee provided valuable research assistance in an earlier stage of the project, and Anastasia Guscina assisted in the collection of data.
Global holdings of international reserves have increased rapidly in recent years. This increase has been especially dramatic in emerging markets, both in absolute terms as well as in comparison to the reserves held by advanced countries. At the end of 2005, the average reserves-to-GDP ratio reached 19 percent in emerging markets, compared with a ratio of 10 percent in the advanced countries. Emerging markets have accumulated reserves well above the levels suggested by traditional rules of thumb based on current account transactions and short-term external liabilities. Also, the recent record pace of reserve accumulation in emerging markets is at odds with the prediction of a standard reserve holding model (IMF, 2003).

With increasing financial liberalization and openness to cross-border transactions, managing a country’s liquid assets to facilitate current and future international transactions—what we call “sovereign liquidity”—has become a key element in macroeconomic management. Clearly, the desired level of reserves and the availability of liquidity depend on a sovereign’s access to international capital markets. This paper examines the relationship between reserve holdings and net capital flows (capital inflows minus capital outflows) in the changing global financial environment.

In recent decades, currency and/or financial crises accompanied by reversals in capital flows have become more frequent and severe. With increased financial integration, countries are more vulnerable to contagion from within and outside their regions (for example, Kaminsky and Reinhart, 2000). In response, central banks in developing countries have accumulated reserves to cushion extreme events, the bunching of external debt maturities, or other shocks that could affect the foreign exchange market and the domestic economy.

Holding large reserves is costly, but the perceived cost may be small relative to the economic and social cost of a crisis. Many of the currency and financial crises of the past 10 years have been associated with the contractionary effects of currency depreciation, with substantial output losses, especially through balance sheet channels (for example, Choi and Cook, 2004; and Frankel, 2005). After the Asian financial crisis, emerging markets have reduced short-term external debt, and stockpiled reserves to reduce their vulnerability to a crisis (for example, Aizenman and Marion, 2004; Rodrik, 2006; and Jeanne, 2007). Because a reserve holding country can always opt not to use its reserves for debt service, reserves have an insurance value specific to the country (Van Wijnbergen, 1990). Recently, Durdu, Mendoza, and Terrones (forthcoming) suggested that financial

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1Before the 1990s, emerging market reserves fluctuated between 3 and 4 months of imports. At the end of 2005, they stood at an average of 5.8 months of imports. Even in terms of the Guidotti – Greenspan rule that countries hold reserves to cover short-term external debt (Rodrik, 2006), emerging market reserve cover has been very high: for example, the ratio was greater than 3.0 for Korea and 6.5 for China in 2005. However, recent research suggests that a broader metric should be used for assessing reserve adequacy (see Lipschitz, Messmacher, and Mourmouras, 2006).
Globalization with endogenous sudden stops can explain the large buildup of reserves. This paper examines the “external financing” of a reserve buildup using net capital flows. Given frictions and information problems in international capital markets, countries facing uncertain growth prospects and volatile capital flows have a stockpiling motive in reserve management, even at the cost of increased external debt. Sovereign liquidity holdings are analogous to corporate liquidity holdings in that they help cope with uncertain income streams and cash flows. In the face of external finance premia, the behavior of corporations suggests that the value of having liquid assets is disproportionately high for financially weak firms (Kim, Mauer, and Sherman, 1998; and Almeida, Campello, and Weisbach, 2004). This external financing view of a reserve buildup is in contrast with New Mercantilist views (for example, Dooley, Folkerts-Landau, and Garber, 2004) that a reserve buildup is financed by internal cash flows from current account surpluses as a by-product of a strategy to protect competitiveness in countries dependent on exports for output growth.

With increasing financial integration, and greater exposure and dependence on international capital, sovereign liquidity management has become crucial for macroeconomic stability. The fact that more capital does not flow from rich to poor countries—the paradox of too little flow (Lucas, 1990)—might be substantially attributable to credit-market imperfections (Reinhart and Rogoff, 2004b). For the high-growth emerging markets, the paucity of capital has been ameliorated by financial globalization. Nonetheless, emerging markets may take advantage of the wave of capital inflows to stockpile reserves, as in other times external financing may be expensive due to credit-market imperfections. With increased capital and financial account transactions, the risk of capital flow reversals that can result in huge output losses has increased the option value of holding reserves. As sovereign liquidity shortages may lead to expensive borrowings or a forced reduction in consumption and investment, countries have an incentive to fend off the risk of binding liquidity constraints in the future by hoarding reserves today. Clearly, countries that can borrow at reasonable risk premia will need a smaller stock of reserves, and the differential in reserve holdings across countries should depend on the degree of access to international financial markets.

To examine how the availability of external finance has affected the link between net capital flows and the accumulation of reserves, we estimate panel data models over the period 1980–2005 for 36 emerging markets and 24 advanced economies. Financial globalization is captured by using the stock value of foreign assets and liabilities taken from Lane and Milesi-Ferretti (2006) or by period dummies. The analysis reveals that the effect of net capital flows on reserves for emerging markets has changed dramatically over time, and their reserve holding pattern has been quite different from that of advanced countries. In the 1980s, reserve holdings were negatively associated with net capital flows for emerging markets, while such a negative link was
less pronounced for advanced countries. In recent years, net capital flows have had a strong positive effect on reserves for emerging markets but a negative effect for the advanced countries. This finding supports our external financing view. In the face of increasing financial globalization, with heightened concerns about the risks of “sudden stops” (Calvo, 1998) and loss of access to international capital markets, emerging markets have built up sovereign liquidity when funds are available at low costs.

I. Determinants of International Reserves

Buffer Stocks and the Precautionary Motive

The pure buffer stock/precautionary model focuses on the opportunity cost, the adjustment cost, and volatility. First, the standard measure of the opportunity cost is the differential between the country’s own-interest rate and the interest rate on comparable U.S. treasuries. Most empirical studies, however, do not find a significant negative effect for the opportunity cost (Flood and Marion, 2002; IMF, 2003; and Aizenman and Marion, 2004). Second, if the import share in output is a proxy for the marginal propensity to import, and thus its inverse is an indication of the required output adjustment to produce a particular level of reserves, then the relationship would be negative as countries can use their reserves to reduce such adjustment costs. However, the relationship between reserves and the import share is ambiguous (Heller and Khan, 1978). If the import share represents the economy’s openness and vulnerability to external shocks (Edwards, 1984; and Aizenman and Marion, 2004), the relationship would be positive as the more open is a country to external shocks, the greater is the need for reserves. Also, the use of the ratio of reserves to imports as an indicator of reserve adequacy for facilitating current account transactions implies that the relationship is positive. Third, reserve holdings increase with the volatility of international transactions (Frenkel and Jovanovic, 1981; Flood and Marion, 2002; and Aizenman and Marion, 2004). This volatility is usually measured by the standard deviation of the trend-adjusted changes in reserves over some period. An alternative measure is the volatility of export receipts used by Edwards (1985) and Aizenman and Marion (2004).

Flood and Marion (2002) find that buffer-stock reserve models work about as well in the modern floating exchange rate period as they did during the Bretton Woods period. The IMF (2003), using a standard buffer stock model based on Aizenman and Marion (2004), suggests that the rapid

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2Aizenman and Marion (2004) note that the opportunity cost variable may not be properly measured because the composition of reserves is not adequately reflected, and until the early 1990s, most emerging markets did not have market-determined domestic interest rates.

3Such a reserve volatility measure, however, can be contaminated because it combines jumps in reserves owing to reserve restocking and sudden declines with speculative attacks (Flood and Marion, 2002).
accumulation of reserves in emerging markets between 1997 and 2001 was broadly in line with fundamentals, but the surge in reserves in 2002 and 2003 was above the level predicted by the model. Recent studies have examined the rapid reserve accumulation in emerging markets over the last 5 years, focusing on the risk of sudden reversals in capital flows. It has been suggested that emerging markets may hold large reserves with a precautionary motive to provide self-insurance against a capital account crisis (Aizenman and Marion, 2003; Aizenman and Lee, 2007; and Jeanne, 2007).


We propose that financial market imperfections and the risk of financial distress at a country level give rise to a precautionary motive of sovereign liquidity buildup when liquidity flows are available at low costs. In a loose analogy, the finance literature suggests that corporate liquidity demand responds to cash flows under financial market frictions: Almeida, Campello, and Weisbach (2004) show that financially constrained firms have a propensity to use cash flows to build up cash reserves or liquid assets because such assets enable a firm to reduce its dependence on costly external funding for future production activities—a precautionary or stockpiling motive.

The paper emphasizes the external financing of reserves in contrast with the New Mercantilism that typically concerns internal financing only. The relationship between reserves and net capital flows can be viewed in light of the identity in the balance of payments: “current account balance plus net capital flows (the latter being the capital and financial account balance) equals the change in reserves (a positive value means an increase).” This identity implies three types of relationships between net capital flows and the stock of reserves: (a) a negative (zero) correlation when current account balance is partially (fully) offset by net capital flows—current account financing by net capital flows; (b) a positive correlation when net capital flows determine whether reserves increase or decrease regardless of current account balance (that is, current account balance is more than fully offset by net capital flows or current account balance and net capital flows move in the same direction); and (c) no systematic correlation when there is little capital mobility. This identity suggests two sources of financing reserves. Types (a) and (c) pertain to the “internal” financing of reserve accumulation using cash flows from current account surpluses. This internal financing of reserves is in line with the New Mercantilist view that emphasizes export-driven reserve accumulation. Type (b) pertains to the “external” financing of a reserve buildup using net capital flows, and this is likely to become more important with increasing financial globalization.

The external financing view suggests that sovereign liquidity demand will be related to country cash flows (that is, net capital flows), but the relationship between sovereign liquidity and net capital flows may shift over time. The cost of restocking reserves through selling assets or expanding liabilities depends on a country’s access to international capital markets and
global financial conditions: in bad times countries with low creditworthiness, facing restricted access to international capital markets, find it more expensive to borrow than in good times. Financial globalization reduces barriers to cross-border capital flows and increases access to international financial markets, and also raises the potential risk of country cash flows. Developing countries have a heightened concern about sudden reversals in capital flows and the risk of financial crises (Calvo, 1998; Aizenman and Marion, 2003; Edwards, 2004; and Caballero and Panageas, 2005). Emerging markets, concerned about sudden stops in capital flows, may have increased their reliance on net capital flows to build up reserves in the face of financial globalization. In contrast, advanced countries have been able to raise external funds at relatively low costs through international financial markets. With increased financial globalization, advanced countries have become more active external liquidity providers and lowered their reserve levels relative to GDP. Therefore, sovereign liquidity demands may entail time-varying responses to net capital flows.

The demand for reserves also depends on a country’s credit quality as it affects the cost of external finance and the degree of access to international markets. Countries with low credit ratings face higher cash flow risk and have to pay a higher rate for external funds compared with countries with greater creditworthiness. Thus, lower rated countries will have a stronger incentive to build up reserve buffers to withstand adverse external shocks compared with higher rated countries, consistent with the precautionary liquidity demand story that liquidity holdings are positively related to cash flow risk (Kim, Mauer, and Sherman, 1998; and Almeida, Campello, and Weisbach, 2004). Country credit quality can be measured by sovereign ratings, which summarize and supplement information on macroeconomic indicators, default history, as well as social and political factors (see, for example, Cantor and Packer, 1996). Thus, sovereign ratings should have a negative impact on reserve holdings. Note that higher reserves could reduce the risk of sovereign default (Ben-Bassat and Gottlieb, 1992; and Mellios and Paget-Blanc, 2006), implying a positive correlation between reserves and sovereign rating. In our empirical analysis we take account of the endogeneity of sovereign ratings.

Exchange Rate Flexibility and Other Considerations

Exchange rate flexibility: Conventional wisdom holds that greater exchange rate flexibility should reduce reserves because central banks then do not need

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4The traditional buffer stock model assumes that the liquidation cost of assets for restocking reserves is known and Bar-Ilan, Marion, and Perry (2007) suggest that reserve adjustment costs are important in explaining the time path of reserve accumulation.

5As creditors in sovereign debt markets have limited enforcement rights (for example, few possibilities for seizing country assets), political and credit risks play an important role in capital flows (Reinhart and Rogoff, 2004b).
a large reserve stockpile to maintain a peg or to enhance the peg’s credibility (Flood and Marion, 2002; and Aizenman and Marion, 2004). However, with volatile capital flows, the need for reserves to temper asset price fluctuations may increase with exchange rate volatility. Also, central banks, to dampen the appreciation of their currencies (New Mercantilist views), may accumulate reserves (Frankel and Dornbusch, 1995; and Dooley, Folkerts-Landau, and Garber, 2004).

Other considerations: First, the literature on the interest rate–exchange rate nexus suggests that countries facing the risk of capital flight use domestic interest rates to counter exchange rate movements and prevent capital outflows and/or a lowering of reserves. Second, the net effect of the fiscal stance on reserves will depend on a combination of political-economy factors and fiscal vulnerabilities. Aizenman and Marion (2003, 2004) argue that a greater chance of opportunisti behavior by future policymakers or political corruption reduces the demand for reserves, and may lead to higher external borrowing for increasing the current consumption of special interest groups, suggesting a positive link between reserves and fiscal surpluses. Conversely, as fiscal vulnerabilities increase the risk of sovereign default and a crisis, countries in fiscally fragile situations may hold larger reserve cushions, implying a negative link between reserves and fiscal surpluses (Reinhart and Rogoff, 2004b). Lastly, to obtain higher returns countries may prefer to invest national resources in foreign assets other than reserves (Devereux and Sutherland, 2007).

II. Empirical Reserve Demand Models and Estimated Results

Data

The empirical analysis uses annual data for 36 emerging markets and 24 advanced countries for the 1980–2005 period (Appendix I). Most of the data have been collected from the IMF’s International Financial Statistics and World Economic Outlook databases, and from the CEIC. The descriptive statistics are reported in Appendix II—for further details, see Choi, Sharma, and Strömqvist (2007). As in IMF (2003) and Aizenman and Marion (2004), international reserves are defined as gross reserves net of gold. Reserves relative to GDP over the entire period are higher for emerging markets (12.5 percent) than for advanced countries (9.6 percent). Net capital flows, measured by the capital and financial account of the balance of payments, are higher for emerging markets than for advanced countries by about 2.5 percent of GDP. Sovereign ratings from Standard and Poor’s are higher on average for the advanced countries (22 compared with 14 for emerging markets).

Figure 1 (panel A) depicts the averages for reserves and the reserves-to-GDP ratio by country group for the 1980–2005 period. In terms of the reserves-to-GDP ratio, emerging markets caught up with the advanced countries in the early 1990s. Thereafter, while for advanced countries this ratio fluctuated around 10–11 percent, for emerging markets it rose
substantially after the Asian crisis until it reached 20 percent by 2003, and then slowed down to 19 percent by 2005. Panel B shows the different patterns of net capital flows for the country groups. For emerging markets, as capital accounts were liberalized, net capital flows increased steadily over the 1991–96 period but fell dramatically during the 1997–98 Asian crisis period. For advanced countries, net capital flows showed a downward trend since 1991, a
sharp increase after the Asian crisis, which reflects the funding of the large U.S. current account deficits. When net capital flows are normalized by GDP, the United States does not change the pattern of net capital flows for advanced countries, even though due to its size it has an important effect on the average level of net capital flows for advanced countries. As shown in panel C, emerging markets exhibit a downswing in the current account for 1991–96 (which mirrors an upswing in net capital flows), and a rising trend in the current account during recent years. For advanced countries, the current account relative to GDP fluctuates around a rising trend but was on average negative until the early 1990s. The negative current account average for advanced countries since 1998 is attributable to the U.S. current account deficit: excluding the United States from the sample, the cross-sectional average has been positive.

Following Lane and Milesi-Ferretti (2006), we consider two de facto measures of global financial integration based on actual capital flows: one is the ratio of total foreign assets and liabilities to GDP, and the other is the sum of the stock of portfolio equity assets and liabilities and the stock of foreign direct investment (FDI) assets and liabilities to GDP. Over 1970–2004, this measure gradually increased during the 1970s and 1980s but accelerated, especially in advanced countries, during the mid-1990s. Figure 2 depicts the cross-section average of the Lane and Milesi-Ferretti measures for the country groups that we use in this paper. For advanced countries, both measures show an upward trend with a sharp increase starting in 1998. For emerging markets, they show a gradual increase over the 1980s, some fluctuations with no apparent trend over the 1990s, and then a modest rise in recent years.
Figure 3 shows the group means of the reserve/GDP ratios under different exchange rate regimes. On the basis of the de facto regime classification used by Reinhart and Rogoff (2004a), regimes are divided into three categories: a fixed regime (index ¼ 1), an intermediate regime (index ¼ 2), and a floating regime (index ¼ 3). For emerging markets, consistent with the conventional wisdom the more flexible regimes are associated with lower reserves. For advanced countries, the floating regime is associated with the lowest reserves-to-GDP ratio, but the intermediate regime has the highest ratio, perhaps reflecting that more reserves are required for ameliorating exchange rate volatility under a closely managed float. Notably, the reserves-to-GDP ratio shows an upward trend under all regimes for emerging markets but only under the intermediate regime for advanced countries. Such shifts suggest that factors other than exchange rate regimes have an important effect on reserves.

Note: The figure depicts the group means of the reserve-to-GDP ratio by regime type over subperiods for emerging markets, advanced countries, and the pooled sample. Regimes are indexed as follows: fixed regime ¼ 1; intermediate regime ¼ 2; and floating regime ¼ 3 (see footnote 6). The subperiods are whole (1980–03); Y80_90 (1980–90); Y91_96 (1991–96); Y97_00 (1997–2000); and Y01_03 (2001–03).

Our index corresponds to the Reinhart-Rogoff classification as follows: index value 1 has categories from “no separate legal tender” to “de facto peg,” index value 2 has categories from “pre-announced crawling peg” to “managed floating,” and index value 3 has categories from “freely floating” to “freely falling.” We exclude sample observations for the category: “dual market in which parallel market data is missing.”
Reserve Regressions

Our regression model incorporates the aforementioned determinants of reserve holdings including sovereign ratings and net capital flows.\(^7\) For country \(i\) at time \(t\), the ratio of reserves to output in regression model 1 is given by

\[
\left( \frac{IR}{GDP} \right)_{i,t} = \alpha_i + \beta_S SIZE_{i,t} + \beta_{\sigma_C} \sigma_C^{i,t} + \beta_{\sigma_T} \sigma_T^{i,t} + \beta_{IM} \left( \frac{IM}{GDP} \right)_{i,t} + \beta_{SR} SR_{i,t} + \sum_{j=1}^{k} \beta_{CF,j} D_j \times \left( \frac{CF}{GDP} \right)_{i,t} + \varepsilon_{i,t},
\]

where \(IR/GDP\) is the ratio of international reserves to GDP (gross domestic product in U.S. dollars), and \(SIZE\) is the natural log of population, a supplementary scale factor in the model as we scale the dependent variable, reserves, by GDP. \(\sigma_C\) and \(\sigma_T\) are volatility measures of the growth rate of the nominal exchange rate. We take the growth rate of the exchange rate to eliminate the level effects in the nominal exchange rate due to currency denominations and control for the nonstationarity of the exchange rate. Cross-country volatility, \(\sigma_C\), is measured by the standard deviation of the growth rate of the nominal exchange rate for country \(i\) over the sample period. Time-varying volatility, \(\sigma_T\), is measured by the conditional standard deviation predicted from a GARCH (1,1) model for emerging markets and an ARCH(2) model for advanced countries—the exchange rate growth series is typically more persistent (yet stationary) for emerging markets than for advanced countries. In these models, we control for fixed effects by allowing for different means across countries and, for the sake of parameter parsimony, impose the same slope coefficient within the same country group. \(IM/GDP\) is the import share in GDP, and \(SR\) is the sovereign rating assigned by Standard and Poor’s. The parameter \(\alpha\) denotes fixed-country effects, and \(\varepsilon\) is the error term.

\(CF/GDP\) is the ratio of net capital flows to GDP. If net capital flows are used mainly for financing current account deficits (stockpiling reserves), \(\beta_{CF}\) will be negative (positive). To account for the effect of financial globalization on reserves, we allow for the effect of net capital flows to vary over subperiods. For emerging markets, net capital flows fell drastically during the Asian financial crisis of 1997–98, reversed in 2001, and increased thereafter (Figure 1). Financial globalization was gradual during the 1980s but accelerated in the mid-1990s for advanced countries (Figure 2). We thus...

\(^7\)Aizenman and Marion (2004) find the volatility of export receipts to be statistically insignificant. We exclude this variable from our regressions because the sign of its coefficient varies depending on which other variables are included in the model and is highly sensitive to normalizations.
interacted the capital flow variable with four time dummies: \( D_{(80–90)} \), \( D_{(91–96)} \), \( D_{(97–00)} \), and \( D_{(01–05)} \) took on the value 1 for 1980–90, 1991–97, 1998–2000, and 2001–05 respectively, and were zero otherwise.

The net capital flow and sovereign rating variables may be correlated with the error term due to reverse causality, which may arise, for example, when countries borrow from abroad to meet a reserves target and when ample reserve holdings have a favorable impact on the assessment of sovereign ratings. To deal with the endogeneity of regressors, we use instrumental variables (IV), with the lagged values of sovereign ratings and net capital flows as the instruments. A first-stage regression showed that the instruments were highly correlated with the endogenous variable. The Hansen test for overidentification suggested that the regression model was correctly specified and the instruments were valid. Statistical inferences about coefficient estimates are based on heteroscedasticity and autocorrelation consistent standard errors.

Table 1 summarizes regression model 1 for emerging markets and advanced countries, along with the pooled sample. For the pooled sample, when country heterogeneity is taken into account by fixed country effects, the adjusted \( R^2 \)s, \( R^2 \), show that the model explains about 88 percent of the variation in reserves. However, if the variation explained by the fixed country effect is excluded, \( R^2 \) drops to 57 percent, suggesting that a large part of the variation is picked by country-specific heterogeneity. For the group-specific regressions, the explanatory power of the model is much higher for advanced countries than for emerging markets, as indicated by their \( R^2 \)s (0.94 vs. 0.79). Excluding the fixed country effects, \( R^2 \) in the ordinary least squares (OLS) regressions drops to 0.41 for emerging markets and 0.72 for advanced countries, suggesting much larger heterogeneity in reserve holding behavior for the former group.

A key finding is that net capital flows have a time-varying effect on reserves that differs across emerging markets and advanced countries. For emerging markets, reserves decreased with net capital flows in the 1980s, suggesting that, before financial integration took off, net capital flows were used to finance current account deficits. In contrast, during the subsequent periods, net capital flows either had little effect or a positive effect on

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8To account for endogeneity (high reserve holdings may reduce exchange rate variability), we also used the lagged values of exchange rate regime dummies as instruments. The results were qualitatively the same.

9We ensure that the instruments are uncorrelated with the model error as follows: (i) an instrument is regressed on the endogenous regressor and country dummies to derive the component of the instrument that is uncorrelated with the regressor; (ii) the dependent variable is regressed on a set of exogenous variables including country dummies to derive the component of the dependent variable that contains the model error; and (iii) regress the filtered dependent variable from (ii) on the filtered instrument from (i), and test if the filtered instrument is statistically insignificant using a robust covariance matrix. Also, the Stock and Yogo (2005) test supports the validity of our instruments: the test rejected the null hypothesis of weak instruments at the 1 percent level for any three endogenous variables.
Table 1. Model 1: The Time-Varying Effects of Net Capital Flows on Reserves

<table>
<thead>
<tr>
<th>Independent Variables</th>
<th>Pooled OLS</th>
<th>Pooled IV</th>
<th>Emerging OLS</th>
<th>Emerging IV</th>
<th>Advanced OLS</th>
<th>Advanced IV</th>
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<tbody>
<tr>
<td>SIZE</td>
<td>0.246***</td>
<td>0.254***</td>
<td>0.319***</td>
<td>0.358***</td>
<td>0.132***</td>
<td>0.068</td>
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<tr>
<td></td>
<td>(5.74)</td>
<td>(5.02)</td>
<td>(5.64)</td>
<td>(5.08)</td>
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<td>(1.37)</td>
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<tr>
<td>( \sigma^C )</td>
<td>-0.920***</td>
<td>-0.928***</td>
<td>-0.001***</td>
<td>-0.007***</td>
<td>-0.615***</td>
<td>-0.547***</td>
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<tr>
<td></td>
<td>(-8.36)</td>
<td>(-8.26)</td>
<td>(-3.92)</td>
<td>(-3.60)</td>
<td>(-6.27)</td>
<td>(-5.69)</td>
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<tr>
<td>( \sigma^T )</td>
<td>0.002*</td>
<td>0.001</td>
<td>0.001**</td>
<td>0.001**</td>
<td>0.109</td>
<td>0.123</td>
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<td></td>
<td>(1.62)</td>
<td>(1.30)</td>
<td>(2.29)</td>
<td>(2.09)</td>
<td>(0.64)</td>
<td>(0.64)</td>
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<tr>
<td>IM/GDP</td>
<td>0.232***</td>
<td>0.249***</td>
<td>0.238***</td>
<td>0.161**</td>
<td>0.131*</td>
<td>0.144**</td>
</tr>
<tr>
<td></td>
<td>(4.75)</td>
<td>(4.98)</td>
<td>(3.49)</td>
<td>(2.01)</td>
<td>(1.75)</td>
<td>(2.18)</td>
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<tr>
<td>SR</td>
<td>-0.004*</td>
<td>-0.005**</td>
<td>-0.004***</td>
<td>-0.007**</td>
<td>-0.013***</td>
<td>-0.015***</td>
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<td></td>
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<td>(-2.18)</td>
<td>(-2.15)</td>
<td>(-2.17)</td>
<td>(-3.51)</td>
<td>(-3.60)</td>
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<tr>
<td>( D_{(80-90)}(CF/GDP) )</td>
<td>0.121</td>
<td>0.134</td>
<td>-0.488**</td>
<td>-1.086**</td>
<td>-0.006</td>
<td>-0.320**</td>
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<td>(0.21)</td>
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<td>(-2.32)</td>
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<td>( D_{(91-96)}(CF/GDP) )</td>
<td>0.151*</td>
<td>0.344**</td>
<td>0.103</td>
<td>0.145</td>
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<td>-0.561**</td>
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<td>(1.69)</td>
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<td>( D_{(97-00)}(CF/GDP) )</td>
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<td>0.118</td>
<td>0.265</td>
<td>-0.166*</td>
<td>-0.442***</td>
</tr>
<tr>
<td></td>
<td>(-1.09)</td>
<td>(-1.02)</td>
<td>(1.27)</td>
<td>(1.06)</td>
<td>(-1.72)</td>
<td>(-2.87)</td>
</tr>
<tr>
<td>( D_{(01-05)}(CF/GDP) )</td>
<td>0.072</td>
<td>0.127</td>
<td>0.438***</td>
<td>0.835***</td>
<td>-0.333***</td>
<td>-0.502***</td>
</tr>
<tr>
<td></td>
<td>(0.99)</td>
<td>(1.02)</td>
<td>(5.25)</td>
<td>(2.67)</td>
<td>(-2.75)</td>
<td>(-4.11)</td>
</tr>
<tr>
<td>( \bar{R}^2 )</td>
<td>0.889</td>
<td>0.877</td>
<td>0.795</td>
<td>0.781</td>
<td>0.936</td>
<td>0.931</td>
</tr>
<tr>
<td></td>
<td>[0.597]</td>
<td>[0.596]</td>
<td>[0.413]</td>
<td>[0.388]</td>
<td>[0.724]</td>
<td>[0.723]</td>
</tr>
<tr>
<td>No. of observations</td>
<td>897</td>
<td>897</td>
<td>407</td>
<td>407</td>
<td>490</td>
<td>490</td>
</tr>
<tr>
<td></td>
<td>[0.895]</td>
<td>[0.830]</td>
<td>[0.830]</td>
<td>[0.522]</td>
<td>[0.522]</td>
<td>[0.522]</td>
</tr>
</tbody>
</table>

Note: This table shows the results of regression model 1 for emerging markets, advanced countries, and the pooled sample. The reserves-to-GDP ratio (IR/GDP) is regressed on the log of population (SIZE), cross-sectional volatility (\( \sigma^C \)), time-varying exchange rate volatility (\( \sigma^T \)), import share (IM/GDP), sovereign rating (SR), and net capital flows (CF/GDP) interacted with period dummies (D). The conditional standard deviation predicted from a GARCH (1,1) model for emerging markets and that from an ARCH(2) model for advanced countries were merged to measure \( \sigma^T \) for the pooled sample. Both OLS and IV regressions include fixed-country effects. Instruments for IV estimations are the one-period-lagged values of sovereign ratings and the two lags of each of the net capital flow variables. The z-ratios in parentheses are based on standard errors robust to heteroscedasticity and autocorrelation (Bartlett kernel; bandwidth = 2). ***, **, and * indicate significance at the 1, 5, and 10 percent levels, respectively.

\( \bar{R}^2 \) is the adjusted R-squared, and the figure in brackets excludes variance explained by the fixed-country effects.

Hansen’s test is an overidentification test for all instruments. Under the null hypothesis that the model is correctly specified and the instruments are valid, the test statistic is distributed as a chi-square with the degree of freedom equal to the number of overidentifying restrictions (with p-values in square brackets).
reserves. During the 1991–96 period, a substantial net flow of capital into emerging markets led to only small increases in reserves, reflecting the use of net capital flows primarily to finance domestic expenditures. A positive but insignificant coefficient for 1997–2000 may reflect the mix of capital flight that caused reserve losses during the Asian crisis and the return of capital that helped the restocking of reserves in its aftermath. The relatively high and significant coefficient for 2001–05 implies that net capital flows led to a substantial accumulation of reserves. In contrast, for advanced countries, net capital flows were associated with lower reserve levels relative to GDP, especially in recent years. This suggests that financial globalization helped emerging markets, which had greater concern about the risk of sudden stops after the Asian financial crises, to build up their reserve levels through external financing, whereas it helped advanced countries to become more active external liquidity providers, reducing reserves.

For variables other than net capital flows, the analysis shows, first, that the population coefficient is positive and larger for emerging markets than for advanced countries. One could argue that the need for reserves relative to output initially increases as an economy grows to a certain threshold, and then flattens out or even declines. Second, the cross-country exchange rate volatility coefficient is negative, consistent with Flood and Marion’s (2002) finding that countries with greater flexibility in the exchange rate hold lower reserves. The coefficient estimate is much smaller for emerging markets than for advanced countries, partly because the measured volatility on average is much larger for the former than for the latter (see Appendix II). The coefficient on time-varying exchange rate volatility has a positive sign, suggesting that reserve holdings increase to moderate exchange rate volatility: it is significant at the 5 percent level for emerging markets but insignificant for advanced countries. Third, the import share coefficient is positive for all country groups. In addition, the sovereign rating has a negative effect on reserves, implying that the higher a country’s creditworthiness, the less the need for reserve holdings, with a more pronounced effect for advanced countries than for emerging markets.

In the pooled regressions, net capital flows do not have any discernable effects owing to their opposing effects in the two country groups, except that they had somewhat significant positive effects in the 1990s before the Asian crisis. As the main point of this paper is to show how the effects of capital flows have differed across emerging markets and advanced countries, we

---

10 To account for the financial dimension of international transactions, we also use the M2-GDP ratio in place of the population variable. Like the population variable, the M2-GDP ratio has a statistically significant positive effect only for emerging markets. This result suggests that the level of monetization is not a determinant of reserve holdings for advanced economics, which already have relatively well developed financial markets.

11 Since the data on sovereign ratings are often not available for emerging markets in the 1980s, we also estimated regressions for the 1990–2005 period. The results were very similar to those reported in this paper.
henceforth report results only for the two different country groups. Also, we
do not report the regressions including the standard opportunity cost
measure of reserves (the differential between the domestic interest rate and
the rate of return on U.S. treasuries), because, as in previous studies, we did
not find that it had any discernible effect on reserves.\(^{12}\)

The Lane and Milesi-Ferretti measure of financial globalization is
employed in place of the time dummies interacted with the capital flow
variable in regression model 2:

\[
\left( \frac{IR}{GDP} \right)_{i,t} = \alpha_i + \beta_S \cdot SIZE_{i,t} + \beta_{cC} \cdot \sigma_i^C + \beta_{aT} \cdot \sigma_i^T + \beta_{IM} \left( \frac{IM}{GDP} \right)_{i,t} + \beta_{SR} \cdot SR_{i,t} \\
+ \beta_{CF} \left( \frac{CF}{GDP} \right)_{i,t} + \beta_{GL} \left[ GLOB_{i,t} \times \left( \frac{CF}{GDP} \right)_{i,t} \right] + \epsilon_{i,t}, \tag{2}
\]

where \(GLOB_{i,t}\) denotes the logarithm of the country-specific globalization
measure for country \(i\) at time \(t\). We use two alternative country-specific
globalization measures: the two-year average of the ratio of total foreign
assets and liabilities to GDP; and the two-year average of the ratio of the sum
of the stock of equity assets and liabilities and the stock of FDI assets and
liabilities to GDP (see Appendix II for descriptive statistics). As shown in
Table 2, the two alternative measures give the same qualitative results. The
effect of net capital flows per se—measured by \(\beta_{CF}\)—is significantly negative
for emerging markets but positive for advanced countries.\(^{13}\) Importantly,
the coefficient net capital flows interacted with the financial globalization
measure (\(\beta_{GL}\)) is significantly positive for emerging markets but negative
for advanced countries. This finding reinforces the earlier results: the effect of net
capital flows on reserve accumulation drifts up over time (from negative to
positive) for emerging markets, and shifts down over time for advanced
countries. The effects of other variables are similar to those in regression
model 1.

Next, by first differencing Equation (2) and including lagged dependent
variables, we also estimate a dynamic panel regression model to account for
possible dynamic adjustments of the reserves-GDP ratio to its time-varying

---

\(^{12}\)To account for the cost of acquiring international currencies for building up reserves,
we also used the EMBI and EMBI plus spreads for Latin American and non-Latin American
countries for 1992–2005 as a proxy for the opportunity cost in emerging markets—the IV
estimate was negative but not statistically significant.

\(^{13}\)The average total effect of net capital flows during a period can be measured by \(\beta_{CF} + \beta_{GL} \cdot GLOB^{AVE}\), where \(^{AVE}\) denotes the period average value. In particular, for 2001–05, the
total effect based on the IV estimator in panel B is \(0.127 (=-1.338 + 0.436 \times 3.36)\) and \(-0.506\)
\((=-1.448-0.398 \times 4.91)\) for emerging markets and advanced economies, respectively.
However, its effect at very low levels of financial integration as in the 1980s can be positive
for advanced economics, compared to a negative effect during the 1980s in Table 1. Thus, for
advanced economics, the effect of net capital flows on reserves is mixed in the early stages of
financial integration.
Table 2. Model 2: Effects of Net Capital Flows and Financial Globalization

<table>
<thead>
<tr>
<th>Independent Variables</th>
<th>Emerging</th>
<th>Advanced</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>OLS</td>
<td>IV</td>
</tr>
<tr>
<td><strong>A. Globalization measure based on total foreign assets and liabilities</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>SIZE</td>
<td>0.289***</td>
<td>0.288***</td>
</tr>
<tr>
<td></td>
<td>(4.47)</td>
<td>(3.99)</td>
</tr>
<tr>
<td>(\sigma^C)</td>
<td>-0.003***</td>
<td>-0.003***</td>
</tr>
<tr>
<td></td>
<td>(-5.16)</td>
<td>(-5.12)</td>
</tr>
<tr>
<td>(\sigma^T)</td>
<td>0.001**</td>
<td>0.001**</td>
</tr>
<tr>
<td></td>
<td>(2.49)</td>
<td>(2.53)</td>
</tr>
<tr>
<td>IM/GDP</td>
<td>0.298***</td>
<td>0.294***</td>
</tr>
<tr>
<td></td>
<td>(4.62)</td>
<td>(4.27)</td>
</tr>
<tr>
<td>SR</td>
<td>-0.004**</td>
<td>-0.004</td>
</tr>
<tr>
<td></td>
<td>(-2.25)</td>
<td>(-1.56)</td>
</tr>
<tr>
<td>CF/GDP</td>
<td>-1.467***</td>
<td>-1.579***</td>
</tr>
<tr>
<td></td>
<td>(-3.42)</td>
<td>(-2.55)</td>
</tr>
<tr>
<td>GLOB*(CF/GDP)</td>
<td>0.343***</td>
<td>0.362***</td>
</tr>
<tr>
<td></td>
<td>(3.79)</td>
<td>(2.81)</td>
</tr>
<tr>
<td>(\overline{R}^2) 1</td>
<td>0.810 [0.461]</td>
<td>0.810 [0.437]</td>
</tr>
<tr>
<td>No. of observations</td>
<td>356</td>
<td>356</td>
</tr>
<tr>
<td>Hansen’s test: (\chi^2(1)) 2</td>
<td>0.296 [0.587]</td>
<td>0.077 [0.782]</td>
</tr>
</tbody>
</table>

**B. Globalization measure based on the stock of equity and foreign direct investment**

|                       |          |          |          |          |
|                       | OLS      | IV       | OLS      | IV       |
| SIZE                  | 0.305*** | 0.303*** | 0.087*   | 0.030    |
|                       | (4.95)   | (4.66)   | (1.83)   | (0.59)   |
| \(\sigma^C\)         | -0.003***| -0.003***| -0.460***| -0.358***|
|                       | (-5.53)  | (-5.62)  | (-4.57)  | (-2.99)  |
| \(\sigma^T\)         | 0.001**  | 0.001*** | 0.183    | 0.183    |
|                       | (2.52)   | (2.82)   | (1.11)   | (1.05)   |
| IM/GDP                | 0.325*** | 0.293*** | 0.116*   | 0.119*   |
|                       | (4.78)   | (4.00)   | (1.63)   | (1.94)   |
| SR                    | -0.003** | -0.003   | -0.014***| -0.018***|
|                       | (-1.89)  | (-1.49)  | (-4.44)  | (-5.12)  |
| CF/GDP                | -0.639***| -1.338***| 1.243*** | 1.448*** |
|                       | (-3.23)  | (-3.40)  | (5.12)   | (2.80)   |
| GLOB*(CF/GDP)         | 0.252*** | 0.436*** | -0.318***| -0.398***|
|                       | (4.48)   | (4.31)   | (-5.18)  | (-3.60)  |
| \(\overline{R}^2\) 1| 0.818 [0.459] | 0.810 [0.427] | 0.946 [0.813] | 0.943 [0.810] |
| No. of observations   | 356      | 356      | 461      | 461      |
| Hansen’s test: \(\chi^2(1)\) 2 | 2.023 [0.155] | 0.399 [0.527] |

Note: In this table, regressors include net capital flows (CF/GDP) interacted with a constant and the logarithm of the globalization measure (GLOB) as specified in regression model 2. Both OLS and IV regressions include fixed-country effects. Instruments for IV estimations are the one-period lagged values of sovereign ratings, the two lags of net capital flows, and the globalization measure (GLOB) multiplied by the one-period lagged CF/GDP (as we treat GLOB as exogenous). Luxembourg was excluded from the sample since the globalization measure was available only from 2000. The \(z\)-ratios in parentheses are based on standard errors robust to heteroscedasticity and autocorrelation (Bartlett kernel; bandwidth = 2). ***, **, and * indicate significance at the 1, 5, and 10 percent levels, respectively.

1See footnote one in Table 1.
2See footnote two in Table 1.
determinants. The cross-sectional exchange rate volatility reflects a “fixed” characteristic over the sample period. The population variable also shows little variability around a trend over time and is not suitable for a dynamic setting. Hence, variables for fixed country effects, population, import share, and cross-sectional exchange rate volatility are dropped from the dynamic panel model. The dynamic panel counterpart of Equation (3) is represented by regression model 3:

\[ \Delta \left( \frac{IR}{GDP} \right)_{i,t} = \sum_{j=1}^{2} \beta_j \Delta \left( \frac{IR}{GDP} \right)_{i,t-1} + \beta_{\sigma} \Delta \sigma_{i,t}^T + \beta_{SR} \Delta SR_{i,t} + \beta_{CF} \Delta \left( \frac{CF}{GDP} \right)_{i,t} + \beta_{GL} \Delta \left[ GLOB_{i,t} \times \left( \frac{CF}{GDP} \right)_{i,t} \right] + u_{i,t}, \]  

(3)

where the sovereign rating and net capital flows are assumed to be endogenous. We estimate this model, using Arellano and Bond’s (1991) generalized method of moments (GMM) estimator.

Table 3 shows how financial globalization may have changed the effect of net capital flows on reserve accumulation. The results are generally in line with those obtained in the OLS and IV regressions in Table 2, suggesting that the effect of net capital flows on reserves has increased with financial integration for emerging markets, whereas the converse is true for advanced economies. The coefficient on the one-period lagged dependent variable has a statistically significant coefficient of 0.775, indicating substantial persistence in reserve ratios. The effect of exchange rate volatility (over time) on reserves is positive and significant at the 5 percent level for both country groups. Sovereign ratings had a negative effect on reserves for both country groups, indicating that smaller reserve cushions were needed by economies with higher ratings. The Wald test statistics indicate the joint significance of the regressors. As the Arellano-Bond estimator assumes first-order autocorrelation and no second-order autocorrelation of the residuals, the serial correlation tests do not indicate misspecification.

Some Additional Hypotheses

In this subsection, we examine some additional hypotheses related to reserve accumulation. First, we show that the external financing view is relevant even if there is some truth to the new mercantilism (that countries accumulate reserves as a by-product of resisting currency appreciation to guard competitiveness). In doing this, we examine whether countries with current account surpluses and appreciating currencies have accumulated more reserves than other countries. Table 4 (columns 2–3) indicates that the effects of net capital flows are largely unaffected by exchange rate appreciations and current account balances. Current account surpluses, currency appreciations, and net capital inflows after the Asian crisis are associated with a
replenishing of reserve stocks in emerging markets, whereas they have been associated with a lowering of reserves in advanced economies.

Second, we consider how monetary policy may have affected reserve accumulation. To this end, we include a “nexus” variable defined as the spread between domestic and foreign interest rates relative to exchange rate growth. In doing this, we control for the (overall) fiscal balance because the attractiveness of higher interest rates may be affected by the presence of fiscal liabilities (Flood and Jeanne, 2005). Table 4 (columns 4–5) shows that the nexus coefficient is positive and significant for emerging markets but insignificant for advanced economies, implying that an increase in the spread relative to exchange rate growth helped emerging markets increase reserve holdings, likely through attracting capital inflows. On the other hand, the fiscal balance coefficient is not statistically significant for emerging markets, perhaps implying that the positive effect of political-economy factors counterbalances the negative effect associated with the precautionary

<table>
<thead>
<tr>
<th>Table 3. Model 3: Dynamic Panel Regressions</th>
</tr>
</thead>
<tbody>
<tr>
<td>Independent Variables</td>
</tr>
<tr>
<td>Δ(IR/GDP)_{-1}</td>
</tr>
<tr>
<td>(7.78)</td>
</tr>
<tr>
<td>Δ(IR/GDP)_{-2}</td>
</tr>
<tr>
<td>(-1.09)</td>
</tr>
<tr>
<td>ΔσT</td>
</tr>
<tr>
<td>(2.54)</td>
</tr>
<tr>
<td>ΔSR</td>
</tr>
<tr>
<td>(-2.63)</td>
</tr>
<tr>
<td>Δ(CF/GDP)</td>
</tr>
<tr>
<td>(-0.59)</td>
</tr>
<tr>
<td>Δ[GLOB*(CF/GDP)]</td>
</tr>
<tr>
<td>(2.75)</td>
</tr>
<tr>
<td>Constant</td>
</tr>
<tr>
<td>(3.74)</td>
</tr>
<tr>
<td>Wald test</td>
</tr>
<tr>
<td>No. of observations</td>
</tr>
<tr>
<td>Arellano-Bond tests: order 1</td>
</tr>
<tr>
<td>order 2</td>
</tr>
</tbody>
</table>

Note: This table shows the results of regression model 3 for emerging markets and advanced countries over the period 1980–2004 using the Arellano and Bond (1991) generalized method of moments. The last year in the sample is dictated by the availability of the globalization measure that is based on the sum of the portfolio equity assets and liabilities and the stock of foreign direct investment assets and liabilities (the other measure based on total assets and liabilities gave similar qualitative results). The sovereign rating and net capital flows are treated as endogenous. Luxembourg was excluded from the sample since the globalization measure was available only from 2000. The z-ratios in parentheses are based on standard errors robust to heteroscedasticity. Significance at the 1, 5, and 19 percent level is shown by ***, **, and *, respectively. The statistics for the Arellano-Bond tests are based on the null hypothesis of no autocorrelation of order 1 and 2 (with p-values in parentheses).
## Table 4. Instrumental Variables Regressions with Additional Factors

<table>
<thead>
<tr>
<th>Independent Variables</th>
<th>Currency Appreciation and Current Account Status</th>
<th>Interest Rates-Exchange Rate Nexus and Fiscal Stance</th>
<th>Foreign Asset Position</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Emerging</td>
<td>Advanced</td>
<td>Emerging</td>
</tr>
<tr>
<td><strong>SIZE</strong></td>
<td>0.329***</td>
<td>0.067</td>
<td>0.491***</td>
</tr>
<tr>
<td></td>
<td>(4.83)</td>
<td>(1.32)</td>
<td>(4.21)</td>
</tr>
<tr>
<td><strong>σC</strong></td>
<td>−0.003***</td>
<td>−0.560***</td>
<td>−0.003***</td>
</tr>
<tr>
<td></td>
<td>(−4.18)</td>
<td>(−5.54)</td>
<td>(−3.50)</td>
</tr>
<tr>
<td><strong>σT</strong></td>
<td>0.001**</td>
<td>0.128</td>
<td>0.001**</td>
</tr>
<tr>
<td></td>
<td>(2.08)</td>
<td>(0.62)</td>
<td>(2.01)</td>
</tr>
<tr>
<td><strong>IM/GDP</strong></td>
<td>0.168**</td>
<td>0.137**</td>
<td>0.096</td>
</tr>
<tr>
<td></td>
<td>(2.11)</td>
<td>(2.05)</td>
<td>(1.00)</td>
</tr>
<tr>
<td><strong>SR</strong></td>
<td>−0.006**</td>
<td>−0.015**</td>
<td>−0.008**</td>
</tr>
<tr>
<td></td>
<td>(−2.11)</td>
<td>(−3.63)</td>
<td>(−1.96)</td>
</tr>
<tr>
<td><strong>D(80-90)*(CF/GDP)</strong></td>
<td>−1.001**</td>
<td>−0.381**</td>
<td>−1.211</td>
</tr>
<tr>
<td></td>
<td>(−2.21)</td>
<td>(−2.23)</td>
<td>(−2.53)</td>
</tr>
<tr>
<td><strong>D(91-96)*(CF/GDP)</strong></td>
<td>0.240</td>
<td>−0.610**</td>
<td>0.285</td>
</tr>
<tr>
<td></td>
<td>(1.09)</td>
<td>(−2.17)</td>
<td>(0.89)</td>
</tr>
<tr>
<td><strong>D(97-00)*(CF/GDP)</strong></td>
<td>0.341</td>
<td>−0.410***</td>
<td>0.355</td>
</tr>
<tr>
<td></td>
<td>(1.39)</td>
<td>(−2.70)</td>
<td>(1.12)</td>
</tr>
<tr>
<td><strong>D(01-05)*(CF/GDP)</strong></td>
<td>0.935**</td>
<td>−0.560***</td>
<td>0.934**</td>
</tr>
<tr>
<td></td>
<td>(2.63)</td>
<td>(−4.23)</td>
<td>(2.26)</td>
</tr>
<tr>
<td><strong>I(appreciation)</strong></td>
<td>0.053***</td>
<td>−0.014**</td>
<td>0.053***</td>
</tr>
<tr>
<td></td>
<td>(4.02)</td>
<td>(−2.24)</td>
<td>(−1.00)</td>
</tr>
<tr>
<td>Nexus</td>
<td>0.001**</td>
<td>0.013</td>
<td></td>
</tr>
<tr>
<td>---------------</td>
<td>-----------</td>
<td>--------</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(2.12)</td>
<td>(0.70)</td>
<td></td>
</tr>
<tr>
<td>Fiscal balance</td>
<td>−0.050</td>
<td>−0.205**</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(−0.016)</td>
<td>(−2.29)</td>
<td></td>
</tr>
<tr>
<td>Foreign assets</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>0.023***</td>
<td>−0.007*</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(4.05)</td>
<td>(−1.65)</td>
<td></td>
</tr>
<tr>
<td>( R^2 ) i</td>
<td>0.797 [0.451]</td>
<td>0.932 [0.723]</td>
<td></td>
</tr>
<tr>
<td>No. of observations</td>
<td>407</td>
<td>490</td>
<td></td>
</tr>
<tr>
<td>Hansen’s test: ( \chi^2(4)^2 )</td>
<td>1.811 [0.771]</td>
<td>3.172 [0.529]</td>
<td></td>
</tr>
<tr>
<td></td>
<td>0.773 [0.436]</td>
<td>0.954 [0.374]</td>
<td></td>
</tr>
<tr>
<td></td>
<td>0.810 [0.435]</td>
<td>0.939 [0.854]</td>
<td></td>
</tr>
<tr>
<td></td>
<td>0.623 [0.891]</td>
<td>0.340 [0.952]</td>
<td></td>
</tr>
<tr>
<td></td>
<td>4.147 [0.386]</td>
<td>3.861 [0.425]</td>
<td></td>
</tr>
</tbody>
</table>

Note: This table reports the IV estimation results of regression model 1 with additional factors for emerging markets and advanced countries. Regressions in columns 2–3 include an indicator function for exchange rate appreciation interacted with another indicator function for current account balance: \( I_{\text{appreciation}} = 1 \) if exchange rate growth > 0, and zero otherwise; and \( I_{\text{surplus}} = 1 \) if the current account balance > 0, and zero otherwise. Regressions in columns 4–5 include a nexus variable defined as \( (1 + \text{the money market rate minus LIBOR})/(1 + \text{exchange rate growth}) \) and a fiscal balance variable measured by the overall fiscal balance to GDP ratio, and are estimated for 1990–2006 since the money market rate is often unavailable in the 1980s for emerging markets. Regressions in columns 6–7 include the ratio of total foreign assets net of reserves to GDP for the period 1980–2004. All regressions use instrumental variables (IV) and include fixed-country effects. The z-ratios in parentheses are based on standard errors robust to heteroscedasticity and autocorrelation (Bartlett kernel; bandwidth = 2). ***, **, and * indicate significance at the 1, 5, and 10 percent levels, respectively.

1See footnote one in Table 1.
2See footnote two in Table 1.
motive that stems from fiscal fragility. For advanced economies, the coefficient is negative and statistically significant, suggesting that the precautionary motive outweighs political-economy factors.

Third, to see how investment in foreign assets other than reserves affects reserve holdings, we used the ratio of foreign assets (net of reserves) to GDP from the Lane and Milesi-Ferretti data set. Table 4 (columns 6–7) shows that the coefficient on this ratio is positive for emerging markets but negative (significant at the 10 percent level) for advanced economies. This suggests that advanced economies, which are less liquidity-constrained than emerging markets, prefer to hold a smaller proportion of their foreign assets in reserves.

Fourth, we considered the influence of exchange rate regimes and world interest rates (for the detailed results, see Choi, Sharma, and Strömqvist, 2007). Consistent with Figure 3, emerging markets tend to hold smaller reserves under the floating regime than other regimes, whereas advanced economies tend to hold larger reserves under the intermediate regime than other regimes. The world interest rate (measured by LIBOR), a proxy for external financing costs, tends to be positively correlated with country-risk premia (for example, EMBI spreads) for emerging markets with limited access to international financial markets. For emerging markets, the world interest rate coefficient is significantly negative, suggesting that lower external financing costs in recent years may have contributed to higher reserve accumulation.

III. Concluding Remarks

This paper suggests that, despite greater financial integration and moves toward more flexible exchange rate arrangements, emerging markets have used capital inflows to build up large reserve stocks. In contrast, advanced economies, given their better access to international capital markets, have not shown this pattern. Our findings are consistent with the external financing view: emerging markets, as they integrate into the international financial system and face the risks of sudden stops in capital flows, have built up reserves after the Asian financial crisis, whereas advanced economies have balanced reserves accumulation with investments in higher yielding foreign assets.

An important issue in the context of global financial stability is how to make an assessment of comfortable reserve levels and put the savings of emerging markets to better use without compromising their financial stability. Arrangements among central banks for liquidity risk sharing (for example, through reserve pooling and currency swaps) and better access to international financial markets will help improve sovereign liquidity management in the face of potentially volatile capital flows. As emerging markets mature, they will probably have smaller reserve cushions and prefer to hold a larger proportion of their foreign assets in higher return investments. The increasing creation of sovereign wealth funds is already
leading to a move in this direction, and the adoption of investment and operational norms for such funds could accelerate the process. It would also be interesting in future research to examine how the sterilization operations conducted by emerging market central banks, to manage domestic monetary conditions and prevent exchange rate appreciations, have facilitated the accumulation of reserves.

Appendix I. Country Group List

The emerging market country group (36) comprises Argentina, Brazil, Bulgaria, Chile, China, Colombia, Croatia, Czech Republic, Egypt, Estonia, Hungary, India, Indonesia, Israel, Jordan, Kazakhstan, Latvia, Lithuania, Malaysia, Mexico, Pakistan, Peru, Philippines, Poland, Romania, Russia, Slovak Republic, Slovenia, South Africa, South Korea, Taiwan Province of China, Thailand, Turkey, Ukraine, Uruguay, and Venezuela. The advanced country group (24) comprises Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Hong Kong SAR, Ireland, Italy, Japan, Luxembourg, Netherlands, New Zealand, Norway, Portugal, Singapore, Spain, Sweden, Switzerland, United Kingdom, and United States. Our country groups correspond to the groups “emerging market countries” and “industrial countries” in the IMF Research Department’s Global Data Source. Our emerging market group is similar to the “emerging market countries” except that it includes Croatia, Egypt, Jordan, Kazakhstan, and Uruguay and excludes Hong Kong SAR and Singapore. Our advanced country group includes Hong Kong SAR and Singapore because their per capita incomes have been well above the sample mean of the advanced economies for at least the last 10 years. Also, their financial integration measures have been among the highest over the period considered.

Appendix II. Descriptive Statistics

See Table A1 for descriptive statistics of key variables.

<table>
<thead>
<tr>
<th>Table A1. Descriptive Statistics</th>
</tr>
</thead>
<tbody>
<tr>
<td>Variable</td>
</tr>
<tr>
<td>----------------------------------</td>
</tr>
<tr>
<td>A. Emerging markets</td>
</tr>
<tr>
<td>Reserves</td>
</tr>
<tr>
<td>Reserves/GDP</td>
</tr>
<tr>
<td>In (Population)</td>
</tr>
<tr>
<td>Exchange rate volatility</td>
</tr>
<tr>
<td>Import/GDP</td>
</tr>
<tr>
<td>Net capital flows/GDP</td>
</tr>
<tr>
<td>Current account/GDP</td>
</tr>
<tr>
<td>S&amp;P rating</td>
</tr>
<tr>
<td>GLOB (total)</td>
</tr>
<tr>
<td>GLOB (equity)</td>
</tr>
<tr>
<td>Fiscal balance/GDP</td>
</tr>
<tr>
<td>M2/GDP</td>
</tr>
</tbody>
</table>
REFERENCES


Table A1 (concluded)

<table>
<thead>
<tr>
<th>Variable</th>
<th>Units</th>
<th>N</th>
<th>Mean</th>
<th>SD</th>
<th>Min</th>
<th>Max</th>
</tr>
</thead>
<tbody>
<tr>
<td>B. Advanced economics</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Reserves</td>
<td>Billions of U.S. dollars</td>
<td>610</td>
<td>30.2</td>
<td>66.0</td>
<td>0.0</td>
<td>834.3</td>
</tr>
<tr>
<td>Reserves/GDP</td>
<td>Ratio</td>
<td>607</td>
<td>0.096</td>
<td>0.138</td>
<td>0.002</td>
<td>0.747</td>
</tr>
<tr>
<td>In (Population)</td>
<td></td>
<td>624</td>
<td>2.582</td>
<td>1.415</td>
<td>−1.022</td>
<td>5.698</td>
</tr>
<tr>
<td>Exchange rate volatility</td>
<td></td>
<td>622</td>
<td>0.096</td>
<td>0.039</td>
<td>0.000</td>
<td>0.146</td>
</tr>
<tr>
<td>Import/GDP</td>
<td>Ratio</td>
<td>624</td>
<td>0.453</td>
<td>0.385</td>
<td>0.071</td>
<td>2.159</td>
</tr>
<tr>
<td>Net capital flows/GDP</td>
<td>Ratio</td>
<td>608</td>
<td>−0.001</td>
<td>0.052</td>
<td>−0.281</td>
<td>0.156</td>
</tr>
<tr>
<td>Current account/GDP</td>
<td>Ratio</td>
<td>606</td>
<td>0.005</td>
<td>0.055</td>
<td>−0.192</td>
<td>0.290</td>
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<tr>
<td>S&amp;P rating</td>
<td>Number (1–23)</td>
<td>534</td>
<td>22.0</td>
<td>1.8</td>
<td>14.0</td>
<td>23.0</td>
</tr>
<tr>
<td>GLOB (total)</td>
<td></td>
<td>552</td>
<td>5.230</td>
<td>0.821</td>
<td>3.357</td>
<td>7.527</td>
</tr>
<tr>
<td>GLOB (equity)</td>
<td></td>
<td>552</td>
<td>3.823</td>
<td>1.136</td>
<td>1.145</td>
<td>6.525</td>
</tr>
<tr>
<td>Fiscal balance/GDP</td>
<td>Ratio</td>
<td>620</td>
<td>−0.041</td>
<td>0.441</td>
<td>−0.208</td>
<td>0.169</td>
</tr>
<tr>
<td>M2/GDP</td>
<td>Ratio</td>
<td>531</td>
<td>0.727</td>
<td>0.391</td>
<td>0.000</td>
<td>2.658</td>
</tr>
</tbody>
</table>

Note: This table shows descriptive statistics (units; number of observations (N); mean; standard deviations (SD), minimum (Min); and maximum (Max)) of key variables for 1980–2004 for two country groups in Appendix I. Population is in million persons. Reserves are defined as gross reserves net of gold in U.S. dollars. Exchange rate (cross-section) volatility is the standard deviation of nominal exchange rate growth, and import is measured by imports of goods and services. S&P rating is Standard and Poor’s sovereign ratings (annual average). GLOB (total) is the logarithm of the two-year average of the percentage ratio of total foreign assets and liabilities to GDP, and GLOB (equity) is the logarithm of the two-year average of the percentage ratio of the sum of the portfolio equity assets and liabilities and the stock of foreign direct investment assets and liabilities to GDP. Fiscal balance is overall fiscal balance.


Special Section: Current Account Sustainability in Major Advanced Economies

Introduction

AKITO MATSUMOTO
Special Section Editor


This issue of IMF Staff Papers presents five papers from the conference on “Current Account Sustainability in Major Advanced Economies” held at the University of Wisconsin on May 2–3, 2008. The conference was organized by Charles Engel and Menzie Chinn, sponsored by the Center for World Affairs and the Global Economy and the Robert M. La Follette School of Public Affairs, and cosponsored by the Department of Economics, the European Union Center of Excellence, and the Center for International Business Education and Research.

Engel and John H. Rogers challenge the traditional intertemporal approach to the current account by studying direct observations on the expectations drawn from a survey of forecasters. With these expectations data, the authors extend previous work dispensing with the assumption of unbiased expectations (sometimes termed the rational expectations hypothesis). They find that expected national consumption growth rates are driven largely by expected national income growth rather than world consumption growth or world real interest rates. These results cast additional doubt on models built assuming free trade in assets driven by consumption-growth differences.

Suparna Chakraborty and Robert Dekle develop a model to study the roles of productivity differentials and financial frictions in the recent
evolution of the U.S. current account deficit. They show that productivity
differentials alone cannot explain the deficit because higher productivity
growth in the rest of the world would have induced capital outflow from the
United States, and would have narrowed the current account deficit. They
show, however, that the lower cost of acquiring the U.S. assets abroad can
explain the recent current account pattern.

Carol C. Bertaut, Steven B. Kamin, and Charles P. Thomas present a
model-based evolution of the U.S. external position. They show that while
the U.S. current account deficit will continue to expand, it will likely take
many years before the size of its external debt poses a threat to the United
States’ ability to obtain financing from foreign investors. They also find little
evidence that countries’ indebtedness prompts abrupt changes in asset prices
and exchange rates to correct current account positions. While their findings
suggest that adjustment of the U.S. current account is not imminent, they
also warn that factors excluded from the model could trigger adjustment
sooner than predicted in their model.

Marcel Fratzscher and Roland Straub study the effect of asset price
movements on current accounts in structural vector autoregressive models.
They find that a 10 percent equity price increase in the United States relative
to the rest of the world will worsen the U.S trade balances by 0.9 percent of
GDP after 16 quarters, although the effect is different among G-7 countries.
They attribute these differences to cross-country differences in equity-driven
wealth effects on consumption.

Finally, Hamid Faruqee and Jaewoo Lee document the changes in the
cross-country distribution of external positions. They find that the global
dispersion of the current-account-to-GDP ratio is steadily increasing as trade
and financial integration progress. However, they also find that the recent
substantial expansion of the U.S. current account deficit—at the center of so-
called “global imbalances”—cannot be explained as a trend phenomenon.

While each paper has its own focus, all papers have the common theme
that current account sustainability is linked tightly to asset market
developments. Instead of explaining current account deficits based on
capital account surpluses or vice-versa, the papers provide explanations for
current account and capital account dynamics in an integrated fashion. It is
impossible to fully understand current account developments without
looking at asset market developments. It is not an accident that these
papers studying current account sustainability are published in this issue of
IMF Staff Papers, as the International Monetary Fund continuously
monitors balance of payment situations worldwide.

Finally, I want to state that it is quite an honor for me to be a guest editor
of a special section of IMF Staff Papers publishing such excellent papers
presented at a conference whose organizers, Charles Engel and Menzie
Chinn, are my mentors.
Expected Consumption Growth from Cross-Country Surveys: Implications for Assessing International Capital Markets

CHARLES ENGEL and JOHN H. ROGERS*

Survey data show that the expected growth rates of consumption across countries vary widely and are not highly correlated. This data contradict the simplest of open-economy models in which there is a freely traded non-state-contingent bond and purchasing power parity holds. This paper explores two alternative explanations for the finding: that households in each country in effect face different ex ante real interest rates or that there are significant credit constraints, so that expected consumption growth rates are driven largely by expected income growth. The empirical evidence strongly supports the latter hypothesis. These findings challenge the modeling of consumption that is at the heart of many, if not most, macroeconomic models. [JEL E21, F32, F41]


The present-value model (PVM) of the current account is an intuitively appealing approach in understanding international imbalances. The starting point of the model is the observation that a current account imbalance implies a change in a country’s net external asset position. A country running a surplus, for example, must be acquiring claims on the

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*Charles Engel is a Professor of Economics at the Department of Economics at the University of Wisconsin. John H. Rogers is Deputy Associate Director of the International Finance Division of the Board of Governors of the Federal Reserve System. The authors thank Ben Chiquoine for outstanding research assistance and Andrea Ferrero and Robert Kollmann for comments.
rest of the world. The model focuses on the intertemporal decision of households and firms. Specifically, a country will borrow when desired investment by firms exceeds desired saving by households. An essential building block of the model is the Euler equation for the household, which determines, in essence, expected consumption growth as a function of real interest rates.

Although the PVM is intuitively appealing, its empirical performance is spotty. Nason and Rogers (2006, p. 159) put it more bluntly: “Tests of the present-value model (PVM) of the current account are frequently flatly rejected by the data.” Examples of such rejections can be found in Sheffrin and Woo (1990a, 1990b), Otto (1992), and Ghosh (1995). Some modifications of the basic PVM in these early papers have been found to improve the fit of the model. Bergin and Sheffrin (2000) allow for time-varying real interest rates and real exchange rates, and find the fit of the model is improved. Similarly, Nason and Rogers (2006) find that allowing for a time-varying world real interest rate enhances the empirical performance of the model. Gruber (2004) modifies the standard representation of preferences to allow for habit persistence. This modification allows the PVM to produce a current account whose volatility is more in line with empirical observations. Kano (forthcoming) argues that the crucial factor in the failure of the PVM of the current account in early tests is that they failed to capture the consumption-tilting. According to Kano (2007), standard tests of the PVM are not able to distinguish between time-varying world real interest rates and habit persistence in preferences.1

A potential problem for the PVM is the Euler equation for household consumption. From a number of different sources, there appears to be little support for the Euler equation as a description of consumption behavior, at least using standard models of preferences. In the PVM, the ability of households to borrow and lend freely on international capital markets allows for smoothing of intertemporal income risk. The PVM derives the conclusion that consumption at any point in time should equal the annuity value of wealth, including lifetime discounted labor income. With a constant intertemporal rate of substitution (CIES), expected consumption growth approximately equals the real interest rate.

In the simplest versions of the PVM, if capital markets are well integrated, all households in the world face a common world interest rate. It follows that under the CIES assumption, expected consumption growth should be equalized across countries. Our paper uses a unique data set to

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1Engel and Rogers (2006) provide evidence in favor of a modified version of the present-value model to explain the U.S. current account deficit. That study makes use of the long-run forecasts of GDP used in the current paper. However, that paper does not use the forecasts of consumption used in this study, and it appears that those forecasts may not be consistent with some of the underlying equations in Engel and Rogers (2006). That paper, in essence, assumes that differences in expected consumption growth across countries can be attributed to expected real exchange rate changes.
provide some direct evidence on this building block of the PVM. Specifically, we make use of surveys of expected consumption growth conducted by Consensus Forecasts. We make use of surveys of economic forecasters in the G7 countries that twice a year (April and October) record the median forecasts of long-run consumption growth.

It is evident that expected consumption growth is not equalized across countries (see Figure 1). Each survey records expected consumption growth for the current year, each of the next five years, and then the expected average consumption growth rate for years 6–10 from the current year. From these surveys, we can construct expected 10-year consumption growth rates as the compounded annual expected growth rates. Figure 1 plots the 10-year expected growth rates from each survey. Clearly the forecasters did not expect consumption to grow at equal rates across countries. According to the figure, cross-country expected consumption growth differentials can be as large as 20 to 25 percentage points, two or more percentage points per year on average. In addition, there appears to be very little co-movement in the expected consumption growth rates, contrary to a basic prediction of the PVM of the current account.

One possible explanation of Figure 1 is that households in each of the G7 countries face different ex ante real interest rates. There are two possible general reasons why households in different countries might not be able to borrow and lend at the same real interest rate. The real interest differential between two countries can be written as the deviation from uncovered interest parity (which pertains to nominal interest rates and exchange rates) and the expected change in the real interest rate. Deviations from uncovered interest parity might arise because markets for interest-bearing securities are not perfectly integrated, or because there are risk considerations. On the other hand, even if capital markets are well integrated and risk premiums are low, there may be large real interest differentials if purchasing power parity (PPP) does not hold. If there are expected changes in the real exchange rates between a pair of countries, then the effective consumption real interest rates will be different in the two countries.

Although we cannot distinguish the cause for real interest differentials, our survey data provide a direct measure of the differential. That is because we can use a measure of the long-term (10-year) nominal interest rate in
combination with the survey’s measure of expected inflation over the upcoming 10 years. The same Consensus Forecast survey that collects forecasts of consumption growth also gathers inflation forecasts for a horizon that parallels the consumption forecasts. A possible explanation for Figure 1—the dispersion across countries of consumption growth forecasts—is that indeed household Euler equations hold, but that households in different countries face different real interest rates. Our data allow us to gauge how much of the differences in forecasts of consumption growth can be attributed to real interest differentials.

But perhaps consumption Euler equations (in conjunction with CIES preferences) are not a useful description of consumption behavior. Figure 2 hints at an alternative class of models for aggregate consumption. The Consensus Forecasts constructs measures of expected GDP growth, which again parallel the forecasts of consumption growth and inflation in terms of the horizon and coverage of the surveys. In Figure 2, we have plotted the
expected 10-year GDP growth rates and consumption growth rates for each of the G7 countries. It is clear that these forecasts move closely together.\footnote{It is important to realize that this is not simply a necessary outcome of national income accounting: $Y = C + I + G + NX$. The whole point of the present-value model of the current account is that well-functioning international capital markets break the link between current income and consumption. The difference gets reflected in the current account.}
But Figure 2 shows that expected consumption growth may be too closely related to expected income growth. There are a couple of possible explanations for such a tight link. It may be that many households are credit constrained. Borrowing and lending is not so easy, and consumption is more closely tied to current income than the PVM allows. Or, it may be that capital markets are available to households, but consumption behavior is not as forward looking as the PVM assumes. In either case, we might find that expected consumption growth is more closely tied to expected income growth than to ex ante real interest rates. In Figure 3 we plot expected consumption growth against a measure of an expected 10-year real interest rate for each country. The latter is constructed from the Consensus Forecasts of nominal interest rates and inflation rates. The correlation does not appear to be nearly as tight as that between expected consumption growth and expected real income growth.

The remainder of this paper tests more formally the propositions outlined above. We investigate how much of differences in expected consumption growth rates across countries can be explained by real interest differentials versus differences in expected income growth. As Figures 2 and 3 suggest, we will find a much stronger role for expected income growth. This undermines one of the key building blocks of the PVM.

I. Model Background

Here we review the implications of the PVM for cross-country expected consumption. Assume households have a period utility given by \( \frac{1}{(1-\rho)C_t^{1-\rho}} \). \( C_t \) is the consumption aggregate for the household. They maximize the expected present discounted value of utility, with a constant discount factor given by \( \beta \). As is well known, the Euler equation representing intertemporal substitution between consumption at time \( t \) and consumption in 10 years, \( C_{t+10} \), is given by

\[
C_t^{1-\rho} = \beta E_t(R_{t,10}^{-\rho}C_{t+10}^{1-\rho}).
\]

Here, \( R_{t,10} \) is the real rate of return on an interest-bearing bond with a maturity of 10 years.

Hereinafter, we will assume that we can approximate this Euler equation by

\[
C_t^{1-\rho} = \beta R_{t,10}^{-\rho}C_{t,t+10}^{1-\rho}.
\]

\( C_{t,t+10} \) refers to the expectation at time \( t \) of period \( t+10 \) consumption as taken from our survey data. \( R_{t,10} \) is the expected real return on a nominal bond with a 10-year maturity. As discussed above, we construct this measure by deflating the nominal yield on a 10-year bond by the survey measure of expected inflation. Section II provides details on the data construction. Appendix Figures A1–A4 elaborate on the implications of this approximation of the Euler equation.
With this approximation, we get

\[ \frac{C_{t,t+10}}{C_t} = \beta^{1/\rho} (R_{t,10})^{1/\rho}. \]  

(1)

If households in every country face the same real interest rate, and have the same preferences, they have the same expected consumption growth rate.
As noted earlier, the survey data are not consistent with this prediction. We consider two general types of explanations for the observed differences in expected consumption growth: the real interest rate is different across countries, because capital markets are not perfect or PPP does not hold so the real return is different; or, consumption is tied to income. Under the former assumption, all households maximize utility, and Equation (1) holds for everyone in each country.

Under the alternative assumption, consumption growth is tied to income growth. We model this in an ad hoc way. Perhaps consumers follow a simple rule of thumb, or perhaps there are capital market imperfections that tie consumption to current income. We let $C^R_t$ denote consumption of this group, and posit an equation of the form:

$$C^R_{t,t+10}/C^R_t = a + c(Y^R_{t,t+10}/Y^R_t).$$

We can nest the two theories in a general model, where a fraction $m$ follow Equation (1) and $1-m$ follow rules of thumb. Then, the aggregate expected growth rate of consumption is approximately a weighted average of the right-hand side of the two equations. Specifically, let $C^U_t$ be consumption of utility-maximizing consumers. Rewrite Equation (1) as:

$$C^U_{t,t+10} = \beta^{1/\rho}(R_{t,10})^{1/\rho} C^U_t = \beta^{1/\rho}(R_{t,10})^{1/\rho} \mu C_t.$$

Rewrite Equation (2) as

$$C^R_{t,t+10} = [a + c(Y^R_{t,t+10}/Y^R_t)]C^R_t = [a + c(Y^R_{t,t+10}/Y^R_t)](1 - \mu)C_t.$$

Adding (3) and (4) together, and dividing by $C_t$, we get

$$C_{t,t+10}/C_t = (1 - \mu)a + \mu\beta^{1/\rho}(R_{t,10})^{1/\rho} + (1 - \mu)c(Y^R_{t,t+10}/Y^R_t).$$

Equation (5) suggests the empirical specification we pursue. In each country, we ask whether we can explain expected consumption growth using expected real interest rates and expected output growth.

II. Data and Empirical Specification

Our data come from surveys conducted by Consensus Forecasts. Beginning in 1989, Consensus Forecasts has sent a form each month to professional forecasters inquiring about their outlook for several macroeconomic and financial market variables for the next two years. Each April and October, the survey participants are also asked about their forecasts at longer horizons, up to 10 years out; we use these long-horizons forecasts in this paper. Forms are sent the first Tuesday or Wednesday of the month, with a request that responses be received by the following Monday. The response deadline of the second Monday of the month is chosen because most announcements of monthly macroeconomic data occur in the first week of the month. Survey results are published on the Thursday after the Monday deadline. Consensus reports the mean forecast across survey respondents,
which number in the range of two or three dozen depending on the country.\(^5\)

The Consensus Forecasts survey data has been used in many academic papers, including Ghysels and Wright (forthcoming), Engel and Rogers (2006), and Engel, Mark, and West (2008).

We construct our measures of expected consumption growth, real interest rates, and expected output growth as follows.

### 10-Year Expected Consumption Growth

The April and the October surveys are treated differently depending on how the forecasts of current year variables are handled.

**April:** Take April 2006 as an example. The survey reports expected consumption growth rates for each year between 2006 and 2011, and then an average for 2012–2016. Let \( E_t G_{t+1}^C \) be the gross growth rate forecast for 2006. Similarly, \( E_t G_{t+1}^C \), \( E_t G_{t+2}^C \), ..., \( E_t G_{t+5}^C \), refer to the growth rates expected in 2006 for 2007–11. Then use the notation \( E_t G_{\infty}^C \) to denote the expected growth rate for the 2012–16 period. The 10-year expected growth rate for April 2006, \( G_{t,10}^C \), is

\[
G_{t,10}^C = (E_t G_{t+1}^C)(E_t G_{t+1,t+2}^C)(E_t G_{t+1,t+3}^C)(E_t G_{t+1,t+4}^C)(E_t G_{t+1,t+5}^C)
\]

\[
\times (E_t G_{t+5,t+6}^C)(E_t G_{t,1}^C)\]

**October:** Now take October 2006 as an example. Here, \( E_t G_{t,t+1}^C \) is the expected growth rate of consumption for 2007. In April of each year, the first expected growth rate is the expected growth rate for the current year, but for the October survey, the first expected growth rate is the expected growth rate for the next year. \( E_t G_{t+1}^C \), \( E_t G_{t+2}^C \), ..., \( E_t G_{t+5}^C \) refer to the expected growth rates for 2008–11. \( E_t G_{\infty}^C \) represents the expected consumption growth rate for the final 2012–16 period. Then we calculate the 10-year expected growth rate for October 2006, \( G_{t,10}^C \), as

\[
G_{t,10}^C = (E_t G_{t+1}^C)(E_t G_{t+1,t+2}^C)(E_t G_{t+1,t+3}^C)(E_t G_{t+1,t+4}^C)(E_t G_{t+1,t+5}^C)(E_t G_{t,1}^C)\]

### 10-year expected output growth

—Calculated the same way as expected consumption growth. Call this \( G_{t,10}^Y \).

### 10-year expected inflation

—Calculated the same way as expected consumption growth.

### 10-year nominal interest rate

—In both April and October of each year, Consensus Forecasts reports the actual 10-year bond yield on the date of the survey.

\(^5\)Note that the panel of forecasters is not identical across the G7 countries nor across time. However, many panelists do provide forecasts for several countries. Furthermore, although participants do come in and out of the survey, this does not occur very often. Response rates are in general very high (close to 100 percent), but can be somewhat smaller for the long-horizon forecasts and/or for some of the smaller countries.
10-year real interest rate—We take the 10-year gross nominal interest rate described above, and divide by the 10-year gross expected inflation described above. Call this $R_{t,10}$.

Equation (5) inspires the following empirical specification, which we estimate for each of the G7 countries individually:6

$$C_{t,t+10} / C_t = b_0 + b_1 (R_{t,10})^{1/\rho} + b_2 (Y_{t,t+10} / Y_t). \tag{6}$$

We do not estimate this regression over the entire sample, but instead perform rolling regressions using 10 years of data (and therefore 20 observations.) Our reasoning is that if indeed the consumption growth rates of the expected utility maximizing households and the rule-of-thumb households are different, then their weight $\mu$ in aggregate consumption might evolve over time.

Alternatively, the rolling regressions can be viewed as a simple way of determining whether the importance of the two explanatory variables has changed over time, or as a way of accounting for structural breaks in the series that are due, say, to revisions to data collection procedures.

We estimate Equation (6) using OLS regressions imposing $\rho = 2$ (a typical value used in macroeconomic calibrations). In addition to the country-by-country estimates, we undertake panel estimation of Equation (6). Here and in all panel estimates we impose $\rho = 2$ for all countries, impose that $b_1$ and $b_2$ are common across all countries, but allow the intercepts, $b_0$, to be country-specific; panel OLS standard errors are calculated.

Ultimately, we want to explain why $C_{t,t+10} / C_t$ is different than $C_{t,t+10}^{W} / C_t^{W}$ where the $W$ superscript refers to world expected consumption growth. That suggests the following specification, where we divide variables by $C_{t,t+10}^{W} / C_t^{W}$:

$$[(C_{t,t+10} / C_t) / (C_{t,t+10}^{W} / C_t^{W})] = d_0 + d_1 [(R_{t,10})^{1/\rho} / (C_{t,t+10}^{W} / C_t^{W})]$$

$$+ d_2 [(Y_{t,t+10} / Y_t) / (C_{t,t+10}^{W} / C_t^{W})]. \tag{7}$$

$C_{t,t+10}^{W} / C_t^{W}$ is a measure of the expected growth rate of G7 consumption, where G7 consumption growth is measured as a weighted average of consumption growth in each country.7

Under the standard model, $C_{t,t+10}^{W} / C_t^{W}$ should be proportional to $(R_{t,10})^{1/\rho}$, the world real interest rate (to the $1/\rho$ power) if each country faces the same real interest rate. Equation (7) then explains the deviation of expected consumption growth in each country from expected G7 consumption growth as a function of the deviation of each country’s real

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6This is similar to one of the regressions run by Campbell and Mankiw (1999), who use actual (U.S.) data, rather than forecasts.

7The weight applied to any country’s growth rate is essentially that country’s share of G7 real GDP. As in our earlier paper (Engel and Rogers, 2006) we use 1990 real exchange rates to convert nominal GDP shares to real shares; see that paper for more details.
interest rate from the world real interest rate implied by the model (again, taken to the $1/r$ power), and to expected GDP growth in each country relative to expected consumption growth in the G7.

As with Equation (6), we estimate Equation (7) country-by-country and as a panel, imposing $r = 2$ so that the model is linear.

In addition to Equations (6) and (7), we consider logarithmic specifications:

$$c_{t,t+10} - c_t = e_0 + e_1 \ln(R_{t,10}) + e_2(y_{t,t+10} - y_t),$$

and

$$[(c_{t,t+10} - c_t) - (c_{t,t+10}^W - c_t^W)] = f_0 + f_1 \ln[(R_{t,10})^{1/r}/(C_{t,t+10}^W/C_t^W)]$$

$$+ f_2[(y_{t,t+10} - y_t) - (c_{t,t+10}^W - c_t^W)],$$

where the lower-case $c$ and $y$ refer to logs of consumption and output. Finally, specification (9) suggests a more general equation in which the effect of world consumption growth on domestic consumption growth is not restricted:

$$c_{t,t+10} - c_t = g_0 + g_1 \ln(R_{t,10}) + g_2(y_{t,t+10} - y_t) + g_3(c_{t,t+10}^W - c_t^W).$$

As with Equations (6) and (7), we perform both country-by-country and panel estimation of each of these equations.

III. Empirical Results

The results for all of our specifications are displayed graphically in Figures 4–8. Each figure corresponds to results from one equation. (Figure 4 reports results of estimation of Equation (6), Figure 5 of Equation (7), and so on.) In each figure, we report first the panel estimates of the slope coefficients, followed by estimates of the slope coefficients for each country. We report the results graphically because we estimate rolling regressions, so we can conveniently report coefficient estimates and confidence intervals visually.

In Figure 4, the column on the left-hand side of the figures displays the rolling-sample estimates of $b_1$, the coefficient on GDP from Equation (6). The right-hand-side columns display the estimates of $b_2$, the coefficient on the real interest rate. Plus and minus two (OLS) standard error bands are also displayed. The first row of the figures displays the estimates of $b_1$ and $b_2$ for the panel of countries; each of the remaining rows presents estimates for one of the G7 countries, beginning with the United States in row two. The sample begins in 1990; hence the first estimate displayed is for the 10-year period 1990–99. The final estimate is for 1998–2007.

The results are striking. The coefficients on GDP growth are always positive and statistically significant. For many countries the estimate is near unity. In fact, except for the United States, the confidence interval contains unity for the estimates over most of the windows, and for all of the estimates
for some countries. The panel estimated coefficient on expected income growth is slightly below one, and the confidence interval never includes one, reflecting the role of the United States in pulling down the coefficient estimate. But even in the United States, the coefficient on expected income growth is quite high, in the range of 0.60.

On the other hand, for the real interest rate, the point estimates are more often negative than of the expected positive sign, and are rarely statistically different from zero in the country-by-country analysis. For the panel, there is
a small window where the coefficient is significantly different than zero, but it is significantly negative. These findings offer very little comfort to the theory that expected consumption growth in each country is determined by the local ex ante real interest rate.

Figure 5 displays the estimates from Equation (7), the specification where we transform the variables in the baseline specification to be ratios of world consumption growth. The transformation has no material effect on the estimated coefficients. The only noticeable change is that the coefficient on expected income growth in the U.S. regression is lower than in the unscaled
version of the model. However, this mechanically follows because the U.S. share of total “world” consumption is large.

In Figures 6 and 7 we display estimates of Equations (8) and (9), respectively, the log specifications of Equations (6) and (7). Once again the results are quite robust: expected consumption growth tracks expected GDP growth very tightly for each of the G7 countries. In many cases, the two series move nearly one for one. On the other hand, expected consumption growth is essentially uncorrelated with expected real interest rate in the G7.
As we have noted, the primary exception to the general pattern of results is the United States, where consumption and GDP growth move considerably less than one for one. Engel and Rogers (2006) present favorable results for a simple PVM of the current account in U.S. data, a finding that is not replicated (in unpublished work) for other G7 countries. It may appear, then, that U.S. consumers are more forward looking than consumers in the other countries. However, Figures 4–8 of the current paper

Note: As in Figure 4, with estimates from Equation (7).

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show that the correlation between U.S. consumption growth and real interest rate is zero or negative.

Finally, in Figure 8 we display results when we add world consumption growth to the right-hand side of the log specification whose results were displayed in Figure 6. In the country-by-country analysis, the general pattern of results continues to hold: the coefficients on expected income are significantly different from zero and usually near unity (except for the
United States), and the coefficients on real interest rates are generally not significantly different than zero. Usually the coefficient on world consumption growth is near zero. We note, however, that for some windows for a few countries, the coefficient is significantly positive, and in those windows the estimated coefficient on expected income is correspondingly lower. Finally, we note that in the panel estimates, the coefficient on world consumption and the world interest rate are generally
insignificantly different than zero, but the estimated coefficient on expected income growth remains significantly positive and close to unity.

Robustness: Results Using One-Year Expectations

Figures A1–A4 present the results of a check for robustness to constructing our expectations variables at one-year horizons rather than 10-year. Figure A1 is analogous to Figure 1, and hence depicts the expected
consumption growth rates by country. One-year ahead expected consumption growth rates also differ significantly across countries. Differentials are typically larger at the one-year horizon than at the 10-year horizon. This is as one would expect, as the 10-year ahead expectations are presumably less influenced by business cycle considerations.

Figure A2 plots the expected one-year GDP growth rates and consumption growth rates for each of the G7 countries. As seen in Figure 2 for the 10-year growth rates, these short-horizon forecasts of GDP and

Note: As in Figure 5, with variables now in logs.
consumption growth rate move closely together. Figure A3 plots expected consumption growth against the expected one-year real interest rate for each country. The correlation once again does not appear to be nearly as tight as that between expected consumption growth and expected real income growth.

Finally, Figure A4 presents the regression results for the specification in logs, Equation (8). Figure 6 presents the analogous results for the long-horizon expectations. We see that expected consumption growth has little to do with the real interest rate. It is connected closely to expected GDP growth, as observed in earlier results, though the coefficient with the one-year expectations is lower than unity and in many cases appears to be falling over time.
Comparing our results with our findings using 10-year expectations, we see that expected consumption growth over the one-year horizon is less responsive to expected income growth over the same horizon than was true for the 10-year expectations. That is, the coefficient on expected income growth in Equation (8) is significantly less than one for all countries in the one-year horizon regressions, but it is close to one for nearly all countries in the 10-year horizon regressions.

On the other hand, in neither the one-year nor the 10-year data is there much of a link between expected consumption growth and real interest rates. Frequently the coefficients are insignificantly different than zero, or negative. Apparently, intertemporal consumption decisions are not very much influenced by changes in real interest rates, at least to the extent that these survey measures of expectations reflect household expectations.

Note: As in Figure 6, adding log world consumption as a regressor.
This pattern of empirical results is puzzling from the standpoint of standard models of consumption. The model based on expected utility maximization relates the level of consumption to the expected present discounted value of income flows, and the expected growth in consumption to the real interest rate. Here, in essence, we find that there is a sort of consumption smoothing going on—the one-year expected growth rate of consumption and the one-year expected growth rate of income are less closely tied than the 10-year rates. But it is the expected growth rate of consumption in the long run, not the level, that is connected to the long-run expected growth of income. And the expected real interest rate seems to have little influence on consumption decisions at any horizon.

IV. Conclusions

Our survey data show that the expected growth rates of consumption across countries vary widely and are not highly correlated (Figure 1). These data contradict the simplest of open-economy models in which there is a freely traded non-state-contingent bond and PPP holds. We have explored two alternative explanations for the finding: The first posits that the inequality in expected consumption growth rates arises because households in each country in effect face different ex ante real interest rates. The second hypothesis is that there are significant credit constraints, so that expected consumption growth rates are not determined by ex ante real interest rates, but instead are driven largely by expected income growth. The empirical evidence strongly supports the latter hypothesis.

These findings are not just a challenge for open-economy macro-economists, because they challenge the modeling of consumption that is at the heart of many, if not most, macroeconomic models. We find that expected consumption growth is not determined by ex ante real interest rates. That relation is central in many macro models.

One point worth emphasizing is that these findings also pertain to the recent literature on the consumption–real exchange rate anomaly, or Backus-Smith puzzle. When agents can trade a complete set of contingent claims denominated in a single numeraire (such as a currency), but PPP does not hold, then if agents have power utility functions, the relative growth rates of consumption in two countries should equal the real exchange rate growth. This theory is a special case of the one we have examined. If markets are complete, actual ex post relative consumption growth should equal ex post real exchange rate growth. But in that setting, the relationship must obviously hold ex ante as well if expectations are rational: expected relative consumption growth should equal the expected real exchange rate change, which in turn will equal the ex ante real interest differential. This is the equation that we test.

See Backus and Smith (1993) for an early formulation of the problem. Corsetti, Dedola, and Leduc (2008) is a recent paper that addresses the puzzle.
There are many caveats to our study. The first is that our survey may not measure the market’s true expectations of consumption growth. The forecasters in our survey may not do as good a job forecasting consumption growth as households do, or there may be some bias in their reporting of expected consumption growth.9

Although we recognize that our survey measures expected consumption growth with error, the usual method of testing whether expected consumption growth is equal to the ex ante real interest rate is also subject to errors. The standard methodology is to assume expectations are rational, and then to use ex post consumption growth to measure ex ante consumption growth, with an error that is uncorrelated with the information used to formulate expectations. If expectations are rational, that approach should work if the data sample is sufficiently large. However, if consumption growth is persistent and/or the data generating process for consumption growth is subject to periodic regime changes, then ex post measurement may not yield a very useful measurement of ex ante expectations given available sample sizes.

Of course, the other criticism of the standard methodology is that it requires incorporating rational expectations as part of the null hypothesis. Survey-based measures of expectations place no such restriction. Moreover, there is a middle ground where survey measures still may be advantageous. If income is subject to regime changes, households may not be immediately aware of the new data generating process. There may be a period of learning. Even when learning is optimal, the problems with using a small data sample of ex post consumption to measure ex ante expectations are compounded.

Another shortcoming of our approach arises from the fact that the Euler equation does not exactly relate expected consumption growth to real interest rates, even when households have a power utility function. Instead, it equates the expectation of a function of consumption growth to the real interest rate. We are using a first-order approximation to the Euler equation, but there may be some information lost that a higher-order approximation would capture. It seems unlikely, however, that inclusion of these higher moments could account for the high correlation of expected consumption growth with expected income growth.

Our approach does not allow for heterogeneity among agents. Perhaps households have different information sets, and therefore have different

9There is of course a long literature evaluating various properties of survey responses of professional forecasters. Much of this work has evaluated the survey forecasts of inflation. In two recent such contributions, Ang, Bekar, and Wei (2007) and Croushore (2007) present quite favorable evidence on the properties of the survey forecasts. These recent papers argue that earlier papers that had found strong evidence against the rationality of survey forecasts were ignoring real-time issues, putting too much weight on the time around 1980 when there was learning about structural breaks, and not doing the econometrics of relationships among highly persistent variables correctly. The evidence on post-1985 data that is more careful on econometrics and real-time issues does not find a whole lot of evidence of substantive deviations from forecast efficiency. As noted, this work has all focused on inflation forecasts; we are not aware of a comparable literature evaluating the properties of the long-horizon forecasts of consumption that we use in this paper.
expectations of consumption. Indeed, our measure of expected consumption averages forecasts over the reported forecasts from several dozen professionals. Smith and Yetman (2007) have examined the properties of forecasts of consumption and real interest rates (at horizons up to one year), using data drawn from individual forecasters. They use a panel approach to estimating the parameters of the Euler equation based on individual forecasts, and do not attempt to relate consumption forecasts to income forecasts. However, their findings are not encouraging for Euler equations, as they find overall a negative relationship between expected consumption growth and ex ante real interest rates.

There may be heterogeneity of another sort, arising from the fact that individual lifetimes are finite. Models with overlapping generations of agents with imperfect bequest motives have been used in recent literature to explore the behavior of current accounts.¹⁰ Further exploration may determine if such models could account for either our empirical findings or those of Smith and Yetman (2007).

A possible explanation for our finding that expected consumption growth is not closely related to ex ante real interest rates is that we have done a poor job modeling preferences. Although the class of models that relate ex ante consumption growth to ex ante real interest rates, to a first-order approximation, extends beyond the simple power utility function, this still represents a potentially narrow assumption about utility. We reiterate, however, that this relationship is an essential element of many macroeconomic models. But allowing for a general specification of preferences may break down the relationship between expected consumption growth and the real interest rate. One possible elaboration of preferences is simply to allow nonseparability between leisure and consumption. Head, Mattina, and Smith (2004) consider this as a possible explanation for the Backus-Smith puzzle, but in fact find that this nonseparability is of little help. Recently, Jaimovich and Rebelo (2007) have proposed a specification in which consumption and leisure are nonseparable, to help explain the macroeconomic reaction to news in an open economy.

Cochrane (2005) has, in essence, advocated an approach in which minimal assumptions are imposed on preferences. Under his approach, we might do one of two things. We could assume the Euler equation is true, and use the Euler equation to back out implications about preferences. Campbell and Mankiw (1999) is an example of this approach. Alternatively, we could examine aspects of Euler equations that are not dependent on preferences, such as examining the consistency of Euler equations for pricing different assets. Brandt, Cochrane, and Santa-Clara (2006) is an application of this approach to international asset prices.

The challenge, however, is to reconcile this approach with our finding that expected consumption growth is so closely related to expected income growth. Are there plausible preference specifications that can deliver this result?

Now we come full circle. If indeed there are significant impediments to international capital flows, aggregate consumption growth might look like it

¹⁰Recent examples include Caballero, Farhi, and Gourinchas (2008) and Ferrero (2007).
would in a closed economy. Specifically, in a closed economy, consumption growth must equal income growth (or, at least, income net of investment and government spending.) One possible reaction to our findings then is the following: Expected consumption growth matches expected income growth because opportunities for international asset trade are limited. Expected consumption growth is not closely determined by the ex ante real interest rate because there is a time-varying pricing kernel that we do not capture well in our model with power utility. (That is, à la Cochrane, the Euler equation does not put a directly measurable restriction on expected consumption growth.)

In the end, it is an open question whether this explanation for our findings is more plausible than the alternative that credit constraints are important so that aggregate consumption behavior is not well described by Euler equations. Or perhaps there are elements of truth to both stories. In either case, the empirical findings present a challenge to the building block of many macroeconomic models, in which expected consumption growth is driven by the ex ante real interest rate.

APPENDIX

Figure A1. Expected Personal Consumption Growth

Note: As in Figure 1, with one-year expectations.
Figure A2. Expected Personal Consumption Growth and Expected GDP Growth

Note: As in Figure 2, with one-year expectations.
Figure A3. Expected Personal Consumption Growth and Expected Real Interest Rate

Note: As in Figure 3, with one-year expectations.
Figure A4. Coefficients on Expected GDP Growth and Expected Real Interest Rate
(Specification with natural log of variables)
Figure A4 (concluded)

Note: As in Figure 6, with one-year expectations.

REFERENCES


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Global Dispersion of Current Accounts: Is the Universe Expanding?

HAMID FARUQEE and JAEWOO LEE

This paper examines the global distribution of current accounts. Using a panel of more than 100 countries, the analysis establishes a set of stylized facts regarding the collective behavior of current accounts over the past four decades. In particular, we find that the global dispersion of current accounts has been steadily rising, which is qualitatively consistent with the view that ongoing financial globalization has allowed countries to maintain larger current account imbalances. However, this underlying trend is not quantitatively large enough to explain “global imbalances”—that is, the noticeable widening in external imbalances among major economies (for example, United States) seen in recent years. [JEL F3, F4]


Few external variables are as closely or widely watched as the current account balance. Rightly or wrongly, it has been used as a barometer for a wide range of economic conditions—from the state of the business cycle to the sustainability of external financing. In recent years, attention to current account imbalances has taken on a global dimension, reflecting concern over

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“global imbalances.” At the center is a large (albeit moderating) current account deficit in the United States, reflecting a shortfall of domestic saving relative to investment on the order of 5 to 6 percent of GDP. As a counterpart, large and/or growing current account surpluses have been recorded in Japan, China and other emerging Asia, and oil-exporting countries most recently, but less so in Europe (excluding Russia). Whether such a global constellation of widening external imbalances can be sustained and for how long constitutes a key macroeconomic risk facing the world economy.\(^1\) Namely, the possibility of a hard landing in the U.S. dollar—the international currency of choice—has raised concerns in many parts of the world over the potential fallout from a disorderly global rebalancing.

A countervailing argument to such concerns was perhaps most notably voiced by former Federal Reserve Chairman Alan Greenspan. Turning matters around, he argued that the unprecedented size of the U.S. deficit was itself a testimony to the increasingly efficient functioning of international capital markets and its ability to mobilize such a large share of net saving from the rest of the world to the United States. Specifically, he noted the following stylized fact regarding global trade and capital flows:

> Uptrends in the ratios of external liabilities or assets to trade, and therefore to GDP, can be shown to have been associated with a widening dispersion in countries’ ratios of trade and current account balances to GDP. A measure of that dispersion, the sum of the absolute values of the current account balances...has been rising as a ratio to GDP at an average annual rate of about 2 percent since 1970 for the OECD countries, which constitute four-fifths of world GDP... More generally, the vast savings transfer has occurred without measurable disruption to the balance of international finance... Accordingly, the trend... will likely continue as globalization proceeds.\(^2\)

This paper reexamines the global distribution of current accounts viewed from a longer term perspective. Using a panel of over 100 countries that comprise over 95 percent of world output, the analysis establishes a set of “stylized facts” regarding the individual and collective behavior of current accounts over the past four decades. In particular, we examine the dispersion properties of external imbalances and interpret these empirical regularities in the context of increasing openness in trade and financial flows—often referred to as “globalization.” While an emergent literature on financial globalization has documented that gross financial flows (including international reserve accumulation) has increased dramatically in recent

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\(^1\)See IMF (2005), Blanchard and Sa (2005), Chinn and Lee (2009), Faruqee and others (2006a,b), and Gourinchas and Rey (2005).

\(^2\)Remarks at the 21st annual monetary conference at the Cato Institute, November 2003.
years, what does this imply (if anything) for net flows? More specifically, the central issues that the paper addresses include the following:

- Is the universe of current accounts expanding or narrowing? And, at what rate? What fraction of the U.S. external deficit specifically (and global imbalances broadly) can be attributed to the underlying changes in global dispersion?
- What does changing global dispersion imply for current account persistence?
- What economic factors help explain underlying trends in the dispersion of external imbalances?

Besides risk and policy implications, the question of rising dispersion has a direct bearing on the celebrated Feldstein and Horioka (1980) puzzle. Their basic finding that savings are closely correlated with investments across countries has remained more or less intact, despite several prominent exceptions (for example, Blanchard and Giavazzi (2002) for Europe). Our query on rising dispersion would help to answer whether the background for the Feldstein-Horioka findings remains intact. If there is no trend change in the dispersion of current accounts, Feldstein-Horioka correlations should continue to be confirmed in the data with statistical significance as strong as the original results. If instead a rising trend is identified in the dispersion of current accounts, it would suggest that these findings would likely weaken over time, though not necessarily becoming extinct.

I. Dispersion and Convergence

We first ask if the global constellation of current account imbalances has been expanding or narrowing over time. Conceptually, in the case of convergence, there is a universally unique end point—that is, zero balance—around which all current accounts should converge (up to a discrepancy term). But predictions from economic theory are generally ambiguous on this and whether external balances should gravitate toward some long-run equilibrium or even be path-dependent based on the history of shocks. The answer typically depends on the class of model—for example, representative agent vs. overlapping generations framework—and its assumptions regarding market completeness, initial conditions and the history of shocks. Hence, whether current accounts actually converge or diverge and over what horizon are essentially empirical questions. To examine these issues more closely, we employ both nonparametric and parametric methods—including concepts from the growth literature on convergence—to determine if the universe of current accounts is expanding.

3For recent studies on financial globalization, see Prasad and others (2003), Kose and others (2006), and the references cited therein.
Unconditional Distributions

The unconditional distribution of current account ratios (in percent of GDP) at different points in time is shown in Figure 1. Kernel density estimates of the cross-sectional distribution for 101 countries suggest that the universe of current accounts has been generally expanding. The distribution of current accounts shows a steady increase in dispersion from 1960 to 2004, with the mass of the distribution being less concentrated in the area around zero and moving further out toward the tails. This suggests that, on a global basis, current account imbalances or the net flows of goods and services have tended to rise in proportion to the overall economy over time. The notable exception to this progressive pattern of expansion is the distribution of the ratio of the current account to trade (imports and exports) which has remained stable over the years.

\[ f(x_i) = \frac{1}{Nh} \sum_{j=1}^{N} K \left( \frac{x_i - x_j}{h} \right) \]

where \( x_j \) is the \( j \)th data observation, \( N \) is the number of observations, \( h \) is the window size (that is, the degree of smoothing), and \( K \) is the kernel or weighting function. The non-parametric estimates in Figure 1 are based on the Epanechnikov kernel (see Silverman, 1986). Results using the less efficient Gaussian kernel (that is, standard normal) are very similar.

Jarque-Bera tests strongly reject normality for each of these years. Skewness in the distribution was found significant for each of these years, except 1960; excess kurtosis (that is, "fat tails") was statistically significant throughout.

In contrast to the rising dispersion in the distribution of the current account as a share of GDP, the distribution of the ratio of the current account to trade (imports and exports) has remained stable over the years. In other words, current account positions (largely net trade
year 1980, when presumably the effects of oil shocks widened the dispersion of current accounts temporarily beyond that seen in later years. Notice too that the distributions for each year are not exactly centered around zero (but a small negative value), consistent with the global current account discrepancy.\(^7\)

\(\sigma\)-Convergence

Figure 2 presents supporting evidence from a time-series perspective. The figure plots annually two dispersion measures of current account ratios over the past 45 years. They are the global standard deviation of current accounts (in percent of GDP) or \(\sigma\), and the global mean absolute deviation of the current account (in percent of GDP) or \(\mu\), both calculated across countries for each year.

Two features of the figure are worth noting. First, dispersion shows significant time variation from year to year. Consistent with the impression from the previous figure, notice the considerable increase in the global spread around the time of the two major oil shocks in the mid- and late 1970s. Second, an underlying trend increase in dispersion is apparent, consistent with the global distributions shown in Figure 1. Specifically, the universe of current account positions has been expanding on average by almost 2 percent per annum, measured by its standard deviation.\(^8\) This latter finding suggests a lack of so-called \(\sigma\)-convergence in external positions over this time horizon.

Note that both measures are unweighted, treating each country symmetrically. For comparison, a third measure of dispersion \(\sum\) is shown in the figure by computing the global sum of current accounts (in absolute value) in percent of world GDP. This is equivalent to a weighted mean absolute deviation of current accounts (in dollar terms), where country weights are determined by own GDP (in dollar terms) as a share of world GDP (in dollar terms). All three measures are highly correlated (with correlation coefficients between 0.60 and 0.95). However, the third measure shows a steeper increase, particularly in recent years, well ahead of the other (unweighted) measures.\(^9\) This corresponds to emergence of “global imbalances” where large deficits and surpluses emerged in the largest countries such as the United States and China.

\(^7\)The global current account discrepancy—usually expressed in dollar terms or in percent of world imports or GDP—has mostly been negative since the early 1970s, reflecting discrepancies in both trade and income accounts; see Marquez and Workman (2001) for a discussion.

\(^8\)A regression of the (log) standard deviation on a time trend yields the following results (with corrected standard errors given in parentheses): \(\ln(\sigma_{CA/GDP}) = 1.4 + 0.017t + \epsilon_t; R^2 = 0.33.\)

\(^9\)Over the sample, the rate of increasing global dispersion is \(1 \frac{1}{2}\) to \(1 \frac{3}{4}\) percent per year on an unweighted basis and \(3 \frac{1}{4}\) percent on a weighted basis.
Another convergence perspective—commonly used in the growth literature—is the notion of “β-convergence.” In the context of current accounts, β-convergence would require that countries accumulating past imbalances eventually unwind these positions. This would allow current accounts (and trade balances) globally to converge to more similar values around zero—that is, the convergence point. For example, countries with a large stock of net external debt, reflecting flow deficits in the past, would need to run current account surpluses in the future to pay down the debt or, at least, smaller current account deficits to decrease the share of debt relative to the overall economy. Comparing the initial net foreign asset ratio to the average current account ratio in subsequent years, however, provides very little support for this type of convergence. The cross-country regressions, in fact, show that countries with larger net indebtedness are more likely to run larger (not smaller) current account deficits in the years that follow (Figure 3).

The results of the cross-section regression are:

\[
CA_{avg} = -1.94 + 0.09 \times NFA_0 + v; \quad R^2 = 0.13; \quad NOBS = 94.
\]

Moreover, the slope of the line drawn is greater than typical estimates of nominal growth in GDP, suggesting these subsequent flow imbalances tend

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10A weaker form of convergence posits that current accounts, but not trade balances nor net foreign asset positions, converge toward balance. This would require that trade (not current account) surpluses be achieved in the years following current account deficits to stabilize the accumulation of net foreign liabilities. Net foreign assets data are based on Lane and Milesi-Ferretti (2001, 2007).

11Reflecting limitations on historical data for net foreign assets, the cross-section consists of 94 countries. Similar positive-sloped regression results follow when the sample is split into surplus and deficit countries, as well as between high vs. low deficit countries. Chinn and Prasad (2003) find a similar effect of net foreign assets on current accounts based on multivariate panel estimation that controls for a wide range of explanatory variables (for example, demographics, fiscal positions, economic development, and so on).
to augment the net stock of external assets or liabilities in relation to the size of the economy. Alternative coefficient estimates may be more comparable to nominal dollar growth rates. This would imply a reversion to the initial ratios of net foreign assets (NFA) to GDP, but this is still at variance with the notion of convergence to a common value (for example, zero balance).12

This result is comparable to the findings of Kraay and Ventura (2000). Using the data from less than 20 industrial countries, they found that current account imbalances are proportional to the net external balance sheet positions. In response to an increase in savings, a creditor country tended to run surplus while a debtor country tended to stay in deficit. They view this to be the result of a portfolio choice in the presence of a large investment risk. While data limitation makes it difficult to examine the validity of their prediction for a wider set of countries, their model is one possible explanation for the result that we find for a very large set of countries.

To recap, the distributional and convergence properties of current account balances suggest an expanding universe. The β-convergence results further suggest that countries who have had current account imbalances historically are the group more likely to run subsequent current account imbalances (of the same sign) in ensuing periods, leading to further accumulation of net foreign assets or liabilities.

12Dropping outlier countries with average current account imbalances (net external assets) greater than 10 percent (50 percent) of GDP in absolute terms would slightly lower the coefficient on initial NFA (to 0.06) but raise its significance level (p-value = 3 percent).
II. Stationarity and Persistence

We now examine aspects of the time-series properties of current accounts—in particular, stationarity and persistence. Trehan and Walsh (1991) showed that the stationarity of the current account is a sufficient condition for the intertemporal budget constraint to hold. Stationarity has since been an indirect test of the basic premise of the intertemporal view of the current account. Thus, this type of behavior would indicate whether the expanding global dispersion of current accounts has also been compatible with respecting intertemporal budget constraints.

To examine the stationarity and persistence properties of current accounts, a battery of unit root and stationarity tests were conducted. In particular, the well-known augmented Dickey-Fuller (ADF) test and nonparametric Phillips-Perron (PP) test for a unit root against a stationary alternative were applied to the individual country series for the current account ratio (in percent of GDP). In addition, the Kwiatkowski and others (1992) (KPSS) test for stationarity against a unit root was also used. The corresponding test statistics and significance levels are shown in Appendix II.

Figure 4 summarizes the rejection and nonrejection rates (in percent) for these unit root tests. For more straightforward comparisons, the rejection of stationarity under the KPSS test is reported as a nonrejection of the unit root. Overall, the picture is quite mixed. One test finds the majority of current accounts to be nonstationary (ADF test), another tests finds the majority to be stationary (KPSS test), and the third test is split down the middle (PP test).

Individually, for nearly a quarter of the sample (22 of 101), these tests failed to reject both nonstationarity and stationarity for the same series (see Appendix II).

This finding highlights two widely known features of these tests and the time-series data: (1) unit root and stationarity tests tend to have low power (that is, fail to reject too often) in finite samples, and (2) the current account is generally a very persistent series, making it difficult to distinguish between nonstationary and stationary alternatives over limited time spans.

For 21 countries—including, notably, the United States and Japan, the tests indicated (at least, statistically) a nonstationary current account ratio over this time span. That is, the unit root tests failed to reject nonstationarity and the KPSS test further rejected stationarity. But for more than half of the sample (55 of 101 countries), at least one of the two unit root tests rejects and

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13 Trehan and Walsh showed that the stationarity of the current account was the necessary and sufficient condition, but the necessity was debated lately by Bohn (2005).
14 In all cases, the model specification includes a constant but no time trend. Including a time trend in the unit root tests marginally increase the number of rejections.
15 See Campbell and Perron (1991). A peculiar finding is that for three countries, these low-powered tests rejected both stationarity and nonstationarity for at least one unit root test.
the stationarity test accepts their respective null hypotheses, suggesting a stationary series.

Moreover, on the basis of panel unit root tests (Table 1), nonstationarity is strongly rejected. The tests were applied to three panels comprising different groups of countries according to data availability. The first panel comprises data from 1960 to 2004 for 49 countries, the second panel comprises data from 1975 to 2004 for 77 countries, and the third panel comprises data from 1985 to 2004 for 101 countries (excluding Kuwait). The null of nonstationarity is strongly rejected for all possible specifications suggested by Levin, Lin, and Chu (2002), Im, Pesaran, and Shin (2003), and Breitung (2000). Test statistics reported in the table correspond to specifications without time trend, but a unit root was rejected for specifications with time trend, too.
Overall, these various tests suggest that the current account is a stationary but persistent series. When a simple AR(1) specification is estimated from a panel perspective, pooled ordinary least squares and fixed-effects estimates, respectively, yield the following equations (with standard errors in parentheses):

\[ CA_{it} = k + 0.75 CA_{i,t-1} + \varepsilon_{it}; \quad R^2 = 0.59, \]  
\[ (0.02) \]

\[ CA_{it} = k + 0.60 CA_{i,t-1} + \varepsilon_{it}; \quad R^2 = 0.62. \]  
\[ (0.01) \]

Under either specification, there is significant AR(1) coefficient on the lagged current account, though with panel fixed effects, the degree of inertia is reduced somewhat.

But these specifications are, in a sense, incomplete—failing to recognize a common component associated with the particular pattern in the movement of global dispersion over the past several decades. Moreover, the β-convergence results indicate that countries with nonzero initial NFA positions continue to accumulate assets (liabilities) on a net basis by running current account surpluses (deficits) in subsequent periods. In other words, countries tend to run significant imbalances of the same sign (either positive or negative) as in the past. To introduce this trend feature into the analysis, we include a sign-preserving time trend (sptrend) constructed as follows:

\[ sptrend_t = \text{sign}(CA_{t-1}) \times t; \quad \text{where} \quad \text{sign}(CA_{t-1}) = CA_{t-1}/|CA_{t-1}|. \]  
\[ (3) \]

The time trend specifies increasing surpluses or deficits depending on the sign of the current account in the previous period. Note too that this sign-preserving trend is also broadly consistent with preserving current account adding-up, but a simple time trend is not.\(^{16}\) Including this term into the panel fixed effects regression yields:

\[ CA_{it} = k_i + 0.01 \text{sptrend}_t + 0.59 CA_{i,t-1} + \varepsilon_{it}; \quad R^2 = 0.63. \]  
\[ (0.004) \quad (0.01) \]

The trend capturing increasing dispersion is statistically significant (\( p \text{ value} = 0.08 \)).\(^{17}\) The fit of the equation is marginally improved and the persistence parameter (AR1 coefficient) is smaller, as one would expect.\(^{18}\) In other words, some of the observed persistence in external

\(^{16}\)Current account adding-up would be more apparent if balances were defined in a common unit (say) U.S. dollars—but this would raise issues of nominal drift. Using the current account ratio to GDP broadly preserves the level of the current account discrepancy (in percent of GDP) provided that surplus and deficit countries (as respective groups) are of similar economic size.

\(^{17}\)Owing to the unbalanced panel, the sign-preserving trend coefficient in (4) is estimated for the vast majority (but not all) countries. Including the full sample (which has countries with only few observations at the end of sample) would reduce the coefficient estimate given the accumulated value of the trend term itself. A “resetting” trend for these countries would raise the point estimate. For further sensitivity analysis, see the following footnote.
balances appears to reflect an underlying trend phenomenon—a slowly increasing global dispersion. This is perhaps better viewed as an evolving longer-run process occurring over many decades rather than the inertia in external balances seen from year to year.

Taking an average trend estimate over different specifications, one can examine the extent to which the recent increase in the U.S. current account deficit is due to this underlying global trend. Figure 5 shows the observed U.S. current account deficit ratio to GDP (line) and the long-run contributions (bars) that obtain from equation (4), reflecting the common or global trend term on average. Accounting for increasing global dispersion goes part way to explaining the burgeoning U.S. deficit in past years, but clearly the widening imbalance has gone far beyond these considerations. In level terms, the long-run component (adding the trend and either a common or country-specific constant) would narrow but not nearly close the “gap” with the observed deficit.

Figure 5. U.S. Current Account (In percent of GDP)

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18When the sign-preserving trend is constructed using the contemporaneous rather than lagged current account, the trend coefficient is always substantially larger and highly significant; and the AR1 coefficient is also substantially reduced. But this specification is susceptible to simultaneity issues. Across various specifications and samples, the range of trend estimates is roughly between 0.005 to 0.008. As a cross-check, the fitted values and implied rate of increase in global dispersion for a given trend estimate are compared to the rates from the nonparametric estimates discussed earlier.
III. Dispersion and Openness

We have seen that the cross-section dispersion of current accounts has been rising but the time series of current accounts have remained stationary. In particular, equation (4) based on sign-preserving trends suggests that the cross-section distribution of long-run average current accounts (measured by the constant terms in AR(1) regression and the sign-preserving trends) has been spreading out. What are behind these trends? Are they in fact driven by the force of globalization, experienced as a rising integration of goods and financial markets across different countries? We look into these factors empirically and theoretically in this section.

Some Evidence

We first consider two economic variables that are likely to affect the behavior of external imbalances: openness to trade and financial flows. We measure trade openness by the ratio of exports and imports to GDP, and financial openness by the Chinn-Ito index, scaled to lie between 0 and 1 (see Appendix I for details). As most economies have been opening up to rising external flows in trade and finance, a deterministic time trend will capture a large part of the common trend toward greater openness. Countriespecific measures of openness will help us to extract more information on the role of openness by exploiting different speeds toward openness among countries.

Countries with more open trade regimes would find it easier to sell goods produced beyond the need for domestic use and to import goods for which demand exceeds domestic production. The trade imbalance is the aggregate accumulation of such imbalances over the whole economy. A country with more open trade regime will thus be more likely to run a trade imbalance, and also find it easier to finance them given the wider base for international lending and borrowing.

More directly, countries with more open financial account will find it easier to lend or borrow to balance its savings capacity and investment need. In addition to enabling countries to put savings to the most productive use and to finance investment needs in the most efficient manner, a greater availability of investment and funding opportunities will tend to stimulate savings and investment, and increase international financial flows further.

To develop a framework on how to incorporate these effects on the dispersion of current accounts, we consider the following AR(1) representation of the current account of country $i$.

$$CA_{it} = \mu_i + \beta CA_{i,t-1} + \epsilon_{it}.$$ 

Idiosyncratic shocks $\epsilon_{it}$ are uncorrelated across countries and time ($i$ and $t$), and have mean zero and unit variance ($\sigma^2(\epsilon_{it}) = 1$), evaluated over $i$ at each point in time. The long-run average current account, $\mu_i/(1-\beta)$, is allowed to differ across countries. Because current accounts have been found
to be stationary ($0 < \beta < 1$), we obtain the following moving-average (MA) representation of current accounts.

$$CA_{it} = \sum_{l=0}^{\infty} \beta^l (e_{it-l} + \mu_i).$$

The global dispersion of current accounts at time $t$ is

$$\sigma^2(CA_{it}) = \sum_{l=0}^{\infty} \beta^{2l} \sigma^2(e_{it-l}) + \sigma^2(\mu_i))$$

$$= (1 - \beta^{2})^{-1}[1 + \sigma^2(\mu_i)].$$

There are three ways that the global dispersion of current accounts can increase, which need not be mutually exclusive. For one, the cross-section distribution of the long-run average current account can spread out over time—reflecting an increasing ability to borrow and lend internationally, thereby increasing the global dispersion of the current account. Alternatively, a rise in the underlying persistence of a current account deviation from its long-run average can increase the observed global dispersion of current account as the effect of idiosyncratic shocks die out more slowly. This latter channel, however, is not pursued in this paper, for it appears to require a much longer span of data than used in this paper. (See Taylor, 2002) for suggestive evidence of this channel in traditional Organization for Economic Cooperation and Development countries over a span of 100 years.) Finally, the variance of idiosyncratic shocks, which has been normalized to unity here, could actually be rising over time. Indeed, Kose and others (2008) find evidence that the share of idiosyncratic shocks in global business cycles was a little larger in the 1985–2005 period than in the 1960–84 period. However, the change is quantitatively too small to account for a large part of the rising dispersion in current accounts. Nor is it possible to estimate such decompositions for each year separately, and this channel is not pursued further in this paper either.

To capture the spreading-out of the cross-section distribution, we construct sign-preserving indicators of openness. Using $\text{tradeopen}$ and $\text{finopen}$ to measure openness in trade and financial accounts, respectively, sign-preserving openness indicators are:

$$sp\text{tradeopen}_t = \text{sign}(CA_{i-1}) \times \text{tradeopen}_t,$$

$$sp\text{finopen}_t = \text{sign}(CA_{i-1}) \times \text{finopen}_t.$$

This leads to an expanded version of equation (4).

$$CA_{it} = \mu_i + \beta_1 CA_{it-1} + \beta_2 sp\text{trend}_t + \beta_3 sp\text{tradeopen}_{it} + \beta_4 sp\text{finopen}_{it} + \nu_{it}. \quad (5)$$

Table 2 reports the results of estimation for 77 countries whose current account data are available starting in no later than 1975. In the upper panel,
The openness was measured by the average over the whole sample period, thus using current accounts since 1960 for countries with available data. In the lower panel, the openness in trade and finance was measured as three-year moving averages, thereby restricting the sample to 1970 onwards (Chinn-Ito index available from 1970). In both upper and lower panels, the two left columns report estimates obtained from using the sample of all 77 countries, the two middle columns report estimates from the sample of 21 advanced countries, and the two right columns report estimates from the remaining (emerging-market and developing) countries.

The estimates indicate that financial opening (integration) played a large role behind the sign-preserving trend that was documented in the previous section. In estimations (1) and (2) of the upper panel where the average measures of openness are used, the sign-preserving trend loses statistical significance when the sign-preserving financial openness term is included.

### Table 2. Current Account Dispersion and Openness

<table>
<thead>
<tr>
<th></th>
<th>All countries</th>
<th>Advanced countries</th>
<th>Other countries</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
</tr>
<tr>
<td>CA(−1)</td>
<td>0.58**</td>
<td>0.57**</td>
<td>0.69**</td>
</tr>
<tr>
<td>(0.02)</td>
<td>(0.02)</td>
<td>(0.03)</td>
<td>(0.03)</td>
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<tr>
<td>SPTrend</td>
<td>0.006*</td>
<td>−0.001</td>
<td>0.014**</td>
</tr>
<tr>
<td>(0.004)</td>
<td>(0.005)</td>
<td>(0.006)</td>
<td>(0.006)</td>
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<tr>
<td>SPTrOpenAvg</td>
<td>−0.357</td>
<td>(0.231)</td>
<td>−0.300</td>
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<tr>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>SPFinOpenAvg</td>
<td>0.947**</td>
<td>(0.277)</td>
<td>−0.130</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>R²</td>
<td>0.53</td>
<td>0.52</td>
<td>0.71</td>
</tr>
<tr>
<td>(7)</td>
<td>(8)</td>
<td>(9)</td>
<td>(10)</td>
</tr>
<tr>
<td>CA(−1)</td>
<td>0.58**</td>
<td>0.54**</td>
<td>0.68**</td>
</tr>
<tr>
<td>(0.02)</td>
<td>(0.02)</td>
<td>(0.03)</td>
<td>(0.04)</td>
</tr>
<tr>
<td>SPTrend</td>
<td>0.005</td>
<td>−0.019**</td>
<td>0.013**</td>
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<tr>
<td>(0.003)</td>
<td>(0.007)</td>
<td>(0.003)</td>
<td>(0.012)</td>
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<tr>
<td>SPTrOpenMav</td>
<td>0.054</td>
<td>(0.300)</td>
<td>0.594</td>
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<td></td>
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<tr>
<td>SPFinOpenMav</td>
<td>1.464</td>
<td>(0.315)**</td>
<td>0.893**</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>R²</td>
<td>0.53</td>
<td>0.51</td>
<td>0.73</td>
</tr>
</tbody>
</table>

Note: Statistically significant at 5 percent (**) and 10 percent (*). All regressions included country fixed effects. Based on 77 countries whose current account data were available starting before 1975, comprising 21 advanced economies and 56 other (emerging and developing) economies.
This evidence turns out to be a combined effect between advanced and other economies: in estimations (3)–(6), the sign-preserving trend remains statistically significant under the advanced-economy sample, but the sign-preserving financial openness is statistically significant only for other economies. The indirect support for the role of financial openness in advanced economies is probably due to the limited “cross-section” variation in financial openness among advanced economies. Nevertheless, a clear contrast can be drawn vis-à-vis the conspicuous absence of statistical evidence that trade openness played an important role behind the sign-preserving trend.

More convincing evidence on the role of financial openness is provided in the lower panel, where moving averages of openness is used, thereby allowing a greater variation in financial openness among advanced economies as well as among other economies. In addition to the decline in the effect of sign-preserving trend observed in estimations (7) and (8) for the whole sample, estimates (9) and (10) show a clear decline in the effect of sign-preserving trend, combined with a strong statistical significance of the sign-preserving financial openness. The same pattern is also observed in estimates (11) and (12) for other countries.

A Simple Model of Expanding Dispersion and Financial Integration

Building upon the evidence that financial openness appears to have played a pivotal role in expanding the global dispersion of current accounts, we present an illustrative (steady-state) model where the ongoing integration in international financial markets increases the global dispersion of current accounts. Some form of heterogeneity is a necessary condition for global dispersion, and we introduce a heterogeneity in discount rates. Combined with a small cost of financial intermediation, which represents financial market friction, we generate a nondegenerate steady-state distribution of net foreign assets. Further introducing growth in aggregate output, we show that the global dispersion of current accounts rises, as the cost of financial intermediation falls. And we assume a deterministic world.¹⁹

World economy is assumed to comprise \(N\) open economies, each of which commands a fixed stream of endowment \((y_i)\), and has an identical utility function \(u(c_{it})\) but with a heterogeneous discount rate \(\beta_i\). The heterogeneity in the discount rate gives rises to international lending and borrowing (thus current account imbalances). It is further assumed that countries incur a financial transaction cost when they lend or borrow. The period-by-period

¹⁹In a deterministic setup, the ratio of the current account to GDP has a well-defined relationship to the steady-state ratio of NFA to GDP. In a stochastic setting, stationary shocks to the current account (changes in NFA) will have a nonstationary effect on NFA, and the stochastic steady-state relationship between the current account and NFA remains unclear.
budget constraint is written as:

\[ a_{it+1} = (1 + r_t) a_{it} - c_{it} + y_{it} - \frac{\gamma}{2} (a_{it})^2 + \eta(a_{it}), \]

where \( a_{it} \) denotes the net foreign assets of country \( i \), and \( r_t \) the world interest rate. The financial transaction cost, \( \bar{g}(a_{it})^2 \), captures the cost needed to maintain a nonzero international asset position. The quadratic form is akin to the adjustment cost widely used in the macroeconomic literature, and is assumed to be the fee paid to competitive financial intermediaries when international financial position is adjusted away from the zero balance. The fees paid to competitive intermediaries are distributed back to each country (\( \eta(a_{it}) = \gamma/2(a_{it})^2 \)) ex post, and thus financial frictions affect the ex-ante decision making of each country without draining global resources.

The first-order condition for each country’s consumption-saving choice is:

\[
(b_i)^t u'(c_{it}) = (b_i)^{t+1} u'(c_{it+1})(1 + r_t - \gamma a_{it}).
\]

In the steady state, \( c_{it+1} = c_{it} = \bar{c}_t \), and equation (6) simplifies to \( 1/\beta_i = 1 + r - \gamma \bar{a}_i \) and the net foreign asset position is determined by the difference between the world interest rate and the subjective discount rate:

\[
\bar{a}_i = \frac{1}{\gamma} \left( 1 + r - \frac{1}{\beta_i} \right).
\]

The world interest rate is determined at a level that equates the global demand and supply of assets: \( \sum_{i=1}^{N} \bar{a}_i = 0 \):

\[
1 + r = \frac{1}{N} \sum_{i=1}^{N} \frac{1}{\beta_i}.
\]

In a no-growth economy just described, the steady-state distribution of net foreign assets can be maintained with zero balance in all current accounts, thereby generating no dispersion in current accounts. A steady-state dispersion in current accounts can be generated by introducing economic growth; now assume that each country’s population grows at the same rate \( g \). The aggregate output of a country at time \( t \) becomes: \( Y_{it} = (1 + g)^t \bar{y}_i \), normalizing the initial population (that is, at time \( t = 0 \)) at unity. Denoting the aggregate net foreign assets by a capitalized letter, \( A_{it} \), the change in the ratio of net foreign assets to GDP can be rewritten in terms of the current account as follows.

\[
\frac{A_{it+1}}{Y_{it+1}} - \frac{A_{it}}{Y_{it}} = \frac{A_{it+1} - A_{it}}{Y_{it+1}} [1 - (1 + g)] + \frac{A_{it+1} - A_{it}}{Y_{it}} = -g \frac{A_{it+1}}{Y_{it+1}} + \frac{CA_{it}}{Y_{it}}.
\]

\[\text{See Ghironi, Lee, and Rebucci (2007) for further discussion of these costs and their role in determining the steady-state values of the NFA-to-GDP ratio.}\]
In the steady state with a constant ratio of the aggregate net foreign assets to GDP,

\[
\left( \frac{CA_i}{Y_i} \right) = g \left( \frac{A_i}{Y_i} \right) = g \tilde{a}_i = g \left[ \frac{1}{N} \sum_{i=1}^{N} \frac{1}{\beta_i} - 1 \right],
\]

a tighter integration of international financial markets, represented as a decline in \( \gamma \) which lowers the cost of international financial transactions, increases the dispersion in the ratio of the current account to GDP.

**IV. Concluding Remarks**

Examining current accounts for a wide spectrum of countries over the past four and a half decades, we can summarize our key findings or “stylized facts” as follows:

- **The universe of current accounts has been expanding over the past half century.** Based on a variety of measures and methodologies, the global constellation of external current account positions has markedly widened over time. Although dispersion can vary significantly from year to year—ostensibly in response to large international shocks, there is a steady, underlying rate of expansion of around 2 to 3 percent per year.

- **In other words, in a context where global gross trade and financial flows have grown rapidly, net flows have also increased (on a sustained basis) to individual countries.** And sign reversals in the current account are occasional, but not frequent. Reflecting this persistence in current account imbalances, countries that have run larger external imbalances in the past also tend to run subsequent, larger imbalances (of the same sign), suggesting an extenuation of international lending or borrowing patterns. However, the presence of an underlying, long-run trend toward greater global dispersion suggests that inertia in current accounts from year to year may be overstated by simple estimates of persistence.

- **Rising dispersion is also found to be closely associated with increasing financial integration of the world economy.** Rising financial openness or integration appears to have played a large role, perhaps more so than trade openness, in accounting for the expanding universe of current accounts. At the same time, individual current account series and changes in net foreign assets (as ratios to GDP) are found to be stationary (albeit persistent), indicating that while dispersion is rising, basic intertemporal resource constraints are not likely violated for individual countries.

- **Global imbalances though have run well ahead of underlying dispersion trends.** The recent acceleration of external positions in major countries (including the United States) is clearly not fully accounted for by the trend behavior exhibited by the universal expansion.

From an economic standpoint, the results lend support to recent views that some, though not all, of the large global current account imbalances are due
to the ongoing integration of the world economy. In particular, it is not surprising that we would see, in an increasingly integrated global economy, higher levels of current account deficits (including in the United States) and surpluses in key partner countries. The other side of this trend is the likely weakening in the statistical hold of the Feldstein-Horioka results. However, we also find that the underlying pace of the increase in global dispersion is not as fast as sometimes claimed and has bounds, indicating that a sizable part of today’s global imbalances is likely in excess (relative to the underlying trend) and would probably be unwound to a significant degree. Some movements in that direction appear to have finally started in the United States, while the counterpart movements are less evenly distributed.

APPENDIX I

Data Description
The main variable is the ratio of the current account to the GDP, both of which were obtained from various issues of *International Financial Statistics* (IMF) and *World Development Indicators* (World Bank). The capital account liberalization index was developed by Chinn and Ito (2006), and is the first principal component of several variables that reflect the ease of cross-border financial transactions. In our estimation, the index was normalized to take a value between 0 and 1, increasing with the liberalization of capital account regime. For each value of Chinn-Ito index $CI_{it}$, our indicator is defined as follows:

$$finopen_{it} = \frac{CI_{it} - \min\{CI_{it}\}}{\max\{CI_{it}\} - \min\{CI_{it}\}}.$$

APPENDIX II

Alternative Measures of External Positions and Their Behavior
A related but distinct measure is the change in net foreign assets (NFA). It essentially differs from the current account by the amount of capital gains (valuation change), which is driven by asset price fluctuations, including exchange rate variations. Because these asset price movements are broadly described as a random walk, the change in NFA will contain a much larger white-noise component and exhibit smaller persistence than the current account. This is indeed confirmed by the data, as summarized in the following two charts. Note that due to data limitations regarding NFA, the sample size is smaller.

First, the change in NFA (in percent of GDP) is subjected to the same battery of stationarity and unit root tests as for the current account, summarized in Figure 4. The test results uniformly show a higher rejection rate of nonstationarity. See Figure A1. Second, for the change in the ratio of NFA to GDP—that is, $(NFA/y)$—the indications toward stationarity are even stronger; see Figure A2. Changes in the ratio also include a growth term (related to the change in the scaling variable GDP). This helps toward finding stationarity in the ratio given that GDP (that is, the denominator) is growing over time. Excluding the growth factor term (by considering $NFA/y$) weakens the stationarity finding, but does not overturn it. That is, the NFA concept appears to be much more stationary (less persistent) series than the current account.
Figure A1. Unit Root Tests Rejection Rates

(Variable = D[net foreign assets/GDP]; in percent)

1 Variable is change in net-foreign-assets-to-GDP ratio; for KPSS test, failure to reject stationarity reported as rejection of unit root in the figure. N = 81.

Figure A2. Unit Root Test Rejection Rates

(Variable = D[net foreign assets/GDP]; in percent)

1 Variable is change in net-foreign-assets-to-GDP ratio; for KPSS test, failure to reject stationarity reported as rejection of unit root in the figure. N = 81.


(Percent of GDP)

<table>
<thead>
<tr>
<th>Country</th>
<th>ADF 1</th>
<th>PP 2</th>
<th>KPSS 3</th>
<th>Country</th>
<th>ADF 1</th>
<th>PP 2</th>
<th>KPSS 3</th>
</tr>
</thead>
<tbody>
<tr>
<td>United States</td>
<td>0.21</td>
<td>−0.15</td>
<td>1.16*</td>
<td>Argentina</td>
<td>−2.24</td>
<td>−3.45*</td>
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Note: An asterisk * denotes statistical significance at the 5 percent level.

1. Augmented Dickey-Fuller t-test statistic for unit root against level stationary alternative; lag length chosen based on Schwarz BIC.
2. Phillips-Perron Z test statistic for unit root against level stationary alternative.
REFERENCES


How Long Can the Unsustainable U.S. Current Account Deficit Be Sustained?

CAROL C. BERTAUT, STEVEN B. KAMIN and CHARLES P. THOMAS*

This paper addresses three questions about prospects for the U.S. current account deficit. First, is it sustainable in the long term? Projections of a detailed model of the U.S. balance of payments suggest that the current account deficit will resume widening and external indebtedness will continue to expand. Second, how long will it take for indebtedness to rise sufficiently to prompt some pullback by global investors? We project that external debt, net investment income, and the share of U.S. claims in foreigners' portfolios will take many years to reach levels that would test global investors' willingness to extend financing. Finally, if and when levels of sustainable debt burden are breached, how readily would asset prices respond and the current account start to narrow? We find little evidence that, as countries’ indebtedness rises, the changes in asset prices and exchange rates needed to correct the current account materialize all that rapidly. [JEL F21, F32, F37]


Several years ago, as the U.S. current account deficit was expanding to record levels, observers increasingly began to focus on the unsustainability of the U.S. external imbalances, as well as the possibility

*The authors are economists in the International Finance Division of the Federal Reserve Board. This paper has benefitted from comments by our editor Akito Matsumoto, an anonymous reviewer, Trevor Reeve, and participants at the Current Account Sustainability in Major Economies (II) conference at the University of Wisconsin, especially discussant Jeffrey Frankel. Jim Albertus, Sean Fahle and Dao Nguyen provided excellent research assistance.
that the subsequent correction would be abrupt and disorderly. After peaking at 6.6 percent of GDP in the third quarter of 2006, however, the current account deficit has shrunk to about 5 percent in 2008 as a result of declines in the dollar, slower U.S. GDP growth, and expansion abroad. Concerns about a disorderly correction now appear to have become less prominent. This may in part reflect a growing conviction that a correction of the current account is likely to be orderly rather than disruptive. It may also reflect a view that, with the real multilateral dollar about 25 percent below its 2002 peak and the deficit reduced, no further correction of the U.S. current account may be necessary.

This paper addresses three simple questions: Is the U.S. current account now sustainable on a long-term basis? If not, how long might it take for measures of U.S. external indebtedness to expand beyond levels that global investors are willing to finance? And finally, if and when such levels are breached, how rapidly might a correction in asset prices and the current account ensue?

The first section starts by discussing the most common metric for assessing current account sustainability, the stability of an economy’s net debt as a share of GDP. Section II describes projections of U.S. external balance variables, based on simulations of a detailed model of the U.S. balance of payments, and measures them against the sustainability criterion described above. We find that, assuming the real value of the dollar remains flat at current levels, beyond the near term the current account deficit likely will begin widening again and U.S. external debt will rise steadily.

However, just because the current account is unsustainable in the long term does not mean that a correction is imminent. Theory provides no guidance as to how large the external debt must become before developments are triggered that would narrow the U.S. current account deficit. Our baseline projection suggests that the U.S. net external debt will grow from around 20 percent of GDP at present to around 60 percent of GDP by 2020. Is that a lot or a little? To answer this question, we look to the current pattern of external liabilities among industrial countries, and we find a number of countries whose external debt ratios currently are 60 percent or higher.

The net debt/GDP ratio, however, is not a perfect or unique indicator of the strength of a country’s international balance sheet. Accordingly, Section III considers a second metric of current account sustainability: the exposure of investors to U.S. assets in terms of the share of U.S. securities in foreign portfolios. If the financing of the current account deficit means that this exposure is rising without limit, that too would suggest that the current account is unsustainable. We find no evidence that, to date, the exposure of foreign investors to U.S. assets has been rising. Calculations based on our projections of the U.S. balance of payments (described above) suggest that

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1See, among others, Mann (1999 and 2002 and 2003) and Mussa (2004).
this exposure will increase going forward, but not necessarily to a worrisome extent.

Finally, in the event that measures of U.S. external indebtedness were to breach any significant thresholds of creditworthiness and investor exposure, do we know how rapidly global investors might react by pulling back from U.S. assets, thereby engendering a correction of the current account deficit? To address this question, Section IV examines the extent to which, historically, higher levels of external debt or related imbalances have led to changes in an economy’s access to financing. Estimating panel regressions for a sample of industrial countries, we find that higher levels of debt push up interest rates in a country by only a very small extent, and exert no discernable effect on exchange rates.

Section V summarizes the findings and advances some tentative conclusions.

I. Assessing Current Account Sustainability

The NIIP/GDP Criterion

The net international investment position (NIIP) represents the sum of all claims by U.S. residents on foreign residents less the claims of foreigners on the United States. The NIIP is a key determinant (along with rates of return) of U.S. net investment income: the sum of receipts on foreign assets owned by U.S residents net of payments on foreign claims on U.S. residents. Therefore, most analysts underscore that a necessary condition for current account sustainability is that the NIIP/GDP ratio be stable (Mann, 1999 and 2002 and 2003; Mussa, 2004; Cline, 2005). Otherwise, if the (negative) NIIP/GDP ratio were to rise without limit, the ratio of net investment payments to GDP would rise as well, and would eventually exceed GDP.

Qualifications to the NIIP as a Measure of External Sustainability

The NIIP does not fully summarize the sustainability of the external position. First, as will be discussed further below, depending upon the rate of return, the same NIIP may be associated with very different net investment income flows.

Second, changes in the valuation of assets may affect the NIIP without affecting the economy’s underlying capacity to service the external position. For example, all else equal, a rise in U.S. stock prices will raise the value of foreign holdings of U.S. assets and thus cause the NIIP to become more negative, even though U.S. residents have become wealthier and better able to make payments to foreigners. An extreme example is Finland, where a substantial share of its external liabilities consists of foreign holdings of stock in Nokia. In 1999, the NIIP surged to nearly −170 percent of GDP in 1999, driven by a parallel surge in the price of Nokia stock; when that stock declined, so, too, did the size of Finland’s negative NIIP.
Third, and as a related point, one should not confuse the NIIP with the net wealth of an economy’s residents. Net wealth is comprised of total assets owned by residents, both domestic and foreign, less foreigners’ claims on those residents. This amounted to over $55 trillion in 2006, dwarfing the size of both gross claims of foreigners on the United States ($14.4 trillion) and the NIIP ($2.2 trillion). Presumably, a country’s ability to repay external debt will depend not only on its GDP, but also on its total net wealth, just as a homeowner’s ability to repay his mortgage depends not just on his income, but on the value of his assets including, but not limited to, his house.

Finally, the aggregate NIIP may be only loosely related to the ability of U.S. residents to repay their external liabilities. For most international borrowers in an industrial economy, foreigners represent only a small portion of their creditor base, with most liabilities being to domestic residents. The diversity and distribution of creditworthiness across borrowers in a given economy is likely to influence the credit risks faced by foreign investors to a greater degree than the overall foreign indebtedness of the economy. (As discussed in more detail in Section II, the riskiness of U.S. external debt is reduced by the fact that the major debtor is the U.S. government.) This is reinforced by the fact that in most industrial countries, it is not possible to treat foreign creditors differently from domestic creditors.

II. Simulations of the U.S. Balance of Payments

With these qualifications in mind, we now consider projections of the U.S. balance of payments to determine whether the key criterion for current account sustainability—the stability of the NIIP/GDP ratio—is likely to be met. For this exercise, we use the Federal Reserve Board’s partial-equilibrium model of the balance of payments, which is described briefly below. (A more complete description is provided in the appendix to Bertaut, Kamin, and Thomas, 2008.)

Before proceeding, we address the desirability of using a partial equilibrium model—which assumes the key macroeconomic drivers of the current account to be exogenous—to assess current account sustainability. In principle, the new generation of forward-looking dynamic general equilibrium models might be better suited for longer run macroeconomic projections. However, in these models, trade deficits and external debt represent equilibrium responses to shocks, and asset prices and deficits start to correct long before economies reach any putative debt limits (see, for example, Erceg, Guerrieri, and Gust, 2006). Conversely, much of the debate over U.S. external sustainability assumes that U.S. indebtedness may expand until it reaches certain limits, after which correction may be triggered. In this context, it makes sense to use a partial equilibrium model, which does not assume spending adjusts endogenously in response to future financing constraints, to forecast the paths of external imbalances and debt under
plausible assumptions about output growth and prices, and then to assess whether those paths are sustainable. Moreover, the U.S. international transactions (USIT) partial-equilibrium model described below treats the U.S. balance of payments in considerably more detail than any general equilibrium model currently in use, and such detail is essential to assessing current account sustainability.

The U.S. International Transactions Model

The USIT model consists of 491 equations including 26 econometrically estimated behavioral equations, with the rest being identities and other computational equations. The model takes as exogenous projections for the central determinants of the U.S. external accounts, including: U.S. and foreign real GDP growth, U.S. and foreign inflation rates, U.S. interest rates, oil prices, and the foreign exchange value of the dollar. Based on these inputs, it then projects U.S. external balance variables in four broad categories: (1) trade flows, (2) nontrade components of the current account (especially investment income), (3) financing flows, and (4) the investment positions comprising the NIIP. Salient aspects of the modeling strategy and parameters are as follows:

1. Trade sector
   - Import prices for most major categories are projected based on the level of the dollar, foreign consumer price indices (CPIs), and the U.S. CPI; the rate of pass-through from changes in the dollar to changes in merchandise import prices is quite low, about one-third. Export prices depend on measures of U.S. production costs and final prices.
   - Real imports for most major categories depend on both the prices of imports relative to U.S. prices—with an elasticity of about unity—and on U.S. GDP. Real exports for most major categories depend on the price of exports relative to exchange-rate-converted foreign CPIs—also with an elasticity of about unity—and on a trade-weighted aggregate of foreign GDPs. Importantly, these equations incorporate the Houthakker-Magee asymmetry in income elasticities: the elasticity of real imports with respect to U.S. GDP is about 1.75 on average, exceeding the elasticity of real exports with respect to foreign GDP, of about 1.25 on average.²

²The theoretical basis for the Houthakker-Magee asymmetry remains ambiguous and the subject of controversy among trade modelers. Even so, the estimated coefficients in our trade models continue to exhibit the Houthakker-Magee asymmetry for goods trade. For services trade, in contrast, the income elasticity for exports exceeds that for imports. However, as goods trade exceeds services trade, the income elasticity for total imports exceeds that of total exports.
• Both the quantity and price of oil imports are modeled separately, with the former based on trends in U.S. production and consumption of oil, and the latter based on current and prospective market developments.

(2) Nontrade components of the current account balance

• Investment income is projected by applying income rates of return to different categories of U.S. external claims and liabilities.
• Assumptions on U.S. interest rates are used to project income rates on portfolio (equity, bond and deposit) positions. Skipping ahead slightly to Figure 1, the rates of income on U.S. portfolio assets and liabilities have been roughly similar in recent years and are projected to remain so, at a level close to the projected U.S. short-term rate of interest, going forward.
• The income rate of return on foreign direct investment in the United States depends on the U.S. output gap. The rate of return on U.S. direct investment abroad depends on the foreign output gap and the relative price of oil.
• Historically, income rates on U.S. direct investment abroad have exceeded that on foreign direct investment in the United States. Although this gap has narrowed over the past decade, it remains large and we project it to remain large in the future.3
• Transfers are projected exogenously, based on recent trends.

(3) Financing flows

• Once the current account balance is projected, this pins down the amount of net financing flows into (out of) the U.S. economy. However, two additional facets of these flows must be specified. First, financial flows must be allocated among the various categories: direct investment and portfolio investment. Second, a given amount of net financing may be associated with any number of combinations of gross financing flows. For example, an $800 billion net inflow may be achieved by $800 billion in gross inflows from abroad, combined with zero outflows; or it could be achieved by $1,600 billion in gross inflows and $800 billion in gross outflows.

3A number of explanations have been advanced for the asymmetry of rates of return on direct investment, including greater efficiency of U.S. firms, better project selection by U.S. firms, younger and thus less mature investments for foreign firms in the United States, greater competitive pressures in the U.S. market, or differences in tax treatment (see Higgins, Klitgaard, and Tille, 2005). None of these factors seem likely to disappear in the near term.
Gross direct investment flows to/from the United States depend on GDP growth in the recipient country.

Foreign flows into U.S. government assets and U.S. government flows into foreign assets are projected at their recent trends.

This leaves net private portfolio flows to balance the current account/financial account identity. The gross flows are constructed so as to produce growth rates in the stock of claims.
and liabilities that most closely match recent history while still having the implied net flows meet the net financing requirement.4

(4) **Investment positions**

- For each category of investment (direct investment, private portfolio, government portfolio, etc.), the gross asset or liabilities position in a given year will be equal to the position in the preceding year plus (a) financial flows (positive or negative) in that category during the year, and (b) valuation changes in the position.
- For all the simulation exercises, we report the direct investment positions and the NIIP using the current cost measure of direct investment.5
- In the simulation projections presented here, we include valuation changes to the direct investment positions alone, not the portfolio positions. The valuation changes for the direct investment positions reflect changes in exchange rates and in domestic prices of the assets (land, machinery, structures, and so on) comprising the position.

**Key Assumptions for the Projection**

The most important assumptions underlying the baseline balance-of-payments projection are shown in Figure 1.

- The real multilateral exchange value of the dollar is held constant at its level at the beginning of 2008.
- For U.S. real GDP growth, rates in 2008 and 2009 are based on Organization for Economic Development (OECD) projections, while longer-term growth rates of 2.4 percent are based on the OECD assessment of U.S. potential GDP growth in 2009 (OECD, 2008). In between the near term and farther out, growth jumps temporarily to restore the actual level of GDP to potential, after which growth subsides to its potential rate and the output gap remains at zero.
- GDP growth in the foreign industrial economies is projected in the same manner, based on OECD (2008). GDP growth in developing countries is based on the IMF’s *World Economic Outlook* projections for 2008, on the average growth rate from 1997 to 2007 for further out, and also

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4This is accomplished by starting out with trend extrapolations of gross inflows and outflows. If more financing is needed, gross inflows are adjusted up and outflows are adjusted down symmetrically; the reverse occurs if less financing is needed.

5This is the U.S. Bureau of Economic Analysis’s (BEA) preferred measure as it avoids many methodological issues associated with estimating the stock market value of nontraded equity positions. In addition, for our simulations, using the current cost measure means our estimates do not depend on our assumptions about future stock market movements.
incorporates a transitional period to restore levels of GDP to their potential levels.

- For 2008 and 2009, U.S. long- and short-term interest rates are based on OECD (2008). Beyond a transitional period, long-term rates are set equal to the projected growth of nominal GDP: real growth of 2.4 percent plus inflation of 2 percent, based on the OECD 2009 projection. Short-term rates are set 1 percentage point below long-term rates.

- The price of imported oil is set flat at its value at the beginning of this year. Note that imported oil is comprised of a mix of different grades, and its price is accordingly lower than that of West Texas Intermediate, whose price approached $100 per barrel around that time.

Model Projections

The Baseline Projection

As indicated in Figure 2, after some initial wiggles, real exports grow at a pace of nearly 5 percent over most of the projection period. Real imports, reflecting their higher elasticity with respect to GDP, grow a touch faster than 5 percent. In consequence, the trade balance widens gradually as a share of GDP, reaching about 5 percent by 2020. The nonoil trade deficit expands at a faster pace, as the volume of oil imports rises more slowly than other types of imports, consistent with trend declines in the oil-intensity of U.S. GDP. The current account balance also declines gradually as a share of GDP, reflecting not only the widening trade deficit, but also a (long-expected) shift in the balance on investment income from positive to negative as the NIIP gets more negative; net investment income and the current account deficit would deteriorate even faster were it not for the asymmetry in the rates of return on foreign direct investment. Finally, reflecting all of these developments, the NIIP/GDP ratio deteriorates steadily over the projection period, reaching over 60 percent of GDP by 2020.

Because of their importance for net investment income and thus the current account balance, the bottom panels of Figure 1 provide more detail on the composition of projected gross investment positions and on the rates of return on these positions. The bottom right panel indicates that the high rates of return on direct investment claims that we are assuming are well in line with past history, while the rate of return on direct investment liabilities is, if anything, generous by historical standards. The similarity in our projection of rates of return on portfolio claims and liabilities is also supported by history. The bottom left panel shows that increases in U.S. direct investment claims, which have the potential to significantly improve the net investment income balance owing to their high rate of return, do not appear out of line with the evolution of other categories.

\(6^{6}\)The saw-toothed pattern of the trade balance is caused by a residual seasonal pattern (even after seasonal adjustment by the BEA) in oil imports.
We draw four central conclusions from this projection. First, based on the standard criterion—stability of the NIIP/GDP ratio—the U.S. current account balance is not sustainable in the long term. The NIIP/GDP ratio deteriorates by more than 40 percentage points of GDP by 2020, or very roughly 4 percentage points per year. But, second, even if the current account...
is unsustainable, it is probably less unsustainable than would have been the
case had we started the projection back in, say, 2000; at that time, as
indicated at the top of Figure 1, the real value of the dollar was considerably
higher, and U.S. and foreign growth seemed more similar. Moreover, and
third, it is doubtful that, by 2020, the U.S. balance of payments will be
entering any danger zone where external adjustment will be urgently
required. Although the level of the net debt appears quite elevated (this
will be discussed further below), net investment income payments still
represent a paltry $\frac{1}{2}$ percentage point of GDP. This debt burden implies only
a minimal drag on spending, nor would it wave a red flag to investors
concerned about the creditworthiness of the U.S. economy. Fourth and
finally, even if investors took more signal from the NIIP/GDP ratio than
from the net investment income balance, it would take quite a few years
before the net debt rose above 60 percent of GDP.

**Comparison with Other Projections**

Compared with some previous exercises in projecting the U.S. balance of
payments, our baseline projection pushes considerably farther into the future
the date at which the U.S. external debt becomes a concern. For example,
writing nearly a decade ago, Mann (1999) projected that with an unchanged
exchange rate and standard growth assumptions, the U.S. current account
deficit would reach 8 percent of GDP by 2010 and the NIIP/GDP ratio
would reach $-64$ percent of GDP. With updated assumptions, Mann (2004)
projected the current account deficit would reach roughly 13 percent of GDP
by 2010. Cline (2005) projected that by 2010, the current account deficit
would reach $7\frac{3}{4}$ percent of GDP and the NIIP/GDP ratio would reach 50
percent. Comparing model simulations is difficult, but several factors likely
contribute to the more benign outlook in our projections compared
with Mann (1999 and 2004) and Cline (2005): the dollar has fallen further
from the levels assumed in their projections; a combination of valuation
changes and data revisions have boosted the starting point for our
projections of net investment income; and valuation changes and data
revisions have boosted the starting point for our projections of the NIIP.

Two other projections might be mentioned, although they are more
difficult to compare with ours. Higgins, Klitgaard, and Tille (2005) construct
a scenario in which the NIIP/GDP reaches $-65$ percent of GDP in 2015 and
$-89$ percent of GDP by 2025; this represents faster deterioration than in our
projections, but it is based on the assumption that the current account deficit
is fixed at 6 percent of GDP, a higher deficit than we project for most of the
projection period. By contrast, Kitchen (2007) develops a projection in which
the NIIP/GDP ratio reaches only $-39$ percent of GDP by 2015—compared
with about $-50$ percent of GDP in our baseline—but this projection assumes
dollar depreciation of over 1 percent annually. Notably, Kitchen’s analysis,
like ours, assumes a persistent rate of return differential favoring U.S. direct
investment abroad, and thus he also projects a very small deficit on net investment income, notwithstanding a still-substantial net external debt.

**Alternative Projections**

We believe the baseline projection described above to be plausible, but certainly the confidence interval around future projections must be very large indeed. Accordingly, in this section we present several alternative projections to illustrate the range of uncertainty, shown in Figure 3.

In the first alternative projection, the rate of foreign GDP growth is increased by about ½ percentage point, so that it grows about 4 percent annually. As shown by the dot-dot-dashed lines, the trade and current account deficits flatten out and then start narrowing; the net investment income balance is much improved, as higher foreign growth leads to more high-earning U.S. direct investment abroad; and the NIIP/GDP ratio deteriorates more slowly. Hence, with this relatively small alternation of assumptions, the present configuration of asset prices, growth rates, and exchange rates would likely be sustainable in the long term.

In the second alternative projection, the rate of foreign GDP growth is lowered by about ½ percentage point and U.S. GDP growth is boosted about ½ percentage point, so that they both grow at about 3 percent annually. As denoted by the dashed lines, under this scenario, the current account deficit widens to more than 8 percent of GDP by 2020, the NIIP/GDP ratio deteriorates beyond −70 percent, while the balance on net investment income now declines to about −1 percent of GDP. These outcomes are less sustainable than those in the baseline projection, but nevertheless, the debt-service ratio remains benign.

In the third alternative projection, shown by the dotted lines, oil prices rise at 5 percent annually. In this scenario, the trade deficit widens to nearly 8 percent and the NIIP/GDP ratio also deteriorates beyond 70 percent of GDP by 2020. But surprisingly, the balance on net investment income now declines to about −1 percent of GDP. These outcomes are less sustainable than those in the baseline projection, but nevertheless, the debt-service ratio remains benign.

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Finally, we consider a scenario—the dot dashed lines—in which the dollar declines by 1 percent annually going forward and, as investors demand a higher return to compensate, U.S. interest rates rise by 1 percentage point as well. In this scenario, which might be regarded as quite gradual correction, the trade deficit is considerably reduced relative to baseline. This positive effect on external balance is partly offset by the higher payments required on U.S. portfolio liabilities, so that the net investment income deficit is larger than in the baseline. However, on balance the current account deficit is narrower than in the baseline projection and the NIIP/GDP ratio slightly less negative as well.
To sum up, the risks to our projection are both on the upside and the downside, and we believe the baseline represents a plausible modal scenario.
Ex Post Historical Simulation

Obviously, beyond uncertainty about the exogenous variables in our simulations, another source of error in our projections is parameter uncertainty. To at least partially address this concern, Figure 4 presents the results of a historical simulation.
simulation of the model starting in the fourth quarter of 1994 and ending in the fourth quarter of 2007. The exogenous variables are set to their actual historical values, while the endogenous variables are simulated dynamically.

The model does a surprisingly good job of tracking movements in the trade and current account balances. The predicted ratio of the NIIP to GDP generally follows the actual path until 2003 or so, after which the predicted path continues to deteriorate while the actual NIIP/GDP ratio becomes less negative. However, much of the rise (to less negative values) of the actual NIIP/GDP ratio reflected revisions to the data which uncovered more U.S. assets abroad, and this could not be anticipated by the model. The predicted path of net investment income also follows the broad contours of the historical data except during 2001–04, when mispredictions of rates of return on investment income (not shown) and the noted revisions to the NIIP cause predicted net investment income to undershoot actual values.

All told, the ex post historical simulation provides some comfort that our model can give us useful insights into the outlook for the U.S. external balance.

**Sensitivity to the Key Export Elasticity**

The incorporation by our model of the Houthakker-Magee asymmetry in income elasticities leads to forecasts of larger current account deficits and net debt than if we assumed the income elasticities for imports and exports were equal. Although the USIT model’s trade equations have tracked history well, recent work by Thomas, Marquez, and Fahle (2008) suggests that mismeasurement of foreign prices may have introduced a downward bias to the estimated income elasticity for exports. If this is the case, then the Houthakker-Magee asymmetry embodied in the USIT model is unduly pessimistic for external adjustment.

The Thomas, Marquez, and Fahle measure of foreign prices is based on a weighted average relative price (WARP) which combines deviations from purchasing power parity (PPP) price levels with moving trade shares. Unlike conventional measures, the WARP treats a shift in trade shares from high-price countries to low-price countries as an effective decrease in foreign prices, even if all price levels remain unchanged. Since the trade shares of China and other low-price producers have increased markedly over the past 25 years, the WARP-based measure of foreign prices has risen less than standard measures, and thus the WARP-based measure of the relative price term for an export equation (price of U.S. exports/foreign prices in dollars) has increased relative to conventional measures.

We reestimated the USIT equation for goods exports (excluding high tech) using the WARP-based measure of foreign prices and found the income elasticity to be higher than in the standard formulation, while the estimated price elasticity was little changed. 7 The WARP formulation boosts the

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7In effect, the larger impetus to exports owing to the higher income elasticity is offset by the smaller impetus from prices owing to the less pronounced decline in relative export prices.
overall income elasticity for total goods and services exports reported in Section II from 1.25 to 1.43, diminishing but not eliminating the Houthakker-Magee asymmetry.

Results from a simulation incorporating the higher export elasticity, with other assumptions at their baseline levels, are reported in Figure 5. Higher
export growth produces a slight narrowing in the trade and current account balances, as a share of GDP, over the forecast horizon. However, the NIIP and net investment income still deteriorate as a share of GDP, albeit more slowly than in the baseline.

**How Negative a NIIP Is Worrisome?**

Our baseline projection of growing net indebtedness suggests that the U.S. current account is not sustainable in the long run. However, theory provides no guidance as to how large the negative NIIP/GDP ratio, or the ratio of net investment income payments to GDP, could become before triggering an adjustment that narrows the current account deficit. Accordingly, while the NIIP slated to reach roughly −60 percent of GDP by 2020 in the baseline scenario, we do not know whether, in reality, the current account deficit would be forced to adjust much before or much after it reached that point.

To shed some light on this issue, we look at the experience of other industrial countries. Figure 6a presents NIIP/GDP ratios for 19 industrial economies in 2006, the latest year for which these data are available for a broad array of countries. The exhibit makes clear that the current level of the NIIP/GDP ratio for the United States, about 20 percent, is well within normal bounds. More importantly, about five countries had negative NIIP/GDP ratios in the neighborhood of 60 percent or higher: New Zealand, Greece, Portugal, Australia, and Spain.

Figure 6b presents data on the ratio of net investment income to GDP for a similar group of industrial economies. A couple of countries—Ireland and New Zealand—have negative net investment income balances of 5 percent of GDP or more, while several more have balances in the −1½ to −2½ percent range. Under the baseline scenario, the U.S. net investment income balance does not reach −½ percent of GDP until 2019.

In considering sustainability, the NIIP and net investment income are compared with GDP because GDP is a measure of a country’s ability to service its debt. Panels (c) and (d) of Figure 6 present ratios of the NIIP and net investment income to an alternative measure of debt-repayment capacity, exports of goods and services. By these measures, the United States moves up a few rungs on the ladder of highly indebted countries, but remains below the five countries (listed above) with negative NIIP/GDP ratios exceeding 60 percent. According to the baseline projection described above, the U.S. NIIP/exports ratio expands to −440 percent by 2020, a little larger than the 2006 debt/export ratios for the most highly indebted countries shown in Panel (c) of Figure 6. However, in the baseline projection, the U.S. negative net investment income balance as a share of exports remains smaller than 5 percent, which is considerably below many of the ratios shown in Figure 6, Panel (d). Moreover, given that a country can reduce imports or expand exports as needed to repay external debt, we feel the NIIP/GDP and net investment income/GDP ratios shown in Panels
(a) and (b) of Figures 6 are probably better proxies for debt-repayment capacity.

A more relevant set of comparisons may involve international investment positions of economies at the time that they begin to experience current account adjustment. Panel (e) of Figure 6 presents the NIIP/GDP ratio at the onset of current account adjustment; the adjustment years are the same ones identified in research by Croke, Kamin, and Leduc (2006), and data are available for 15 of the 23 episodes identified. There is a very wide range of NIIP/GDP ratios associated with current account adjustment, suggesting either that many adjustments were triggered well before NIIPs reached some notional limit, or that this limit varies across countries. Panel (f) of Figure 6 indicates a similarly wide dispersion of data on the ratio of net investment income to GDP during current account adjustment episodes. Notably,
however, in 16 of the 22 episodes, the negative net investment income ratio exceeded 1 percent, a level that, in our baseline projection, the United States does not reach in the next 12 years.

Is the United States Special?

The above analysis suggests that, even as far away as 2020, the U.S. net external debt will still be no larger than it is today for five industrial economies, and the servicing burden on that debt will be relatively minor. But how relevant is the experience of other industrial economies for the United States?

Some arguments suggest that measures of external indebtedness could become even greater for the United States than for other industrial economies before adjustment was needed. First, an increasingly prevalent view holds that because the United States offers especially deep and liquid financial markets, global investors find U.S. assets unusually attractive and would be especially willing to finance the large deficit for a long period and on relatively cheap terms (Cooper, 2005; Hubbard, 2005).8

Second, among the different categories of U.S. debtors, a key international borrower is, of course, the federal government. At the end of 2007, foreign holdings of U.S. treasuries and agency debt amount to about $3.8 trillion, accounting for 22 percent of total U.S. foreign liabilities. Given the U.S. government’s commitment to the quality of its debt and its access to effective taxation, this likely makes U.S. external debt, overall, more creditworthy than if it had been issued primarily by U.S. households.

Third, as is often remarked, because the United States’ foreign-currency-denominated assets well exceed its foreign-currency-denominated liabilities, a decline in the dollar tends to reduce the net debt. This makes U.S. debt-servicing less vulnerable to dollar depreciation, and hence probably raises the level of sustainable debt.

A number of other considerations, however, suggest that the United States may have a diminished scope to issue external debt compared with other countries. Importantly, some of the highly indebted economies shown in Figure 6 may themselves represent special cases: Given their abundant resources, it may be sensible for Australia and New Zealand to import substantial capital and repay slowly over a long time horizon. Similarly, the large deficits and debts of Portugal, Spain, and Greece may be an artifact of their integration into the European Union and the euro area.

Moreover, put simply, the United States is the largest economy in the world, and heavy issuance of liabilities could saturate global demand. This consideration is addressed below.

8However, Curcuru, Dvorak, and Warnock (2008) and Gruber and Kamin (2008) present evidence contradicting the view that global investors place a special premium on U.S. assets.
III. Measures of Investor Exposure to U.S. Assets

The NIIP/GDP ratio is primarily a signal, albeit a highly imperfect one, of the weight of an economy’s debt service obligations; when the NIIP/GDP ratio reaches a certain size, investors may decide to limit their acquisition of the economy’s assets, fearing that larger NIIPs may not be serviceable. In addition to concerns about creditworthiness, however, investors may also seek to limit their exposure to an economy because that exposure threatens to breach a certain share of their portfolio. The NIIP/GDP ratio is not a very useful measure of this type of exposure, that is, of the weight of U.S. assets in foreign portfolios. In this section, we address several more direct measures of the exposure of investors to U.S. assets.

Recent Measures of Foreigners’ Exposure to U.S. Assets

Panel (a) of Figure 7 plots foreign holdings of all types of U.S. securities. Consistent with the increase in the U.S. net indebtedness position, holdings of all of these securities have risen in recent years. As panel (b) of Figure 7 illustrates, foreign holdings account for a noticeably growing share of the amounts outstanding of Treasury and agency securities, but there has been a less noticeable increase in the foreign share of U.S. corporate bonds and equities, as the total stocks of these securities have also increased rapidly.

A more relevant point for understanding how foreign exposure to U.S. securities may have changed is to consider how rapidly global market capitalization has grown, and how large foreign holdings of U.S. securities are compared with other securities in their portfolios. Figures 7c and d plot the share of U.S. equities and bonds in the global market capitalization of those instruments. These measures gauge whether persistent U.S. current account deficits have been associated with more rapid issuance of equity and bond liabilities than is occurring in other economies. In fact, the data show little net rise since the late 1990s in the U.S. share of global market capitalization for these instruments, likely reflecting two factors. First, financial deepening is progressing rapidly abroad, and increased securities issuance has led to ratios of market capitalization to GDP abroad that approach those of the United States. Second, declines in stock prices, on balance, since 2000 and declines in the

9Mann (1999, 2002, and 2003) and Cline (2005), among others, also draw a distinction between creditworthiness- and exposure-based criteria for sustainability, and present calculations based on these concepts. Concerns about creditworthiness and exposure are not necessarily unrelated. In a portfolio balance model, investors allocate their wealth to different assets, based on expected returns and uncertainties about those returns. Therefore, the more creditworthy an asset is considered to be, the higher the share in wealth allocated to that asset.

10Balakrishnan, Bayoumi, and Tulin (2007) find that declining home bias and financial deepening account for most of the financing of the recent large U.S. current account deficits, rather than increases in the share of U.S. assets in foreign portfolios.
dollar since 2002 have also kept exposures to U.S. assets in terms of the shares held in check.

Panels (e) and (f) of Figure 7 address the question of how large the share of U.S. assets is in foreign portfolios, and compares, for the limited number of countries and years for which survey data are available, the change over

Sources: (a) Authors’ estimates based on Treasury International Capital data; (b) authors’ estimates based on Treasury International Capital data and Flow of Funds accounts; (c) S&P, Global Markets Factbook 2007; (d) Bank for International Settlements data, adjusted for Brady bonds outstanding; (e)–(g): authors’ estimates based on IMF, Coordinated Portfolio Asset Surveys.
time in (1) the aggregate share of holdings of U.S. equities and bonds in overall holdings by foreigners of these instruments (the dark shaded bars), and (2) the aggregate share in overall holdings by foreigners of these instruments of claims on residents outside their own countries (the light shaded bars).\textsuperscript{11} Note that overall holdings include not only a country’s cross-border holdings of equities or bonds, but also its holdings of its own domestic equities or bonds. A number of observations can be made. First, the share of U.S. equities and bonds in the aggregate portfolios of foreigners has not risen much since 1997, and most or all of that increase took place between 1997 and 2001.\textsuperscript{12} Second, compared with the share of U.S. assets in portfolios abroad, the share of all external (to them) assets in portfolios abroad has risen by as much or more. This relationship may be seen more easily in Panel (g) of Figure 7, which indicates the share of U.S. equities and bonds in the external portfolios of foreigners. (In these calculations, holdings of a given country’s domestic securities are subtracted from its overall holdings to arrive at its external holdings.) The shares of both U.S. equities and bonds in these portfolios rose between 1997 and 2001, but have declined or been little changed since 2001.

Panels (a) and (b) of Figure 8 assess the extent to which foreigners are underweight in U.S. assets and compares that to the extent that they are underweight in external assets more generally.\textsuperscript{13} Foreigners are considered to be appropriately weighted (in terms of a standard portfolio allocation model) in U.S. equities, for example, if the share of U.S. equities in their total equity portfolio—the dark shaded bars in Figure 7, Panel (e)—is equal to the share of U.S. equities outstanding in global equity capitalization—as shown in Figure 7, Panel (c). We thus compute foreigners’ relative portfolio weights in U.S. assets (plotted against the vertical axis) by dividing the share of U.S.

\textsuperscript{11}Data on foreign holdings of U.S. and other external securities are derived from the IMF’s Coordinated Portfolio Investment Surveys (CPIS). Because the CPIS captures nonreserve holdings only, we impute an amount for holdings of both U.S. and other external securities held as reserves using data from the IMF SEFER and COFER surveys. Most industrial countries and a number of emerging-market countries now participate in the CPIS. In terms of major holders of U.S. securities, they exclude investments held in some major custodial centers, by most Middle East oil exporters, and by China. Holdings of U.S. securities accounted for by CPIS countries in 2006 represent about 70 percent of total U.S. securities held by foreign investors. See Bertaut, Griever, and Tryon (2006) for a discussion of the methodology for imputing reserve holdings, and for a more complete discussion of the comparability between holdings of U.S. securities as measured by the CPIS and by U.S. liability surveys. Data on holdings of domestic securities are derived from national source financial balance sheet accounts where available, and otherwise from estimates of domestic equity and bond market capitalization.

\textsuperscript{12}The 1997 and 2001 figures are not strictly comparable because more countries participated in the 2001 CPIS than in the 1997 CPIS. However, the difference in coverage is less critical for comparing relative shares than absolute holdings, and indeed the increase in shares held between the two years owes largely to the increases registered by countries that were participants in both years.

\textsuperscript{13}See Bertaut and Griever (2004), for a fuller elaboration of this approach.
securities in the total portfolio of foreigners by the size of the U.S. market relative to the world market:

\[ \text{Share in US securities} = \frac{\text{foreign holdings of US securities}}{\text{foreign holdings of all securities}} \]
A relative portfolio weight of 1 implies appropriate weighting of U.S. assets in foreigners’ portfolios, while a relative weight less than 1 implies an underweighting of U.S. assets.

A similar calculation is undertaken for the foreigners’ relative portfolio weight in all external securities, where for any given country outside the United States, external refers to all countries external to that country (including the United States).¹⁴ A value of this calculation (plotted against the horizontal axis) less than 1 implies that foreigners are underweight in assets outside their own country and overweight in domestic securities—that is, they exhibit home bias.

Observations on the dashed 45 degree line in Panels (a) and (b) of Figure 8 would indicate that foreigners are equally underweight U.S. and other external assets. The data indicate that foreign investors are both underweight in U.S. assets and in external assets more generally—that is, they exhibit home bias. In 1997, the bias against U.S. assets appeared to be considerably greater than the bias against other countries’ assets. Since then, foreigners appear to have reduced their bias against U.S. and other countries’ equities about equally. The same appears true for foreign holdings of U.S. and other external bonds between 2001 and 2006; between 1997 and 2001, they appear to have reduced their bias against U.S. bonds by somewhat more. Although foreigners’ home bias has declined over this period, foreigners remain more underweight in U.S. securities than they do in external securities in general.

The bottom line from Figure 7 and Panels (a) and (b) of Figure 8 is that, in spite of the expansion of U.S. external liabilities since the mid-1990s, neither the shares nor the relative weights of U.S. assets in foreigners’ portfolios have increased to any meaningful extent. This somewhat surprising result is, in part, a reflection of the decline in the dollar since 2002, which has reduced the value of holdings of dollar-denominated instruments relative to those denominated in other currencies. However, this result also reflects the increases of asset holdings abroad more generally, of which increased holdings of U.S. assets are just a part.¹⁵ All told, there is no evidence that an overhang of excessive foreign exposure to U.S. assets is developing which would require further adjustments in the U.S. current account balance or in U.S. asset prices to correct.

¹⁴In this analysis, we consider intra-euro area holdings of other euro area country securities as domestic securities.

¹⁵Higgins and Klitgaard (2007) reach a similar conclusion that financial globalization has had the result that despite increases necessary to finance U.S. current account deficits, there has not been an unusual buildup of U.S. securities in foreign portfolios.
Prospective Future Movements in Foreigners’ Exposure to U.S. Assets

Even if the exposure of foreigners to U.S. assets, relative to a number of benchmarks, does not appear to have increased much over the past decade, it is possible that the financing of continued current account deficits would push this exposure to more worrisome levels in the future. With the projected increase in the NIIP in our baseline projection described above, foreign investors will of necessity acquire many additional U.S. assets. How large a share of their portfolio foreign investors are willing to acquire is an open question. It depends importantly on factors beyond the evolution of the U.S. NIIP, including the growth of securities issuance in foreign economies and thus the share of U.S. securities in global market capitalization.

To devise an estimate of the potential magnitude of the effects of the projected increase in the NIIP on foreign portfolios, we perform the following exercise: We assume that both U.S. and foreign total market capitalization (reflecting equities and bonds combined) grow at their respective rates of nominal GDP, thus keeping their market cap/GDP ratios constant. As described in Section II, the USIT model simulations project gross portfolio flows into and out of the United States so that the resulting net financing is sufficient to meet the balance of payments requirements while also keeping the growth in the gross positions as close as possible to their recent historical rates. Thus, the total portfolio of foreign investors is projected to grow along with foreign market cap (minus the amount acquired by U.S. investors), plus estimated increased holdings of U.S. securities. Finally, we perform an alternative calculation in which the projections for financial flows into and out of the United States remain the same, but both U.S. and foreign market cap grow at the same rate (equal to the average of foreign and U.S. growth), so that U.S. market cap stays constant as a share of global market cap.

As illustrated in Panels (c)–(f) of Figure 8, the results of this projection exercise generally point to increases in the share and relative weight of U.S. assets in foreigners’ portfolios, but it is not clear those increases would be worrisome. Figure 8c shows projections for total U.S. market cap as a share of global market cap. (This panel is comparable to the share concept in Panels (c) and (d) of Figure 7, but combines equity and bond capitalization.) Under the baseline assumption, total U.S. market cap declines to about 32 percent of global market cap by 2020, reflecting the slower projected growth of U.S. nominal GDP and thus of U.S. market cap, compared with foreign GDP and market cap. In the alternative simulation, the U.S. share of market cap stays constant, by design.

---

16Specifically, we assume that as foreigners acquire additional nondirect investment claims on the United States, the share of these claims that are securities (as opposed to bank deposits, trade credits, and so on) will mirror their current share in nondirect investment claims. Similarly, as U.S. residents acquire additional nondirect investment claims on foreigners, the share of these claims that are in the form of securities will mirror their current share.
Figure 8, Panel (d) shows that if foreigners acquire U.S. assets sufficient to accommodate the rise in the NIIP, the share of U.S. market cap held by foreigners rises from around 20 percent currently to near 40 percent in either simulation.

Figure 8e shows the evolution of U.S. securities as a share of the total portfolio of foreigners (comparable to the dark shaded bars in Panels (e) and (f) of Figure 7) and Figure 8, Panel (f) shows the corresponding evolution of the relative portfolio weights in U.S. securities (comparable to the relative portfolio weights in U.S. securities shown in Figure 8, Panels (a) and (b). Under either simulation, the share increases to around 20 percent by 2020, and the corresponding estimated relative weight of U.S. securities in foreigners’ portfolios increases to 55 to 60 percent. This increase in the relative weight is substantial, but even so, a relative weight of 0.55 implies that foreigners would remain underweight in U.S. assets. Moreover, although we cannot perform a similar projection of the relative weight of external securities in foreigners’ portfolios, presumably that weight would also be rising substantially, assuming the trends documented in Figure 8, Panels (a) and (b) continue.

IV. How Responsive are Asset Prices to International Balance Sheet Indicators?

To make an assessment of how imminent is a correction in the current account balance, one must not only project the likely evolution of the external debt burden and the time at which it breaches some threshold of sustainability. One must also be able to predict, at such time as this breach occurs, how rapidly investors will pull back from a country’s assets, thus boosting interest rates, pushing down the currency, and inducing current account adjustment. As an initial rough cut at addressing this issue, we evaluated the sensitivity of interest rates and exchange rates to measures of international balance sheet positions.

17The starting figures for 2006—a 12 percent share of U.S. assets in total portfolios corresponding to a U.S. portfolio weight of about 0.28—are slightly larger than the shares and weights in Figure 7 (a and f) and Figure 8 (a and b) because for this exercise, we base total foreign holdings of U.S. securities on the more comprehensive liabilities estimates that underlie the NIIP calculations, and thus we are able to include all foreign holdings of U.S. securities, including those held by countries not participating in the CPIS surveys, notably international financial centers, Middle East oil exporters, and China. Note also that the average shares and weights will more closely resemble the bond shares and weights in the exhibits because the majority of foreign holdings are in the form of U.S. bonds.

18We are not able to project how changes in foreigners’ shares and weights held in total external securities compare with the projected changes in shares and weights held in U.S. securities. Although we are able to forecast total foreign (non-U.S.) market cap held by foreigner investors, we have no way of allocating what fraction of that foreign market cap reflects foreigner investors’ home country securities and what fraction reflects holdings of other foreign securities.
Long-Term Interest Rates

In our analysis, we estimate panel regressions using annual data for 22 industrial countries over the period 1975 to 2006. The dependent variable is the nominal long-term (usually 10-year) yield on government benchmark bonds. Following Gruber and Kamin (2008) and Warnock and Warnock (2006), a set of control variables includes the overnight money market interest rate, the four-quarter rate of CPI inflation, the four-quarter rate of real GDP growth, the standard deviation of the quarterly change in the long-term nominal interest rates over the preceding 12 quarters, the ratio of the structural (full-employment) fiscal balance to GDP, and two annual lags of the dependent variable.

The first column of Table 1 presents the results of this equation, estimated including only the control variables. The results are, for the most part, consistent with expectations. Increases in money market interest rates, inflation, and real GDP growth all boost nominal long-term bond yields by a statistically significant extent. Also as one would expect, increases in the volatility of interest rates boost yields and increases in the fiscal balance reduce yields, although these effects are not statistically significant.

The next column of the table adds year and country fixed effects, as well as a dummy variable that becomes one in 1999, with the creation of the euro area. The coefficients on the control variables are, for the most part, little changed.

The next four columns add, separately, to this equation four different measures of external balance, all lagged one year: the NIIP/GDP ratio, the ratio of net investment income (NIINCOME) to GDP, the current account (CAB)/GDP ratio, and the share of a countries’ gross external liabilities in the gross external assets of foreigners (external liability). The remaining columns present results of equations that include all four of these external balance measures together, but with different combinations of year and country fixed effects.

By and large, there is little evidence that the external balance measures examined here are associated with a significant effect on long-term yields. To the extent that any of these variables has a consistent and nearly significant effect of the expected sign, it is the NIIP/GDP ratio. At its largest, the coefficient on this term is −0.005, implying that a 100 percent of GDP negative NIIP would be associated with an increase in the long-term nominal interest rate of 50 basis points—this is a discernable, but not especially large, increment. The current U.S. NIIP of about −20 percent of GDP implies a boost to the interest rate of a fifth of that, of only 10 basis points. (This is a much smaller effect than estimated by Lane and Milesi-Ferretti, 2001, using a more limited set of control variables.)

In interpreting the coefficients on the external balance variables, a prominent identification problem should be acknowledged. In principle, countries with highly developed financial systems and strong investor protections should attract foreign investors. These large capital inflows, in turn, ought to lower U.S. long-term yields. Accordingly, a priori, it is not clear whether large external debts should be associated with higher interest
Table 1. Panel (Ordinary Least Square) Regressions for Interest Rates

Dependent Variable: 10-Year Nominal Government Bond Yields

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<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
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<th>(7)</th>
<th>(8)</th>
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<td>(0.060)</td>
<td>(0.062)</td>
<td>(0.064)</td>
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<td>(0.063)</td>
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<td>(0.060)</td>
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<td>−0.392</td>
<td>−0.484</td>
<td>−0.483</td>
<td>0.166</td>
<td>−0.393</td>
<td>0.175</td>
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<td>(0.288)</td>
<td>(0.308)</td>
<td>(0.296)</td>
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<td>Fiscal balance/GDP</td>
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<td>−0.025</td>
<td>−0.031</td>
<td>−0.022</td>
<td>−0.019</td>
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<td>t-stat</td>
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<td>t-stat</td>
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<td>External liability variable (–1)</td>
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<td>–0.119</td>
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<td>SE</td>
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<td>–0.77</td>
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<td>541</td>
<td>532</td>
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</table>


1Excludes Iceland.
2Excludes Canada, Iceland, and Italy.
rates, because they raise concerns among investors, or lower interest rates, because the large debts reflect strong investor interest.

However, the estimates shown in columns 2–6 and 9–10 should, to a large extent, control for the simultaneity problem described above, as they include country fixed effects which should capture any special attractiveness of a country’s assets. In fact, in columns 7 and 8, where no country fixed effects are included, the coefficient on the NIIP/GDP ratio is much smaller than in columns 9 and 10, where country fixed effects have been included. This suggests that the country fixed effects are, indeed, helping to control for the simultaneity problem.

**Exchange Rates**

If investors react to high debt levels by pulling back from a country’s assets, we should also observe a depreciation of the exchange rate. To assess whether this is the case, we re-estimated the panel equations shown in Table 1, but substituted the real exchange rate in place of the nominal long-term interest rate. We use the CPI-deflated multilateral exchange rate published in the IMF’s International Financial Statistics.\(^\text{19}\) (An increase indicates appreciation.)

In Table 2, the dependent variable is the percent deviation of the real exchange rate from its country-specific sample mean. The coefficient on the money market interest rate is positive and significant, while that on the inflation rate is about the same magnitude but negative and significant; together, these results confirm our expectation that increases in the real interest rate should positively affect the real exchange rate. None of the other control variables have a significant coefficient of the expected sign—this is perhaps not surprising, as exchange rate models are notoriously hard to estimate. The measured effects of the external balance variables are a mixed bag as well. The NIIP and net investment income generally have a positive effect on the real exchange rate as expected, but the effect is inconsistent in terms of sign and significance. Higher current account balances lead to lower exchange rates and larger external liabilities lead to higher exchange rates, the opposite of what we’d expect.

Finally, Table 3 reestimates these equations, with the dependent variable specified as the percent change in the real exchange rate from the previous year. Now, the current account balance appears to be positively associated with the real exchange rate, but the coefficients on the other variables remain insignificant or of the wrong sign.\(^\text{20}\)

\(^{19}\)These regressions are estimated for the same 22 industrial countries over the sample 1978–2006.

\(^{20}\)Gagnon (1996) finds evidence that net foreign assets scaled by trade flows are significantly associated with real exchange rates for a panel ending in 1995. In a multi-country probit study of industrial economies, Wright and Gagnon (2006) find that larger current account deficits are significantly associated with sharp real currency depreciations, but the magnitude of the effect is quite small. Other variables do not exert significant, robust effects on the probability of a sharp real depreciation.
### Table 2. Panel (Ordinary Least Square) Regressions for Exchange Rates

<table>
<thead>
<tr>
<th>Dependent Variable: Real Effective Exchange Rate as Percent Deviation from Sample Country Mean (An increase indicates appreciation)</th>
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<tbody>
<tr>
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<tr>
<td>REER percent deviation (−1)</td>
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<td>SE</td>
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<tr>
<td>t-stat</td>
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<tr>
<td>REER percent deviation (−2)</td>
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<td>Money market interest rate</td>
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<td>t-stat</td>
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<td>Inflation</td>
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<td>SE</td>
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<td>Real GDP growth</td>
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<td>SE</td>
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<td>t-stat</td>
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<td>Interest rate volatility</td>
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<tr>
<td>CAB/GDP (−1)</td>
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<td>t-stat</td>
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1Excludes Iceland.
2Excludes Canada, Iceland, and Italy.
Table 3. Panel (Ordinary Least Square) Regressions for Exchange Rates

<table>
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<th>Dependent Variable: Real Effective Exchange Rate Percent Change (An increase indicates appreciation)</th>
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<td>(1)</td>
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<tr>
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<td>REER percent change (−2)</td>
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<td>t-stat</td>
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|                | 0.065       | 0.023                            | 0.178      | 0.139    |
| SE             | 0.077       | 0.022                            | 0.458      | 0.639    |
| t-stat         | 2.00        | 1.28                             | 0.571      | 1.033    |
|                | 0.194       | 0.186                            | 0.639      | 1.033    |
| SE             | 0.124       | 1.033                            | 0.699      |          |
| t-stat         | 1.57        |                                  |            |          |

|                | 0.038       | −2.471                           | −4.995     | −2.187   |
| SE             | 0.654       | −3.17                            | −1.32      | −2.58    |
| t-stat         | 0.06        | −3.17                            | −1.32      | −2.58    |
|                | −1.35       | −1.50                            | −1.37      | −1.44    |
| Country fixed effects | No | Yes | Yes | Yes | No | Yes | No | Yes |
| Time fixed effects | No | Yes | Yes | Yes | Yes | Yes | No | Yes |

R²: 0.0557
SER: 5.5589
No. of observations: 488


1 Excludes Iceland.

2 Excludes Canada, Iceland, and Italy.
Summing Up

In this section, we found that greater external debt was associated with only small increases in interest rates and mixed effects on exchange rates. Accordingly, deteriorations of external balance positions do not seem to have been associated with significant pullbacks by global investors. Of course, modeling financial market prices is notoriously difficult, so it may not be surprising that we found little evidence of strong linkages among indebtedness, interest rates, and exchange rates. By the same token, however, the direst predictions that continued large U.S. current account deficits will lead to financial crisis remain unsubstantiated.

V. Conclusion

This paper has addressed three questions about the prospects for the large U.S. current account deficit. Is it sustainable in the long term? If not, how long will it take for measures of external debt and debt service to reach levels that could prompt some pullback by global investors? And if and when such levels are breached, how readily would asset prices respond and the current account start to narrow?

To address these questions, we started with projections of a detailed partial-equilibrium model of the U.S. balance of payments. Assuming plausible settings of macroeconomic indicators in the United States and abroad, as well as a flat real dollar, our projections indicate that the current account deficit will resume widening and the negative NIIP/GDP ratio will continue to expand. This suggests that the configuration of macroeconomic settings underlying the current account balance at present is not sustainable in the long term. Nevertheless, compared with earlier years when the dollar was much higher, the current account balance is likely somewhat less unsustainable at present.

Moreover, compared with other industrial economies, current levels of the U.S. debt and debt service are relatively modest, and foreign exposure to U.S. assets has been moving down rather than up. Our projections suggest that even by the year 2020, the negative NIIP/GDP ratio will be no higher than it is in five industrial economies today, and U.S. net investment income payments will remain below 1 percent of GDP. Absent changes in the dollar and consequent valuation adjustments, the share of U.S. claims in foreigners’ portfolios will likely rise over the next decade, but not to an obviously worrisome extent. All told, it seems likely it would take more than a decade for U.S. indebtedness to reach any limits of global investors’ willingness to extend financing.

This finding is consistent with the lack of consistent evidence that high levels of external indebtedness lead to subsequent reductions of current account deficits. See, among others, Lane and Milesi-Ferretti (2001 and 2002), Chinn and Ito (2007), and Gruber and Kamin (2008).
Finally, we explored the historical responsiveness of asset prices and the current account in a wide range of industrial economies to increases in different measures of net indebtedness and external imbalances. We discerned only a very small effect of external indebtedness on interest rates, and no clear and concerted effects on exchange rates. Accordingly, it is not clear that, even were the U.S. net debt to approach limits to its sustainability, the developments needed to correct the current account—changes in growth rates, assets prices, exchange rates, and the like—would materialize all that rapidly.

We would emphasize that these findings do not imply that U.S. current account adjustment is necessarily many years away. Many factors could trigger such adjustment, including, inter alia, a surge in foreign growth, declines in U.S. growth, or intensified concerns about U.S. current account sustainability. In fact, judging by the developments of the last several years, we may already be in the middle of an adjustment episode. Our point is rather that international balance sheet considerations likely are not sufficient, by themselves, to require external adjustment any time soon.

REFERENCES


Asset Prices and Current Account Fluctuations in G-7 Economies

MARCEL FRATZSCHER and ROLAND STRAUB*

The paper analyses the effect of equity-price shocks on current account positions for the G-7 industrialized countries during 1974–2007. It uses a Bayesian vector autoregression with sign restrictions for the identification of equity-price shocks and to test empirically for their effect on current accounts. Such shocks are found to exert a sizable effect, with a 10 percent equity price increase, for example, in the United States relative to the rest of the world, worsening the U.S. trade balance by 0.9 percentage points after 16 quarters. However, the response of the trade balance to equity-price shocks varies substantially across countries. The evidence suggests that the channels accounting for this heterogeneity function both through wealth effects on private consumption and to some extent through the real exchange rate of countries. [JEL E2, F32, F40, G1]


Current account positions have hardly ever been so dispersed globally as they are today. It is not only that the largest economy, the United States, has been recording a current account deficit in excess of 5 percent for several years, but other industrialized countries, such as the United Kingdom and Australia, and some emerging markets and transition economies have

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*Marcel Fratzscher is a division chief and Roland Straub a senior economist with the European Central Bank. The authors would like to thank the participants at the conference on “Current Account Sustainability in Major Advanced Economies” at the University of Wisconsin, Madison, and in particular our discussant, Ken West, for comments and discussion.
similar or even larger deficits. By contrast, countries such as China, Japan and oil exporters register corresponding large trade surpluses. At the same time, asset prices have gone through a marked cycle over the past decade, with equity markets rising substantially in the second half of the 1990s and in 2002–06 and declining in 2001–02. The financial market crisis of 2007–08 has made the importance of asset prices for the global economy more than apparent. Despite the financial crisis, the role of asset prices for the global economy will most likely increase further as financial markets deepen and emerging economies liberalize and integrate.

This paper analyses the impact of asset (equity) price shocks on the current account. The objective is not only to grasp the magnitude of the effect of asset prices on trade, but also to understand the channels through which this effect materializes. Asset-price shocks affect net exports through a wealth channel as households adjust saving and consumption decisions, and through an exchange rate and terms-of-trade channel, altering the relative prices of domestic and foreign goods. Equally importantly, asset prices may exert different effects across economies, as those with deeper yet more closed financial markets may respond more strongly.

The focus of this paper is on the G-7 industrialized countries and on the role of equity-price shocks during 1974–2007. We use the sign restrictions derived in Fratzscher and Straub (2008), who build an open-economy dynamic stochastic general equilibrium model, in which changes to asset prices influence private consumption through wealth effects. We then employ a Bayesian vector autoregression (VAR), following Canova and de Nicoló (2002), Uhlig (2005), and Peersman (2003), using sign restrictions to test for the effect of asset price shocks in the data. This methodology not only requires imposing a relatively small and intuitive number of identification restrictions, but importantly it also allows us to distinguish asset-price shocks from other types of structural shocks. Our empirical implementation follows closely that of Fratzscher, Juvenal, and Sarno (2007), who test for the effect of equity-market shocks, housing-price shocks and exchange rate shocks on the trade balance of the United States. They show that equity-market shocks and housing-price shocks have been important drivers explaining more than 30 percent of the variation of the U.S. trade balance, whereas exchange rates account for a much smaller share.

Our empirical findings show that asset prices exert a sizable effect on the trade balance of countries. The channels through which equity prices influence net exports are both through wealth effects on private consumption and to some extent through the exchange rate. An increase in asset prices tends to have a positive impact on short-term interest rates and inflation, and leads to an appreciation of the real effective exchange rate (REER) and a sizeable increase in consumption. Moreover, we find a large degree of cross-country heterogeneity in the impulse response pattern. The U.S. trade balance is among the most sensitive as net exports, on average, decline by 0.91 percentage points after 16 quarters in response to a 10 percent increase in asset prices.
in U.S. equity prices relative to the rest of the world. The trade balances of most other countries react substantially less.

The paper is related to three fields of the literature. A first strand focuses on the drivers of the large and persistent global current account imbalances. Several papers emphasize the importance of a saving glut (Bernanke, 2005) in many emerging markets and commodity-exporting countries, partly stemming from the underdevelopment and lack of integration of financial markets in those economies (Caballero, Farhi, and Gourinchas, 2006; Ju and Wei, 2006; Fratzscher, 2008), as well as the increasing role of ensuing valuation effects on gross international asset positions (Lane and Milesi-Ferretti, 2005; Gourinchas and Rey, 2007) and a precautionary motive as a rationale for high saving rates (for example, Chinn and Ito, 2007; Gruber and Kamin, 2007). Other studies to explain the dispersion in current account positions stress the role of productivity differentials (for example, Bussiere, Fratzscher, and Müller, 2005; Corsetti, Dedola, and Leduc, 2006), or link it to the great moderation, which has induced a decline in income volatility and uncertainty (Fogli and Perri, 2006).

As to the second area, a vast amount of literature identifies and measures the effect of price changes in various financial assets on private consumption (for example, Bertaut, 2002; Case, Quigley, and Shiller, 2005). Most of this literature finds a significant effect of both equity wealth and housing wealth on private consumption. However, there is still substantial controversy as to the magnitude and precise functioning of this channel as for instance exemplified by the conflicting results found by Palumbo and Whelan (2006) and Lettau and Ludvigson (2004). The effect of such a wealth channel on the external dimension of countries, in particular the current account and the exchange rate, has so far received little attention in the literature. From a current policy perspective, it has been argued by some that the U.S. dollar decline would have to be very large, as suggested by several studies (Blanchard, Giavazzi, and Sa, 2005; Obstfeld and Rogoff, 2005; Krugman, 2007).

The third area relates to the crucial issue of the structural interpretation of asset-price shocks. Although we can separate an asset-price shock from the standard macroeconomic shocks usually analyzed, it is not clear what asset-price changes represent structurally. One interpretation of an asset-price shock is that of a news shock, along the line of work by Beaudry and Portier (2006 and 2007), in which asset prices adjust because of altered expectations about the likelihood of future outcomes, such as to economic fundamentals. Such changes in expectations should then, in turn, be reflected in today’s asset prices as these represent the net discounted value of all future fundamentals.

This is also related to the work by Engel and Rogers (2006), who show that the large size of the U.S. current account deficit is consistent with expectations of an increasing share of U.S. output in the world. An alternative interpretation is that asset-price shocks reflect rational bubbles, as in Kraay and Ventura (2005) and Ventura (2001). They argue that the sharp
increase in asset prices over the past decade may largely reflect a bubble, which is rational because of market expectations that this increase may be persistent. Both interpretations are observationally equivalent to what we understand and see in the behavior of economic fundamentals. We are agnostic about these interpretations; the crucial point is that asset-price shocks reflect factors that function primarily through asset prices. The purpose and intended contribution of this paper is to improve our understanding of how this asset-price channel functions.

I. Methodology

Deriving the Sign Restrictions

This section discusses the set of sign restrictions to identify asset-price shocks. Thereby, we apply the strategy discussed, for example, in Peersman and Straub (forthcoming). In particular, we utilize restrictions, discussed in detail in Fratzscher and Straub (2008), which identify asset-price shocks uniquely and distinguish them from a set of other shocks that are discussed as determinants of current account fluctuations in the literature. Table 1 summarizes the sign restrictions used for the identification in our structural VAR. We associate positive asset-price shocks (that is an exogenous increase in asset prices) with a rise in consumption, inflation, and interest rates. As discussed in Fratzscher and Straub (2008), the rise in current stock market wealth triggers an increase in private consumption. The latter induces a surge in inflation rates, and under the assumption of an active monetary policy rule, an increase in interest rates.

Note that the latter reaction is fundamentally different from the response following technology and monetary-policy shocks. Technology shocks, for example, trigger a rise in consumption and a fall in inflation rates. Monetary policy shocks induce a positive response of consumption and inflation and are characterized by a fall in interest rates.\(^1\) Note that, although we base our sign restriction identification strategy on the predictions of a theoretical model, we do not have to restrict the response of the current account and the real exchange rate, the main variables of interest. In this respect, we can let the data speak for itself.

A crucial issue is the structural interpretation of asset-price shocks. The identifying restrictions above separate an asset-price shock from the standard macroeconomic shocks usually analyzed (technology, monetary policy, and government spending), without identifying the structural factors behind the asset price increase. Note that other demand-side shocks such as shocks to

\(^1\)In the discussed model a fall in government spending, financed, for example, by lump-sum taxes for simplicity, is associated by a rise in private consumption and inflation, but a fall in aggregate output. As a result, the reaction of policy interest rates depends obviously on the monetary-policy rule. As argued in Fratzscher and Straub (2008), a standard Taylor rule implies a fall in interest rates, as the rise in inflation is relatively small, but the response of output is more pronounced for a wide range of structural parameters.
time preferences or distortionary taxes might imply, under certain assumptions, similar patterns for the endogenous variables as asset-price shocks.

On the other hand, exogenous changes in distortionary taxes or time-preference rates are unlikely to be an important determinant of business cycles at a quarterly level.

What is our interpretation of asset-price shocks? As discussed above, one interpretation of an asset-price shock is that of a news shock (Beaudry and Portier, 2006 and 2007), in which asset prices adjust because of changed expectations about the likelihood of future outcomes, such as to economic fundamentals; or as in Engel and Rogers (2006), where current account changes are consistent with changing expectations of relative output shares.

Alternatively, asset-price shocks can echo rational bubbles, as in Kraay and Ventura (2005) and Ventura (2001). We are agnostic about these interpretations; the main point is that asset-price shocks reflect factors that function primarily through asset prices. The objective of the empirical exercise is to illustrate how this asset-price channel functions.

Model Specification and Data

Consider the following specification for a vector of endogenous variables $Y_t$:

$$
Y_t = a + \sum_{i=1}^{n} A_i Y_{t-i} + B e_t,
$$

where $a$ is a $(n \times 2)$ matrix of constants and linear trends, $A_i$ is an $n \times n$ matrix of autoregressive coefficients and $e_t$ is a vector of structural disturbances. Identification of the impact of structural disturbances requires imposing $n(n-1)/2$ restrictions on $B$, which we achieve by using the sign restrictions shown in Table 1. Our sign restriction approach is based on Canova and De Nicoló (2002), Uhlig (2005), and Peersman (2003), discussed in some detail in the next section.

Our VAR includes six variables: $Y_t = [EQ, c, i, \pi, TB, REER]$, a relatively standard specification as, for instance, also used in Fratzscher, Juvenal, and Sarno (2007), that is, private consumption ($c$), short-term interest rates ($i$), inflation ($\pi$), equity returns ($EQ$), as well as the trade balance ($TB$) and the REER.

<table>
<thead>
<tr>
<th>Table 1. Theoretical Impulse Response Functions</th>
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<tr>
<td>Consumption</td>
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<td>Asset-price shock</td>
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<td>Monetary-policy shock</td>
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Our country sample focuses on the G-7 industrialized countries. The time period for the empirical analysis is 1974–2007, using quarterly data. We use 1974 as the starting point of the analysis as it is the start of the floating exchange rate period after the collapse of the Bretton Woods system. We limit our analysis to the G-7 economies.²

For our empirical estimation we use relative variables, that is, we specify each variable in domestic vs. rest-of-the-world terms. More precisely, consumption $c$ is the difference in log private consumption in the domestic economy and log private consumption in the rest of the world, both expressed in U.S. dollar (using end-of-period exchange rates). Interest rates $i$ are the percentage difference of domestic short-term (money market) rates from those in the rest of the world, whereas inflation $\pi$ is the corresponding percentage difference in consumer price index inflation. The rest of the world for all three variables comprises the other economies in the benchmark sample with each country being weighted by its GDP share in the sample group.

Our preferred measure of asset prices $EQ$ is the difference between domestic equity returns and foreign equity returns, both measured in local currency terms. We use local currencies to express returns, rather than U.S. dollars, because we want to obtain a measure of asset-price shocks that excludes exchange rate movements.³ Moreover, we use shocks to equity prices, rather than changes to market capitalization, as our preferred measures because our primary interest is in the cross-country heterogeneity in the responses of the trade balance and the exchange rate. The rest-of-the-world group comprises the other countries in the sample, with each of these countries being weighted by their equity-market capitalization. We use equity-market capitalization weights, rather than GDP weights, because equity shocks are likely to affect the trade balance of countries partly through wealth effects, which in turn should be related to the size of financial wealth held by households, which is better proxied by market capitalization than GDP. In the section on the robustness analysis below we will discuss how alternative specifications of asset-price shocks influence the empirical findings.

The trade balance $TB$ is measured as a ratio to domestic GDP. We use the trade balance, rather than the current account, as we are interested in the effect of asset-price shocks on net exports and want to exclude the effect on income. As the final variable, the REER uses trade weights for a broad set of partner countries, and is expressed in logs.

As to the data sources, the trade balance, consumption, inflation and short-term interest rates come from the IMF’s International Financial

²Appendix Table A1 lists the countries included.
³Hau and Rey (2006) and Andersen and others (2007), for instance, show that there tends to be a negative correlation between equity returns and exchange rate returns in the data for several industrialized countries.
Statistics (IFS). Equity returns and equity-market capitalization are market indices and are sourced from Bloomberg whereas we took the REERs from the IFS and the Organization for Economic Cooperation and Development.4

Implementation of the Sign Restrictions

Before moving on to the empirical results, it is useful to explain how we implement the sign restrictions in our VAR. For a detailed explanation, we refer to Peersman (2003). Consider Equation (1). As the shocks are mutually orthogonal, \( E(\varepsilon_t \varepsilon_t) = I \), the Variance-Covariance of the reduce form residuals of Equation (1) is equal to \( \Omega = BB' \). For any possible orthogonal decomposition \( B \), we can find an infinite number of admissible decompositions of \( \Omega \), \( \Omega = BQQ'B' \), where \( Q \) is any orthonormal matrix, that is, \( QQ' = I \). Possible candidates for \( B \) are the Choleski factor of \( \Omega \) or the eigenvalue-eigenvector decomposition, \( \Omega = PDP' = BB' \), where \( P \) is a matrix of eigenvectors, \( D \) is a diagonal matrix with eigenvalues on the main diagonal, and \( B = PD^{1/2} \). Following Canova and De Nicolo (2002) and Peersman (2003), we start from the latter in our analysis. More specifically, \( P = \Pi_{m,n}Q_{m,n}(\theta) \) with \( Q_{m,n}(\theta) \) being rotation matrices of the form

\[
Q_{m,n}(\theta) = \begin{bmatrix}
1 & \cdots & 0 & \cdots & 0 & \cdots & 0 \\
\cdots & \ddots & \cdots & \ddots & \cdots & \cdots & \cdots \\
0 & \cdots & \cos(\theta) & \cdots & -\sin(\theta) & \cdots & 0 \\
\vdots & \vdots & \vdots & 1 & \vdots & \vdots & \vdots \\
0 & \cdots & \sin(\theta) & \cdots & \cos(\theta) & \cdots & 0 \\
\cdots & \cdots & \cdots & \cdots & \ddots & \cdots & \cdots \\
0 & \cdots & 0 & \cdots & 0 & \cdots & 1
\end{bmatrix}.
\] (2)

As we have six variables in our model, there are \( n(n-1)/2 = 15 \) bivariate rotations of different elements of the VAR: \( \theta = \theta_1, \ldots, \theta_{15} \), and rows \( m \) and \( n \) are rotated by the angle \( \theta_i \) in Equation (2). All possible rotations can be produced by varying the 15 parameters \( \theta_i \) in the range \([0, \pi]\). For the contemporaneous impact matrix determined by each point in the grid, \( B_j \), we generate the corresponding impulse responses

\[
R_{j,t+k} = A(L)^{-1}B_j \varepsilon_t.
\]

A sign restriction on the impulse response of variable \( p \) at lag \( k \) to a shock in \( q \) at time \( t \) is of the form

\[
R_{j,t+k}^{pq} \geq 0 \text{ or } R_{j,t+k}^{pq} \leq 0.
\]

We impose the sign restrictions for \( k = 4 \) lags; choosing a different length, however, does not alter the findings in a meaningful way. Following Uhlig (2005) and Peersman (2003), we use a Bayesian approach for estimation and

4Appendix Table A2 lists the variables and their definitions and sources.
inference. Our prior and posterior belong to the Normal-Wishart family for drawing error bands. Because there are an infinite number of admissible decompositions for each draw from the posterior when using sign restrictions, we use the following procedure. To draw the candidate truths from the posterior, we take a joint draw from the posterior for the usual unrestricted Normal-Wishart posterior for the VAR parameters as well as a uniform distribution for the rotation matrices, using 1,000 draws. We then construct impulse response functions. If all the imposed conditions of the impulse responses are satisfied, we keep the draw, whereas other decompositions are rejected. This means that these draws receive zero prior weight. Based on the draws kept, we calculate statistics and report the median responses, together with 84th and 16th percentile error bands.

II. Empirical Results

This section presents the empirical results from the structural VAR with sign restrictions, applied to the G-7 economies in the period 1974–2007. We also present various extensions to check for the sensitivity and robustness of the findings.

Benchmark Results

Figures 1–7 shows the impulse responses of the six variables, for each of the countries in our country sample of G-7 countries, to a 10 percent positive equity-market shock based on our Bayesian VAR model. The shaded areas indicate the 16 and 84 percentiles of the posterior distribution, following the convention in the literature. Table 2 summarizes the point estimates of the impulse responses at various time horizons.

As to the United States (Figure 1), a 10 percent increase in (relative) U.S. equity prices leads to a substantial worsening in the U.S. trade balance. The effect of the asset-price shock increases gradually over time up to 16–20 quarters, when it reduces the U.S. trade balances by 0.91 percentage points of U.S. GDP. This effect of asset prices on the trade balance appears to stem from two channels, a first one through wealth effects and a second related to the exchange rate. The importance of wealth effects is evident by the strong and quite persistent increase in private consumption, which in turn leads to a higher demand for imports.

The role of the exchange rate channel is underlined by the significant appreciation of the REER after a positive asset-price shock. The real appreciation is likely to be influenced both by the increase in domestic inflation and in domestic interest rates, though both of these responses are more short-lived as inflation and nominal interest rates revert back within 10 quarters. The rise in interest rates and real appreciation of the exchange rate is consistent with the evidence of the presence of a significant forward discount bias found in the literature (for example, Engel, 1996), as well as the more recent evidence stressing the importance of monetary policy or
“Taylor-rule” fundamentals for exchange rate determination (Engel and West, 2005; Clarida and Waldman, 2007).5

Figures 2–7 shows the corresponding impulse responses for the other G-7 countries of the sample. With a few exceptions, the patterns of the impulse responses are quite similar across countries: the trade balance of most countries deteriorates in response to a positive equity-price shock, though the permanence of this response is mostly somewhat lower than that of the United States. Moreover, the real exchange rate and private consumption always increases over the medium run after an increase in equity prices, though again the permanence of this effect differs markedly across countries. The strength of the reaction of private consumption for most countries

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Note: The figure shows the impulse responses of the six variables, for each of the countries in our sample of G-7 economies, to a 10 percent positive equity-market shock based on our Bayesian vector autoregression model. The shaded areas indicate the 16th and 84th percentiles of the posterior distribution.

Moreover, this positive effect of asset prices on the exchange rate is not necessarily inconsistent with the literature that finds a negative correlation between equity returns and exchange rate movements (Hau and Rey, 2006; Andersen and others, 2007) as those correlations are unconditional ones and may stem from other types of shocks.
suggests that wealth effects constitute an important channel through which asset-price shocks affect the trade balance of countries.

Nominal interest rates and inflation also rise in the short run, though recall that we imposed this response for the first four quarters in order to identify equity-price shocks. However, the magnitude and the persistence of the reaction of interest rates and inflation again differ substantially across countries. We also note and show the impulse responses for countries with somewhat puzzling results. For instance, the trade balance for the United Kingdom (Figure 2) improves in response to a positive domestic asset-price shock. We will return to a detailed discussion of these and other cross-country differences in the subsequent section.

Table 2 illustrates the heterogeneity of the point estimates at different time horizons, after one quarter, eight quarters, and 16 quarters, respectively. The table shows the marked differences in the impulse responses across countries, in the magnitude as well as in the dynamics and timing of the transmission of equity-price shocks. For instance, Italy’s trade balance appears to react relatively quickly to asset-price shocks, with the impulse response reverting back to zero relatively quickly. By contrast, the opposite is the case for the United Kingdom, where the reduction in the trade balance materializes only after several quarters.
Table 2 also nicely illustrates the different channels that are at play in transmitting the equity-price shock to the trade balance of countries. Countries that experience a stronger reaction of their trade balance to the asset-price shock also exhibit a larger response of their REER as well as private consumption. For instance, for the reaction after 16 quarters, it is in particular the United States but also Germany and Canada that see the strongest response of their trade balance, yet also experience a relatively larger sensitivity of private consumption and of their REER to the equity-price shock. Hence, this suggests that both a wealth effect on private consumption and an exchange rate channel are at play in explaining the transmission of an asset-price shock to a country’s trade balance.

Finally, the current financial market turmoil has further increased the focus on the role of monetary policy in addressing asset prices, and in particular asset-price bubbles. What do the impulse responses tell us about the potential role of monetary policy as a channel through which asset-price shocks may be transmitted to the trade balance of countries? In principle, one would expect that an aggressive tightening of monetary policy in response to a positive asset-price shock should dampen the effect of this shock on consumption and thus on net exports through the wealth channel. However, on the other hand, such a tightening may lead to an appreciation of the
exchange rate and a worsening of the trade balance. Based on the impulse responses in Figures 1–7, and the summary of these impulse responses shown in Table 2, there seems to be no clear-cut relationship between the response of interest rates, private consumption, inflation, and the trade balance across countries. This of course is no more than suggestive, and does not necessarily imply that monetary policy is not relevant for influencing the impact of asset prices on the trade balance. However, for instance for the United States these findings suggest that the reaction of U.S. short-term interest rates to asset-price shocks is not systematically lower than that of other industrialized countries.

**Robustness and Extensions**

How robust are these findings across alternative specifications, country samples, and time periods? We conduct several robustness tests on the benchmark model.

One important issue is how dependent our empirical findings are on the identification, that is, the sign restrictions we impose. Although these sign restrictions seem sensible, it is nevertheless useful to see how the results change when using alternative identification methods. We do so by
estimating our six-variable VAR using a Choleski decomposition. More precisely, we estimate the VAR using each possible ordering of the six endogenous variables, and then check the distribution of the resulting impulse response functions. Figure 8 shows the impulse responses of the trade balance to a positive asset-price shock for the case of the United States, the United Kingdom, Germany, and France. The top of the shaded area represents the maximum response coefficient among the different Choleski decompositions, whereas the lower end shows the minimum response at any time horizon. Overall, the findings suggest that the direction of the trade balance response to an asset-price shock is mostly the same when taking the Choleski decomposition as when using sign restrictions. However, the range of possible impulse responses is in several cases very large, underlining that the shape of the impulse responses is strongly dependent on the zero restrictions imposed on the Variance-Covariance matrix.

As a next step, we use alternative variables and variable definitions to check how sensitive the findings are to such changes. First, we use the

---

6We show here only the corresponding results for the United States, though the conclusions on the robustness checks are qualitatively similar for other countries.
current account instead of the trade balance, taking into account the fact that the dynamics of both can be considerably different for some countries. Figure 9 shows the impulse responses of this specification for the United States and confirms the basic thrust of the benchmark results as the current account declines considerably after a positive asset-price shock. In fact, the reaction of the current account is somewhat stronger, as one would indeed expect, likely due to the decline not only of the trade balance but also of the income part of the current account.

Second, we use relative equity-market capitalization, rather than equity prices, to define asset-price shocks. Figure 10 shows that the pattern of the impulse responses is unchanged for the United States (as well as for other industrialized countries, which are not shown for brevity reasons).

Note: See note to Figure 1.

Relative equity-market capitalization is measured as the difference in the log domestic-market capitalization and the log rest-of-the-world market capitalization, both measured in U.S. dollars. Using market exchange rates or PPP exchange rates does not change the findings in a meaningful way. More precisely, although the magnitude of the impulse responses may change depending on the specification, the direction and dynamics is very similar across specification.
Figure 7. Japan: Impulse Response Following an Asset-Price Shock

Note: See note to Figure 1.

Table 2. Impulse Response to a 10 Percent Domestic Asset-Price Shock

<table>
<thead>
<tr>
<th></th>
<th>United States</th>
<th>United Kingdom</th>
<th>Germany</th>
<th>France</th>
<th>Italy</th>
<th>Canada</th>
<th>Japan</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>One quarter</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Trade balance</td>
<td>−0.36</td>
<td>1.28</td>
<td>−0.09</td>
<td>−0.17</td>
<td>−0.70</td>
<td>−0.17</td>
<td>0.28</td>
</tr>
<tr>
<td>REER</td>
<td>3.12</td>
<td>2.72</td>
<td>−0.66</td>
<td>0.62</td>
<td>0.95</td>
<td>1.62</td>
<td>3.43</td>
</tr>
<tr>
<td>Consumption</td>
<td>4.88</td>
<td>4.37</td>
<td>0.80</td>
<td>4.34</td>
<td>4.57</td>
<td>3.57</td>
<td>4.06</td>
</tr>
<tr>
<td>Inflation</td>
<td>0.75</td>
<td>2.48</td>
<td>0.71</td>
<td>0.81</td>
<td>1.01</td>
<td>1.23</td>
<td>1.69</td>
</tr>
<tr>
<td>Interest rate</td>
<td>1.88</td>
<td>4.81</td>
<td>0.87</td>
<td>1.41</td>
<td>1.41</td>
<td>2.25</td>
<td>1.83</td>
</tr>
<tr>
<td>Equity market</td>
<td>10.00</td>
<td>10.00</td>
<td>10.00</td>
<td>10.00</td>
<td>10.00</td>
<td>10.00</td>
<td>10.00</td>
</tr>
<tr>
<td><strong>Eight quarters</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Trade balance</td>
<td>−0.66</td>
<td>0.28</td>
<td>−1.02</td>
<td>−0.58</td>
<td>−0.43</td>
<td>−0.04</td>
<td>0.21</td>
</tr>
<tr>
<td>REER</td>
<td>5.11</td>
<td>6.11</td>
<td>3.10</td>
<td>0.47</td>
<td>2.70</td>
<td>6.79</td>
<td>11.26</td>
</tr>
<tr>
<td>Consumption</td>
<td>7.09</td>
<td>5.19</td>
<td>9.31</td>
<td>0.75</td>
<td>2.94</td>
<td>9.59</td>
<td>13.13</td>
</tr>
<tr>
<td>Inflation</td>
<td>0.43</td>
<td>−1.25</td>
<td>0.24</td>
<td>0.29</td>
<td>0.67</td>
<td>0.06</td>
<td>0.38</td>
</tr>
<tr>
<td>Interest rate</td>
<td>0.69</td>
<td>−0.15</td>
<td>0.83</td>
<td>0.59</td>
<td>1.40</td>
<td>0.10</td>
<td>0.03</td>
</tr>
<tr>
<td>Asset prices</td>
<td>6.37</td>
<td>−0.70</td>
<td>0.70</td>
<td>−2.04</td>
<td>5.78</td>
<td>12.04</td>
<td>11.38</td>
</tr>
<tr>
<td><strong>16 quarters</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Trade balance</td>
<td>−0.91</td>
<td>−0.16</td>
<td>−0.87</td>
<td>−0.06</td>
<td>0.32</td>
<td>−0.61</td>
<td>−0.25</td>
</tr>
<tr>
<td>REER</td>
<td>4.64</td>
<td>1.99</td>
<td>1.47</td>
<td>−0.16</td>
<td>1.38</td>
<td>4.83</td>
<td>5.39</td>
</tr>
<tr>
<td>Consumption</td>
<td>7.29</td>
<td>−0.10</td>
<td>5.38</td>
<td>−0.31</td>
<td>1.01</td>
<td>6.74</td>
<td>7.99</td>
</tr>
<tr>
<td>Inflation</td>
<td>−0.03</td>
<td>−0.40</td>
<td>0.07</td>
<td>0.12</td>
<td>0.12</td>
<td>0.07</td>
<td>0.30</td>
</tr>
<tr>
<td>Interest rate</td>
<td>0.38</td>
<td>0.16</td>
<td>0.16</td>
<td>0.28</td>
<td>1.02</td>
<td>0.18</td>
<td>0.08</td>
</tr>
<tr>
<td>Asset prices</td>
<td>0.79</td>
<td>2.38</td>
<td>2.00</td>
<td>−0.16</td>
<td>1.63</td>
<td>10.86</td>
<td>−0.80</td>
</tr>
</tbody>
</table>
Finally, we shorten the time sample to 1990–2007 in order to allow for the possibility that asset-price shocks may have become more important over time as countries have become more integrated financially and through trade. Figure 11 shows that the initial reaction of the trade balance is slightly larger and the response of private consumption significantly larger for the United States, lending some support to this conjecture.

In summary, asset-price shocks appear to have a significant effect on the trade balance of countries, partly through wealth effects on domestic consumption and partly through an exchange rate channel that leads to real appreciation of the domestic currency. Moreover, there are substantial cross-country differences in the effect of equity-price shocks, with the trade balance of the United States in particular exhibiting one of the largest reactions to asset-price shocks.

III. Conclusions

The paper has analyzed the effect of asset-price shocks on the current account. Its focus has been on the experience of the cross-section of G-7

Figure 8. Impulse Response Following an Asset-Price Shock Based on Choleski Decomposition

Note: The figure shows the results following vector autoregressions using each possible ordering of the six endogenous variables and the corresponding application of a Cholesky decompositions for identifying equity-price shocks. The top of the shaded area represents to maximum response coefficient among the different Cholesky decompositions, whereas the lower end shows the minimum response at any time horizon.
industrialized countries. We have employed a Bayesian VAR with sign restrictions in order not only to motivate the identifying restrictions for asset-price shocks, but also to ensure that we can distinguish this type of shock from other shocks, such as to productivity, monetary policy, and government spending. The empirical evidence suggests that equity-price shocks indeed exert a significant effect on the trade balance of countries, partly through a wealth channel of private consumption and partly via an exchange rate channel.

One of the central findings of the paper is the substantial cross-country heterogeneity that we detect in the sensitivity of the trade balance to asset-price shocks. In particular the U.S. trade balance seems to be among the most sensitive to relative asset-price shocks, falling by 0.91 percentage points in response to a 10 percent increase in U.S. equity prices relative to the rest of the world. By contrast, other countries’ trade balances appear to be less responsive to asset-price shocks.

What explains this cross-country heterogeneity? Although the paper does not offer a systematic empirical analysis, the findings from the differences in the impulse responses suggest that both a channel via wealth effects as well as a real exchange rate channel appear to be at play in the transmission of asset-
price shocks to the current account of countries. Specifically, the evidence suggests that differences in these channels are important for understanding the cross-country heterogeneity in the sensitivity of countries trade balances to equity-price shocks.

Many open questions remain and there are various future avenues for better understanding the importance of asset-price shocks, both domestically and globally. In particular against the background of the financial market turmoil of 2007–08, the role of monetary policy for asset prices remains unclear. Similarly, the focus of the present paper has been only on equity markets. Extending the analysis to housing markets seems particularly relevant in the current financial market context. Another important avenue is to extend the analysis to emerging markets, which are rapidly becoming ever more important players in the global economy and international financial markets. We leave these avenues for future research.
Figure 11. United States: Impulse Response Following an Asset-Price Shock with Time Sample, 1990–2007

- Asset Prices
- Inflation
- Consumption
- Interest Rate
- Trade Balance
- Real Exchange Rate

Note: See note to Figure 1.

APPENDIX

Table A1. Country Sample

<table>
<thead>
<tr>
<th>Benchmark Sample</th>
</tr>
</thead>
<tbody>
<tr>
<td>Canada</td>
</tr>
<tr>
<td>France</td>
</tr>
<tr>
<td>Germany</td>
</tr>
<tr>
<td>Italy</td>
</tr>
<tr>
<td>Japan</td>
</tr>
<tr>
<td>United Kingdom</td>
</tr>
<tr>
<td>United States</td>
</tr>
</tbody>
</table>
Table A2. Data Definitions and Sources

<table>
<thead>
<tr>
<th>Variable</th>
<th>Definition</th>
<th>Source</th>
</tr>
</thead>
<tbody>
<tr>
<td>Asset prices</td>
<td>Difference between domestic equity returns and foreign equity returns, both measured in local currency terms</td>
<td>Bloomberg, market indices</td>
</tr>
<tr>
<td>Trade balance</td>
<td>Trade balance as a ratio to domestic GDP</td>
<td>IFS</td>
</tr>
<tr>
<td>Current account</td>
<td>Current account as a ratio to domestic GDP</td>
<td>IFS</td>
</tr>
<tr>
<td>Real effective exchange rate</td>
<td>Log real effective exchange rate using trade weights for a broad set of partner countries</td>
<td>IFS, OECD</td>
</tr>
<tr>
<td>Consumption</td>
<td>Difference in log private consumption in the domestic economy and log private consumption in the rest of the world, both expressed in U.S. dollars (using end-of-period exchange rates)</td>
<td>IFS</td>
</tr>
<tr>
<td>Inflation</td>
<td>Percent difference of domestic consumer price index inflation from that in the rest of the world</td>
<td>IFS</td>
</tr>
<tr>
<td>Interest rate</td>
<td>Percent difference of domestic short-term (money market) rates from those in the rest of the world</td>
<td>IFS, OECD</td>
</tr>
</tbody>
</table>


REFERENCES


Global Imbalances, Productivity Differentials, and Financial Integration

SUPARNA CHAKRABORTY and ROBERT DEKLE

This paper builds a two-country model with differential productivity and financial frictions to quantitatively account for the recent increase in the U.S. current account deficit. An influential literature says that as U.S. productivity surged, capital was attracted to the United States to take advantage of the high returns to investment. We show, however, that when we include emerging Asia, the gap in productivity growth between the United States and the rest of the world cannot explain the U.S. current account deficits, especially since 2000. This is because on a gross domestic product-weighted basis, the rest of the world actually had higher productivity growth during this period; and standard macroeconomic models would predict an outflow of funds from the United States to the rest of the world, and a consequent narrowing of the U.S. current account deficit. This paper shows that greater financial integration abroad can explain this anomaly. However, we still cannot explain why U.S. per capita output growth has been so low, despite the large inflow of capital. [JEL F32, F34, F36]

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The overall U.S. current account deficit rose from a modest $120 billion in 1996 (1.6 percent of gross domestic product) to $666 billion in 2005

*Suparna Chakraborty is an assistant professor at the Department of Economics and Finance, Baruch College, CUNY. Robert Dekle is a professor at the Department of Economics, University of Southern California. The authors thank the participants in the 2007 Econometric Society Summer Meetings, the Conference on Current Account Sustainability in Major Advanced Countries at the University of Wisconsin, and especially, Nelson Mark, the discussant for helpful comments.

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(6.06 percent of GDP) (Figure 1a). The 1990s were a period of almost continuous widening of the U.S. deficit (except for a brief period during 1994 to 1997), with the current account deficit rising especially rapidly since 2000.

In Figure 1b, we decompose the U.S. current account deficit into that with Europe, Japan, emerging Asia, and the Middle East. Although because of data limitations, the decomposition that separates emerging Asia and the Middle East from the remaining countries can only be performed after 1999, it is evident from Figure 1b that in the last few years, the deficit with emerging Asia is the most rapidly growing component of the U.S. current account deficit (by 2005, accounting for about 30 percent of the total U.S. deficit).

Perhaps the most influential explanation of this phenomenon of widening U.S. current account deficits is that of widening productivity gaps between the United States and the rest of the world. Since the mid-1990s, the U.S. economy experienced a productivity surge—and a rise in real returns to capital—while productivity in Europe and Japan stagnated. IMF (2005), Hunt and Rebucci (2005), and Engle and Rogers (2006) attribute the widening U.S. current account deficits to funds from the Europe and Japan seeking higher returns in the United States. A rise in U.S. productivity relative to the world will raise U.S. investment and consumption, and increase the U.S. current account deficit.

This “productivity gap” view, however, cannot explain the rising U.S. deficits since 2000. Later in the paper we will present our calculations of total factor productivity (TFP) growth for the rest of the world, when that region includes (1) only Europe and Japan; and (2) Europe, Japan, and emerging Asia for different subperiods since 1980. While it is true that U.S. productivity growth has outstripped that of the world between 1991 and 2000, since 2000, U.S. productivity growth has lagged by a large margin, the productivity growth in the world, including that of emerging Asia. This is mainly because of very rapid TFP growth in emerging Asia (over 5 percent), particularly in China (Dekle and Vandenbroucke, 2006), and of the growing economic weight of emerging Asia in the world (Figure 1c). If the differential productivity view is correct, then the United States should have been running current account surpluses (or, at the least, experience a decline in the deficits), and funds should be flowing out from the United States in the 2000s, when it fact the opposite happened.

This paper provides an accounting of why the United States ran current account deficits, despite higher productivity growth in the rest of the world in the post-2000 period. Using an explicit, two-region dynamic stochastic general equilibrium model, we attribute U.S. current account deficits to changes in the “cost” of buying U.S. assets in the rest of the world. We

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1See Dekle, Eaton, and Kortum (2006) for the role of China and emerging Asia in explaining U.S. current account deficits; and how much exchange rates need to adjust to equilibrate the U.S. current account deficits.
interpret this “cost” as relating to regulatory and other changes in the global financial sector, and find that our model-derived “cost” of purchasing U.S. assets correlates closely during our period of analysis (1980–2003) with other

Source: U.S. Bureau of Economic Analysis.

Note: Depicts the current account of the United States as a share of GDP during the period 1980 to 2006. Note that for our empirical analysis, we will concentrate on the period 1980 to 2003, because of data availability.

Source: U.S. Bureau of Economic Analysis (BEA).

Note: Plots the share of the United States current account in GDP with its partners, including a set of individual countries or group of countries and the rest of the world. One limitation of the data is that we do not have the U.S. current account balance data with emerging Asia, including China, before 1999. Nevertheless, we do note that the U.S. current account deficit with emerging Asia, particularly China, as a share of U.S. output increases between 1999 and 2005, while that with Japan, the major source of U.S. funds in the 1980s and 1990s declines.
well-known measures of global financial deregulation (Chinn and Ito, 2005); and with distinct, identifiable episodes of global financial shocks, such as the burst of the Japanese “bubble” in the early 1990s, and the Asian currency crisis of the late 1990s.

This paper extends Chakraborty and Dekle (2008) by introducing adjustment costs to changing physical capital, and a government sector. These modifications allow for richer dynamics, and may help us resolve another puzzle: why U.S. per capita output growth has been so low, despite the massive inflow of capital. We find that our extended model does a better job of matching the data on U.S. output per capita than in Chakraborty and Dekle (2008), while accounting for the increase in the U.S. current account deficit. However our extended model still overpredicts the growth in U.S. per capita output as compared to the data. This leads us to conclude that while adjustment costs of capital investment and wasteful government spending are part of the explanation for low output growth, there are other reasons that need to be investigated if we are to solve the puzzle of why U.S. per capita output growth has been so low.

Our benchmark model is a standard two-region business cycle model, where regions trade in final goods and international assets, but not in factors of production. In contrast to standard imperfect asset market models as in Baxter and Crucini (1995), we assume that international asset trading is

Figure 1c. Share of a Country’s (except the United States) GDP in the Aggregate GDP of the Rest of the World

Source: Authors’ calculations.
Note: Plots the changing weights of the countries in the rest of the world except the United States. We measure the weights as the ratio of GDP of an individual country to the aggregate GDP of the rest of the world where the rest of the world comprises the EU-15, Japan, and emerging Asia in our sample. Note that over the last two decades, and particularly since 2000, emerging Asia’s (including China) share has increased. Japan was a major force up to 1991; however its weight has since then diminished.
costly. These costs can be interpreted as monitoring costs or administrative costs associated with international lending. Introducing costs to asset trading helps capture the evolution of financial markets in a simple way. Specifically as a region’s financial markets evolve, and there is a move towards increased financial integration, these trading costs decline. Regions are ex ante symmetric, although ex post they may differ in their productivity shocks and openness of financial markets.

In our quantitative analysis, the two regions represent the United States and the rest of the world. To this, we introduce investment frictions by including a quadratic adjustment cost for investment in physical capital, and a government sector that taxes the representative agent and also provides transfers. Our intuition is that while increases in global financial integration have led to an influx of funds to the United States, thereby leading to increasing deficits, the funds have not been fully utilized for production due to investment frictions embedded in the adjustment cost of capital. At the same time, taxation and wasteful government expenditures have also hampered the full channeling of capital inflows for productive purposes. These twin factors kept the United States from fully realizing the gains from increased financial integration.

Impulse responses using our model show that when there is a positive productivity shock in the rest of the world, capital flows abroad, and the U.S. current account deficit declines. We show, however, that this effect is offset when there is a negative shock to the cost of buying U.S. assets. The U.S. current account can expand if foreigners find that it is easier to invest in the United States.

The idea that changes in overseas—especially emerging Asian—financing behavior can be related to the expansion of the U.S. current account was first floated by Dooley, Folkerts-Landau, and Garber (2004). Mendoza, Quadrini, and Ríos-Rull (2007), and Caballero, Farhi, and Gourinchas (2006) have both formally modeled this idea in a general equilibrium setup. However, in contrast to our work, these papers are all theoretical, and do not perform the empirical exercise that we do here for the first time, taking the model to the data and depicting how a reasonable pattern of changes in global financial structure—amended to a standard two-country macroeconomic model—can quantitatively account for the evolution of the U.S. current account, while not compromising the model’s ability to match other real macroeconomic aggregates.

2Other papers using dynamic optimizing frameworks to analyze the U.S. current account include Cavallo and Tille (2006), and Faruqee and others (2005).

3Choi, Mark, and Sul (2008) also use an imperfect asset market model to trace the actual time path of the U.S. current account. Like us, they too find that the time path of the U.S. current account cannot be explained by U.S. and “rest of the world” productivity differences. The authors attribute the unexplained portion of the U.S. current account to a declining pure rate of time preference of foreign residents.
I. The Model

Our model builds on the incomplete financial markets framework developed by Baxter and Crucini (1995). In their model, instead of a full set of contingent financial contracts, there is only one internationally traded asset, a bond, which can be freely bought and sold. In our setup, we introduce costs to international lending or to the buying and selling of foreign bonds, which are intended to capture various frictions arising from the lack of liberalization in international financial transactions, such restrictions on lending by foreign banks or taxes on the purchases of U.S. bonds and assets.

Time is discrete, indexed by \( t = 0, 1, 2 \ldots \) and the time horizon is infinite. The world is comprised of two countries, Home and Foreign, indexed by \( i \in (H, F) \), each of which is populated by a unit measure of identical, infinitely lived households.\(^4\) In addition to households, each country is also populated by an infinite number of perfectly competitive firms that own the production technology. There is only one good in our model, produced by the firms in each country, using country-specific capital and labor. Once produced, the good is then traded between the two countries, and is used for consumption and investment.

The goal of the paper is to quantitatively study the impact of differential productivity growth on global imbalances in an environment of increasing asset market integration. To this end, we introduce costs to international lending, or to the purchase of the other country’s bond, where costs reflect the relative difficulty of international financial market access. Costs evolve according to changes in domestic and international regulations, with a lowering of costs reflecting increasing ease of access to global financial markets.

Uncertainty in our model arises from country-specific productivity shocks, and shocks to the costs of international lending. The countries are ex ante perfectly symmetric.

Preferences and Technologies

Households in country \( i \) maximize expected discounted utility over consumption \( c_{it} \), and leisure \( 1 - l_{it} \):

\[
E \sum_{t=0}^{\infty} \beta^t u(c_{it}, l_{it}).
\]  

(1)

The budget constraint is given by:

\[
c_{it} + x_{it} + m(k_{it}, k_{it-1}) + s_{it} + f(s_{it}, \phi_{it})s_{it} \\
\leq w_{it}l_{it} + r_{it}k_{it-1} + R_{it}s_{it-1} - T_{it},
\]  

(2)

where \( w_{it}l_{it} \) denotes labor income, and the return to capital is given by \( r_{it} k_{it-1} \).

In addition, households earn returns on international lending where \( R_{it} \)

\(^4\) The assumption of equal population size is for simplicity. Relaxing the assumption would modify the resource constraints. Our main findings would not change.
represents the gross world interest rate at time $t$. $T_{it}$ denotes lump sum taxes, and $m(k_{it}, k_{it-1})$ denotes adjustment costs associated with investment in the physical capital stock. We use a quadratic specification for adjustment costs in capital such that:

$$m(k_{it}, k_{it-1}) = \frac{\chi}{2} \left( \frac{k_{it} - \zeta k_{it-1}}{k_{it-1}} \right)^2 k_{it-1}. \quad (3)$$

Apart from representative consumers and firms, we include the government in our extended model that balances its budget every period such that:

$$g_{it} = T_{it}. \quad (4)$$

Income is used to finance consumption $c_{it}$, physical capital investment, $x_{it}$, and for international net lending $s_{it} - s_{it-1}$. International net lending, or the purchasing (or selling) of the other country’s bonds involves a cost, where $f(s_{it}, \phi_{it})$ denotes the cost of purchasing bonds in terms of the final good. If buying or selling bonds is frictionless, as in the usual Baxter and Crucini (1995) model, then $f(s_{it}, \phi_{it}) = 0$.

We take $f(s_{it}, \phi_{it})$ to encompass all the impediments and frictions that the foreign country may have in lending to the home country. These impediments and frictions may include constraints on international lending such as foreign exchange controls, or bank minimum capital requirements, which restrict overseas bank lending. They may also include frictions arising from asymmetric information that impair the ability of foreign lenders to costlessly acquire information about the behavior of home firms (the borrower). In many models, these asymmetric information problems will lead to additional monitoring costs for the foreign lender, raising the costs of lending (Bernanke, Gertler, and Gilchrist, 1995).

For our purposes, we assume unit costs that are constant with respect to the amount of overseas lending, or the purchase of home country bonds, $s_{it}$, but vary over time:

$$f(s_{it}, \phi_{it}) = \phi_{it}. \quad (5)$$

$\phi_{it}$ captures the impact of external factors that affect the cost of purchasing bonds or of lending internationally. Financial liberalization is captured by a decline in $\phi_{it}$ that allows foreign households to lend more, or to purchase more home country bonds, at lower costs.

The perfectly competitive firms own a production technology that combines labor $l_{it}$ and capital $k_{it-1}$ to produce the traded good $y_{it}$:

$$y_{it} \leq F(A_{it}, k_{it-1}, l_{it}). \quad (6)$$

---

We assume a balanced budget for the government. An alternative way to do our exercise would be to allow for budget deficits, and also allow government to trade in bonds along with households.
$A_t$ represents productivity that is exogenously determined. Ex post, the home and the foreign countries differ with respect to country-specific productivity shocks. A country experiencing a negative productivity shock will divert funds to investment opportunities outside the country, and thus accumulate the other country’s bonds; the country experiencing a positive productivity shock would borrow funds from the rest of the world for investment, and consequently would run current account deficits.

Every period, the firm maximizes profits $\pi_{it}$ subject to the production technology summarized in Equation (6), where profits are given by:

$$\pi_{it} = y_{it} - w_{it}l_{it} - r_{it}k_{it-1}. \quad (7)$$

The world goods market clearing condition requires that aggregate consumption and investment in the world be less than or equal to world production of the traded good:

$$\sum_{i=H}^{F} (c_{it} + x_{it} + m(k_{it}, k_{it-1}) + g_{it}) \leq \sum_{i=H}^{F} y_{it}. \quad (8)$$

Bond market clearing requires that the aggregate bond holdings in the world every period is zero:

$$\sum_{i=H}^{F} s_{it} = 0. \quad (9)$$

We are interested primarily in the dynamics of the current account in our model. The current account in country $i$ is defined as the sum of the trade balance and net interest payments. Given our model, the current account is given by:

$$y_{it} - c_{it} - x_{it} - g_{it} - (R_t - 1)s_{it-1}, \quad (10)$$

where $y_{it} - c_{it} - g_{it} - x_{it}$ is the trade balance, and $(R_t - 1)s_{it-1}$ is the net interest earnings on international lending (bonds).

Given the budget constraint summarized in Equation (2) and the definition of the current account as summarized in Equation (10), we can express the current account in country $i$ as:

$$s_{it} + \phi_{it} \times s_{it} - s_{it-1}, \quad (11)$$

which is of course identical to the net change in international lending in the foreign country, or borrowing by the home country, after adjusting for borrowing costs.

Equilibrium

An equilibrium in our model is given by a vector of allocations $\{c_{it}, l_{it}, x_{it}, k_{it}, y_{it}, s_{it}\}_{i=H}^{F}$, and a vector of prices $\{w_{it}, r_{it}, R_t\}_{i=H}^{F}$, $i = H, F$ such that given a state of the economy summarized by $\{k_{it-1}, s_{it-1}\}_{i=H}^{F}$, $i = H, F$, and exogenous shocks to productivity and cost of

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lending \( \{A_{it}, \phi_{it}\}_{t=0}^{\infty} \) \( i = H, F \), the allocations and the prices solve Equation (1) the representative household’s utility maximization problem (summarized in equations (1) to (5)) and Equation (2) the firm’s profit maximization problem (summarized in equations (6) and (7)), Equation (3) the world resource constraint is satisfied (summarized in equation (8)), and Equation (4) the bond market clearing condition is satisfied (summarized in equation (9)).

Note that balanced growth in our model assumes that the long-run rate of technological progress, denoted in our model by \( \gamma \), in the two countries are identical.

II. Model Predictions

Ex ante we have two equally sized countries, \( H \) and \( F \) that are symmetric in every respect. Ex post we allow them to vary with respect to productivity, \( A_{it} \) and bond trading costs \( \phi_{it} \). Taking the United States as the home country, and the rest of the world as the foreign country, we make the reasonable assumption that the United States and the rest of the world are equally sized, since U.S. GDP is about one-third of world GDP.

Parameters

The utility function is assumed to be quasi-linear (often referred to as GHH after Greenwood, Hercowitz, and Huffman, 1988) where:

\[
 u(c_{it}, l_{it}) = \left(\frac{c_{it} - \psi l_{it}^\gamma (1 + \gamma)^{1-\sigma}}{1 - \sigma}\right)^{1/\gamma}.
\] (12)

The GHH preferences are widely used in the international macroeconomics literature (Mendoza, 2006) as it better fits certain international business cycle facts as compared to Cobb-Douglas preferences.

The production function has a labor-augmenting Cobb-Douglas form:

\[
 F(A_{it}, k_{it-1}, l_{it}) = k_{it-1}^{\theta_i} (A_{it} l_{it})^{1-\theta_i}
\] (13)

The key parameters are summarized in Table 1. We choose the share of capital to be 33 percent of GDP in both countries yielding \( \theta_i = 0.33 \), \( i \in \{ H, F \} \). Capital is assumed to depreciate at an annual rate of 10 percent so that \( \delta_i = 0.10 \), \( i \in \{ H, F \} \). These parameters are taken from Backus, Kehoe, and Kydland (1992). We assume \( \sigma = 2 \). The long-term growth rate of technological progress or \( \gamma \) is taken as 2 percent, and the steady state rate of interest is taken as 6 percent. Further, assuming that in the steady state, households in both countries choose to allocate 33 percent of their time to work (from Backus, Kehoe, and Kydland, 1992), and normalizing the steady state output to be 1, we calculate \( \psi = 3.95 \), given \( \psi = 1.6 \) (from Mendoza, 2006).

\( \phi_{it} \) embodies intangibles like information about domestic and international policies regarding a country’s access to world financial markets. We back out the steady state value of \( \phi_{it} \) from the budget constraint, and the first-order conditions of the model, shown below. To do this, we need steady state values of the capital output \( k_{it}/y_{it} \) ratio, and the net

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The capital output ratio is taken as the average long-run capital to output ratio of the U.S. economy. The net lending (or U.S. borrowing) to output ratio is estimated to match the average ratio of net current account balances of the United States to its output, over the period 1980 to 2003. This yields the net lending to output ratio of 0.12 or gross lending to output ratio $s_i / y_i$ of 0.61. Given these parameters, the steady state capital output ratio, $\phi_{it}$ is calculated to be 0.0127. Given our steady state capital output ratio, we calculate $\beta_l = 0.99$. The quadratic adjustment cost depends on two parameters, $\chi$ and $\zeta$. As in the literature, we assume that in the steady state, the adjustment cost is zero. We can ensure this by setting $\zeta = 1.02$ to match the long-term growth rate of the U.S. economy. Note that in the literature where no balanced growth is assumed, $\zeta$ is set to 1. We measure $\chi$ to be 0.01 to match the volatility of investment in the United States.

### Qualitative Results

#### Solution Algorithm

Given the functional specifications, the necessary first order conditions of our model for country $i \in \{H, F\}$ are given by:

$$ \left( c_{it} - \frac{\psi}{\psi_{it}} F_i^p \right)^{-\sigma} - \lambda_{it} = 0, $$

$$ (14) $$

---

**Table 1. Parameter Specifications**

<table>
<thead>
<tr>
<th>Parameter Description</th>
<th>Parameter Symbols</th>
<th>Parameter Values</th>
</tr>
</thead>
<tbody>
<tr>
<td>Growth rate</td>
<td>$\gamma$</td>
<td>2%</td>
</tr>
<tr>
<td>Preferences</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Coefficient of risk aversion</td>
<td>$\sigma$</td>
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</tr>
<tr>
<td>Discount factor</td>
<td>$\beta$</td>
<td>0.99</td>
</tr>
<tr>
<td>Leisure weight</td>
<td>$\psi$</td>
<td>3.95</td>
</tr>
<tr>
<td>Depreciation rate</td>
<td>$\delta$</td>
<td>0.1</td>
</tr>
<tr>
<td>Income shares</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Capital</td>
<td>$\theta$</td>
<td>0.33</td>
</tr>
<tr>
<td>Labor</td>
<td>$1 - \theta$</td>
<td>0.67</td>
</tr>
<tr>
<td>Adjustment cost</td>
<td>$\zeta$</td>
<td>0.02</td>
</tr>
<tr>
<td></td>
<td>$\chi$</td>
<td>0.01</td>
</tr>
</tbody>
</table>

Note: The parameters are calibrated from the necessary first-order conditions and the steady state values of the variables. We assume that both countries are ex ante symmetric, so we take the parameter values to be the same in both countries.
\( (1 - 0)\gamma_{it} - \psi_{it} = 0, \) \hspace{1cm} (15)

\[
\left( \lambda_{it+1} \left( \frac{\gamma_{it+1}}{\kappa_{it+1}} + 1 - \delta_{it} + \frac{1}{2} (1 + \gamma)^2 \left( \frac{k_{it+1}}{x_{it}} \right)^2 - \frac{1}{2} \xi_{it} \right) \right.
\]

\[
- \frac{(1 + \gamma)^{\alpha_{it}}}{\beta_{it}} \lambda_{it} \left( (1 + \gamma) \frac{k_{it}}{x_{it}} - \chi \xi_{it} \right) = 0
\]

\( \lambda_{it+1} R_{t+1} - \frac{(1 + \gamma)^{\alpha_{it}}}{\beta_{it}} \lambda_{it} (1 + \phi_{it}) = 0, \) \hspace{1cm} (17)

where \( \lambda_{it} \beta_{it} (1 + \gamma)^{(1 - \sigma)} \) is the shadow price of consumption; and the variables are detrended by their long-term growth rates. Equations (14) and (15) are standard. Equation (14) equates the marginal utility of consumption to its shadow price, and Equation (15) equates the marginal rate of substitution between consumption and leisure to the marginal product of labor. Note that under quasi-linear (or GHH) preferences, the marginal rate of substitution between consumption and leisure is independent of consumption, making labor choice immune to wealth effects. Equations (16) and (17) are the intertemporal conditions for investment in capital, and for international net lending. For our analysis, we assume that government expenditure is exogenous. The steady state share of government expenditure is taken as 20 percent (the average for the period 1980 to 2003).

We solve our model using the technique of log-linearization. To that end, we first need to specify the stochastic processes underlying our exogenous variables.

The stochastic processes are vector autoregressive processes of order one and are given by:

\[
\begin{bmatrix}
\bar{A}_{Ht} \\
\bar{A}_{Ft} \\
\bar{\phi}_{Ht} \\
\bar{\phi}_{Ft} \\
\bar{g}_{Ht} \\
\bar{g}_{Ft}
\end{bmatrix}
= P
\begin{bmatrix}
\bar{A}_{Ht-1} \\
\bar{A}_{Ft-1} \\
\bar{\phi}_{Ht-1} \\
\bar{\phi}_{Ft-1} \\
\bar{g}_{Ht-1} \\
\bar{g}_{Ft-1}
\end{bmatrix}
+ \begin{bmatrix}
\bar{e}_{AHt} \\
\bar{e}_{AFt} \\
\bar{e}_{Ht} \\
\bar{e}_{Ft} \\
\bar{e}_{gHt} \\
\bar{e}_{gFt}
\end{bmatrix}
\] \hspace{1cm} (18)

where \( \bar{A}_{it} \) and \( \bar{g}_{it} \) denote the log deviation of productivity and government expenditures from their steady state. \( \bar{\phi}_{it} \) denotes the deviation of the cost of bond trading from its steady state. Epsilon or the error terms capture the shocks. For our numerical analysis, we assume that government expenditures are uncorrelated across the two countries, and with other variables. We estimate the parameters driving government expenditures from our data on U.S. government spending, and report the parameters in Table 2. \( P \) is a 6 x 6 matrix that summarizes the parameters underlying the stochastic process. The innovations are serially independent, multivariate
normal random variables. The variance-covariance matrix of the innovations is summarized by another 6x6 matrix that we call $Q$. We initially assume no contemporaneous correlation of the innovations in the two countries.\footnote{We could not detect any statistically significant correlation between our calculated “foreign” (European Union and Japan; and EU, Japan and emerging Asia, respectively) and “home” (U.S.) TFPs. Despite this lack of correlation in the productivity shocks between the two regions, as a robustness check of our basic results presented in the paper, we performed impulse response analysis, imposing some spillover of shocks between the regions (contemporaneous correlation of shocks of 0.05). None of the impulse responses were affected.}

For our analysis, we take the parameters determining the evolution of productivity from Kehoe and Perri (2002). While we have backed out the steady state value of $\phi_i$, the cost of international lending, the stochastic process driving $\phi_i$ over time is unknown. To this end, we make two assumptions regarding the evolution of $\phi_i$: (1) $\phi_i$ is not correlated with $A_{it}$, and (2) $\phi_i$ between countries are not correlated.

Given that $\phi_i$ captures policy-related and other external effects on the cost of lending and the purchase of bonds, we believe that these assumptions are realistic. As for the persistence of the initial shock to $\phi_i$, we experiment with a very high (0.91), as well as very low persistence (0.5). In addition, we assume that the variance of $\phi_i$ is low.

In any period $t$, given the state of the economy summarized by the vector $\{k_{it-1}, s_{it-1}\}_{i=H,F}$, and the realization of the exogenous shocks summarized by the vector $\{A_{it}, \phi_{it}, g_{it}\}_{i=H,F}$, the numerical solution to our model expresses the endogenous control and state variables as functions of the state and the exogenous variables, where the coefficients of the functions depend on the parameters underlying the stochastic processes defined in Equation (18).

<table>
<thead>
<tr>
<th>Table 2. Stochastic Process of Exogenous Shocks</th>
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<tbody>
<tr>
<td>( P = )</td>
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<tr>
<td></td>
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<tr>
<td>( Q = )</td>
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<td></td>
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<td></td>
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<tr>
<td></td>
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</tbody>
</table>

Note: We assume the shocks to be uncorrelated across countries and with each other.
Impulse Responses

Before carrying out the quantitative exercises, that is, apply our model to the actual data, we examine the qualitative properties of our model by performing impulse response exercises. As mentioned, our research is motivated by observed changes in the United States and the rest of the world current accounts between the 1980s and today.

There are two important trends: (1) during the 1980s, productivity growth in the rest of the world, particularly in the European countries, and in Japan exceeded productivity growth in the United States, but the trend has reversed since the 1990s; and (2) the 1980s and the 1990s were also periods of gradual financial liberalization in the European countries, in Japan, and particularly in East Asia. In our model, we assume that country $F$ or foreign is the rest of the world, and country $H$ or home is the United States.

While tracing the impulse responses, we are interested in the impact of three events: (1) a 1 percent positive shock to the productivity of the home country ($A_H$); (2) a 1 percent negative shock to the productivity of the foreign country ($A_F$); and (3) a 1 percent reduction in the cost international lending, or of international bond purchases, measured by a 1 percent decrease in $\phi_F$.

We are primarily interested in the impact of these three shocks on the current account of the United States that is, of the home country $H$. Note that according to Equation (10), the amount of borrowing in any period determines the current account, with an increase in bond purchases by country $F$ increasing the current account deficit of country $H$.

Therefore for our impulse responses, we chart the effect of shocks on U.S. borrowing. We also trace the impact on real macroeconomic aggregates like output and consumption per capita.

In addition to concerns about ballooning current account deficits in the United States, researchers have tried to reconcile the high U.S. current account deficits with the low levels of the global real interest rate. Typically, we would expect that the excess of investment over saving (the current account deficit) in the United States would lead to higher global real interest rates. However, in reality, U.S. real interest rates have fallen from about 4.5 percent in 2000 to about 1.8 percent in 2005.

We plot the transition dynamics of our model in Figures 2a to 2c. Output in the home country increases under all three shocks. A fall in productivity in rest of the world lowers investment incentives in the foreign country leading to capital flows to the home country or the United States, and an increase in investment and output in the United States. A similar effect occurs when productivity in the home country increases. This attracts foreign capital flows into the U.S. market, and increases the inflow of funds to the United States, thereby worsening the current account deficit.

A fall in the rest of the world (synonymous with financial liberalization in our model) sharply raises home (U.S.) output. A decline in the cost of lending makes it easier for foreign investors to invest in the U.S., thereby increasing capital inflows to the U.S. and the U.S. current account deficit expands sharply.
Our impulse responses also show increased consumption in the home country, thus supporting the conclusions reached by Mendoza, Quadrini, and Rios-Rull (2007) that financial globalization has welfare enhancing effects for the borrower.

As for interest rate implications, shocks to productivities imply an increase in global real interest rates, which is at odds with the data. However, we do find that a 1 percent decline in $\phi_{Fr}$, the cost of lending, generates a decline in global interest rates, along with an increase in the flow of funds to the home country (United States). This helps reconcile why global real rates have fallen, despite an increase in U.S. international borrowing. Increased financial liberalization in the rest of the world raised the supply of funds, and despite higher productivity growth in the United States, lowered global interest rates.

One of our aims in this extended version was to improve our model’s performance with respect to the data. If we compare the impulse responses traced here with that of Chakraborty and Dekle (2008), the addition of adjustment costs to capital to our model increases the magnitude of the

![Figure 2a. Impulse Responses to a 1 Percent Positive Shock to Home Productivity, $A_H$](image)

Source: Authors’ calculations.

Note: Variable description: The variables of interest are borrowings by the home country, borrowH (or equivalently lending by the foreign country); the real interest rate on international lending, Real interest; consumption of the representative agents of the home country that is the recipient of international funds, consH; and the output of the home country, outputH.
response of U.S. borrowing, but lowers the response of real macroeconomic aggregates to financial liberalization, which is closer to the data.

Below, we perform a quantitative exercise to see if given the paths of home and foreign TFP shocks, we can explain the path of the U.S. current account. We show that TFP shocks alone cannot explain U.S. current account deficits. Another factor, say changes in the ease of investing in U.S. assets, is necessary to explain the evolution of U.S. current account imbalances. We also show that including adjustment costs to capital, and wasteful government spending can help explain why U.S. per capita output growth has been so low, despite the increased access to global funds.

III. Quantitative Application to the United States and Rest of the World

In bringing our stylized two-country model to the data, we need to define what the rest of the world stands for. In our two-country model, country $H$ or home represents the United States. For country $F$ or the foreign country representing the rest of the world, we consider two alternatives: (1) the rest of
the world comprises the EU-15 and Japan; and (2) the rest of the world comprises the EU-15, Japan and emerging Asia, particularly China. Especially since the late 1990s, Asia (except Japan) has emerged as a major global player in world financial markets, not only because of rapid growth, but also because of the relaxation of capital outflow controls, and the purchase of U.S. bonds as foreign exchange reserves in emerging Asia, especially in China.

Using the two-country model developed above, we try to match the current accounts and the GDPs per capita of the home country (United States) from 1980 to 2003. We show that while differences in TFP shocks do a good job of explaining the U.S. current account deficits in the late 1980s and the early 1990s, they do a poor job in explaining the U.S. current account deficits later on, particularly since 2000.

What is required for our model to explain the U.S. current account is for financial frictions to behave in a nonmonotonic fashion. That is, the ease of investing in the United States has to start declining from about 1985 to about 1996, and then rise from 1997 to 2000, and then sharply decline.

Source: Authors’ calculations.

Note: Variable description: The variables of interest are borrowings by the home country, borrowH (or equivalently lending by the foreign country); the real interest rate on international lending, Real interest; consumption of the representative agents of the home country that is the recipient of international funds, consH; and the output of the home country, outputH.
The decline in financial frictions starting in 1985 corresponds to the financial liberalization in Europe and Japan, including the “Big Bang” reforms at the start of Japan’s “bubble economy.” As domestic stock and land prices rose, Japan used the rising domestic asset prices to borrow, and invest in the United States. In other words, the rise in domestic asset prices mitigated the financing constraints of Japanese firms, enabling them to lend to the United States.

The increase in our measure of financial frictions from 1997 to 2000 is related to the Asian financial and Japanese banking crisis, limiting the ability of emerging Asian countries and Japan to lend to the United States. We show that except for the Asian financial crisis years from 1997 to 2000, the decline in financing frictions (ϕ_F) from 1985 to 2003 is close to monotonic.

Although with nonmonotonic financing frictions, we are able to describe the changes in the U.S. current account, we are still unable to account for the slow growth in U.S. per capita GDP, even in our extended model, with capital adjustment costs and “wasteful” government spending. The growth in per capita GDP simulated by our extended model considerably overshoots the actual growth in GDP per capita in the United States. That is, given the rapid inflow of capital into the United States, U.S. per capita GDP between 1985 and 1991, and between 2001 and 2003 should have been growing much faster than what is observed in the data.

The Data

The data are described in detail in the Data Appendix. We assume that the two regions, home and foreign, are equal sized in the steady state, which we take to be 1980, but the two regions can subsequently diverge in size. As mentioned, for the foreign country representing the rest of the world, we take two sets of countries: (1) Japan and the EU-15; and (2) Japan, the EU-15, and emerging Asia. In our analysis, TFP is measured as a Solow residual where:

\[ A_t = \frac{y_{it}}{k_{i,t-1}^{0.6}(1+\gamma)^{t-0.6}}. \]  

(19)

To calculate TFP, we first must calculate \( y_{it} \), \( l_{it} \), and \( k_{it-1} \) for our two sets of countries that we take as the rest of the world. As described in the Data Appendix, we construct the aggregate variables for the rest of the world by taking a weighted sum of each corresponding variable. For example, as for the GDP per capita \( y_{it} \) of the home country, we take the weighted sum of the GDPs per capita of each of the EU-15 countries and Japan, where the weight for country \( i \) is the share of country \( i \)’s GDP in the sum of the EU-15+Japan’s GDPs. Since the inclusion of emerging Asia changes the set of countries and the associated weights, the estimated TFPs are different for the two sets of countries.

Thus calculated, we interpret changes in \( A_t \) as the deviation of TFP from its steady state, where in the steady state, global TFP is assumed to grow at 2
Figure 3 depicts our estimated $A'$s. The average TFP growth rates during our four subperiods are summarized in Table 3. The first subperiod (1980 to 1986) shows productivity growth in the EU-15 and Japan growing at 0.84 percent above trend on average, while that in the United States grows at 0.14 percent below trend. Similarly, during the next subperiod (1986 to 1991), productivity grows at 0.79 percent below trend in the United States, and 0.87 percent above trend in the EU-15 and Japan. This pattern changes during the third subperiod (1991 to 2000), when productivity in the EU-15 and Japan grows 1.35 percent below trend, while U.S. productivity grows 0.2 percent above trend. Productivity in the foreign country slightly improves in the last period (2000 to 2003), with TFP still growing slightly below trend at (-56 percent). In the home country (United States), productivity growth sharply picks up, with TFP growing at 0.62 percent above trend during 2000–03.

Following Choi, Mark, and Sul (2008), we also experimented with other detrending procedures, including the HP-filter. We find that using alternative measures of long-term trend growth rate like 1.5 percent or 2.15 percent do not alter our results. Nelson Mark pointed out that the results might be sensitive to HP-filtering. We find that while the pattern of TFP shocks changes with HP-filtering, the effects of the HP-filtered TFP shocks on the U.S. current account and other macroeconomic aggregates are essentially the same as the effects of our constant detrended TFP shocks (results available upon request from the authors).
These TFP growth patterns change when we include emerging Asia in our set of countries in the rest of the world, especially during the early 1980s, and during the 2000–03 period. In contrast to the previous case when only the EU-15 and Japan are in the foreign country, adding emerging Asia results in TFP growing at 0.24 percent below trend during 1980 to 1986. Since the mid-1980s, foreign country TFP growth including emerging Asia follows a cyclical pattern, growing above trend during the late 1980s, before declining below trend during the 1990s. It is remarkable that between 2000 and 2003, when emerging Asia is included in the foreign country, TFP growth is 3.9 percent, far above trend.

These patterns in the rest of the world and U.S. TFP growth suggest that if current accounts were affected by TFP growth alone, then in the 1980s and since 2000, U.S. investors should be lending to the rest of the world. We therefore should see a decline in the U.S. current account deficit in the 1980s and since 2000, when in fact the U.S. current account deficit rose. Thus, some other factor than differences in the home and foreign country TFPs must be affecting the U.S. current account. We attribute this factor to changes in the ease of buying U.S. financial assets, or financial liberalization in the foreign country.

### Quantitative Experiments

We conduct our experiments when the rest of the world includes the EU-15, Japan, and emerging Asia. We feed in the calculated annual changes in TFP in the foreign (EU-15 and Japan), and in the home (the United States) countries, assuming no change in financial liberalization. The results for the

<table>
<thead>
<tr>
<th>Year</th>
<th>Foreign (EU-15 + Japan)</th>
<th>Foreign (EU-15 + Japan + Asia)</th>
<th>Home (U.S.)</th>
</tr>
</thead>
<tbody>
<tr>
<td>1980:86</td>
<td>0.84%</td>
<td>−0.24%</td>
<td>−0.14%</td>
</tr>
<tr>
<td></td>
<td>−1.21</td>
<td>−3.06</td>
<td>−2.53</td>
</tr>
<tr>
<td>1986:91</td>
<td>0.87%</td>
<td>0.15%</td>
<td>−0.79%</td>
</tr>
<tr>
<td></td>
<td>−1.64</td>
<td>−1.57</td>
<td>−0.88</td>
</tr>
<tr>
<td>1991:2000</td>
<td>−1.35%</td>
<td>−1.18%</td>
<td>0.02%</td>
</tr>
<tr>
<td></td>
<td>−1.41</td>
<td>−1.88</td>
<td>−1.41</td>
</tr>
<tr>
<td>2000:03</td>
<td>−0.56%</td>
<td>3.88%</td>
<td>0.62%</td>
</tr>
<tr>
<td></td>
<td>−0.23</td>
<td>−3.67</td>
<td>−1.5</td>
</tr>
</tbody>
</table>

Note: Change in total factor productivity (TFP) measured as \( \frac{(A_{i,t+1} - A_{i,t})}{A_{i,t}} \) where \( A_{i,t} \) is measured as a Solow Residual. “Foreign” includes EU-15 and Japan or EU-15, Japan along with emerging Asia while “Home” represents the United States. For each subperiod, we provide the average growth rate of \( A_{i,t} \), and the standard deviations are in brackets.
U.S. current account are shown in Figure 4 (numerical comparison summarized in Table 4). For the sake of comparison, we also depict the predicted current account share from our earlier model (Chakraborty and Dekle, 2008a). The earlier model, which we call “Basic,” has no government sector and capital adjusts instantaneously. Introducing capital adjustment in our “Extended” model presented here allows for richer dynamics.

On the whole, the current account of the United States has been in deficit and the deficit increases from 1.36 percent of GDP during the 1980s to 4.35 percent of GDP between 2000 and 2003. Except for the 1980s, when our model predicts a U.S. current account surplus, the predictions from our model match up well with the data. In the 1980s, our model predicts a current account surplus, because TFP growth in the EU-15 and in Japan are high, while that in the United States is low, which should result in a flow of funds from the United States to the rest of the world, when in fact, the opposite happened. Note that the results for the current account are almost identical for the basic and extended models.

The fit of our model improves somewhat between 1995 and 1999, as the fall in TFP growth below trend in the rest of the world is higher. This narrowing of the negative TFP growth differential with the United States means that the model-simulated U.S. current account deficit increases between 1991 and 1999, corresponding more closely with the data. However, between 2000 and 2003, the fit of the model deteriorates, as the positive gap in the TFP growth differential between the rest of the world and the United States widens. Now given the very rapid TFP growth above trend in the foreign country (mainly China), the model predicts a current account surplus in the United States, when we actually have a U.S. current account deficit. Our quantitative results are robust to different levels of persistence in the productivity shocks. We raised the persistence from our benchmark (0.95) to a random walk (1.0), and also lowered it to a serially uncorrelated process, but the simulated U.S. current account changed very little.8

The Pattern in $\phi_F$

As noted above, with or without emerging Asia in the rest of the world, our model misses much of the action in the evolution of the U.S. current account when we assume no change in global financial integration. We thus conduct the following counterfactual experiment: What would be the pattern in the financial liberalization in the rest of the world, $\phi_F$, if we are to match the evolution of the U.S. current account? Does this derived pattern accord with the actual financial liberalization that has taken place?

The principle behind this accounting exercise is related to the business cycle accounting (BCA) methodology suggested by Cole and Ohanian (2004), and Chari, Kehoe, and McGrattan (2007). The BCA starts with the premise

8For the sake of brevity, we do not present the results of assuming a random walk process here.
Table 4. Current Account and Output Growth
(Data, Basic, and Extended Model)

<table>
<thead>
<tr>
<th></th>
<th>Current Account Share</th>
<th>Output Growth</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Data</td>
<td>Model basic</td>
</tr>
<tr>
<td>1980:86</td>
<td>-1.36%</td>
<td>-0.97%</td>
</tr>
<tr>
<td></td>
<td>-1.46</td>
<td>-2.25</td>
</tr>
<tr>
<td>1986:91</td>
<td>-2.03%</td>
<td>-1.85%</td>
</tr>
<tr>
<td></td>
<td>-1.3</td>
<td>-1.37</td>
</tr>
<tr>
<td>1991:2000</td>
<td>-1.85%</td>
<td>-1.35%</td>
</tr>
<tr>
<td></td>
<td>-1.21</td>
<td>-1.01</td>
</tr>
<tr>
<td>2000:03</td>
<td>-4.35%</td>
<td>-3.16%</td>
</tr>
<tr>
<td></td>
<td>-0.41</td>
<td>-0.24</td>
</tr>
</tbody>
</table>

Note: Model predictions when φ₁ is constructed so that model predictions of the current account levels match the data. In the table, we plot the current account as a share of output. To the extent that our model does not match output exactly, we have a discrepancy between the data and model generated current account share. We provide the results for the basic model (Chakraborty and Dekle, 2008) and the extended model presented in this paper, where the extended model includes adjustment costs of capital and the government sector.
that large classes of general equilibrium models are numerically equivalent to a prototype growth model with wedges, where wedges distort the first-order conditions, thus keeping the economy from reaching the first best outcome. BCA involves measuring these wedges, given the data and then evaluating the importance of each wedge by feeding them one by one and in various combinations in the model to see which wedge best accounts for the data, while keeping in mind that feeding in all wedges jointly would account exactly for the data (by construction).

As mentioned, $\phi_F$ or the cost of trading international assets is intended to capture frictions in financial markets that include, but is not restricted to regulations, and the financial openness of country $F$. Thus, by its very nature, it is a broad concept of increased openness in financial markets for which we do not have an exact equivalent in the data.

Methodologically, we feed in TFP into our model and then extract the time path of $\phi_F$ relying on the principle that TFP and $\phi_F$ jointly would exactly account for the evolution of the current account. The first step in constructing $\phi_F$ requires that we start with an initial guess of the stochastic process driving the evolution of $\phi_F$. Our initial guess is the same as in previous experiments (outlined in matrices $P$ and $Q$). We then iterate until our model predictions best match the data. To save space, we will only report the results of the extended model, since the pattern of the extracted $\phi_F$ in the extended model is almost identical to the pattern of $\phi_F$ in the basic model without capital adjustment costs and the government.

As mentioned, we also try to capture the patterns in U.S. output per capita. In general, both the basic and extended models overpredict the growth in U.S. per capita output (Figure 5, numerical comparison summarized in Table 4), although the overprediction is much less for the extended model. This is because in the extended model, government spending is assumed to be wasteful, and part of the capital inflows from abroad is used as government spending, rather than as physical investment. The introduction of adjustment costs to physical investment also tends to lower per capita output growth.

Is the pattern in $\phi_F$ plausible? First, note that $\phi_F$ starts to fall sharply from 1985 until 1991, and then stabilizes until about 1996. The decline in $\phi_F$ starting in 1985 corresponds to the financial liberalization that took place in Japan and in the United Kingdom in the mid-1980s. In both countries, entry into the commercial banking, insurance, and securities businesses were liberalized. Banks, insurance companies, and securities firms were allowed to lend to more sectors, including to foreigners (Dekle, 1998). Starting in 1996, $\phi_F$ increases until 2000, corresponding to the Asian financial crisis, where collateral values of governments and financial institutions in emerging Asia were damaged, leading to a deterioration of lending to the United States. Also, during this time, Japan was in the midst of its banking crisis, and Japanese bank lending to the United States deteriorated (Dekle and Ken Kletzer, 2003).

The sharp decline in $\phi_F$ starting in 2000 corresponds to the purchases of U.S. assets, particularly U.S. government bonds, by the Japanese and
emerging Asian, especially Chinese, governments, to prevent their local currencies from appreciating. This change in Asian foreign exchange reserve behavior, particularly by emerging market governments is captured by the sharp decline in $\phi_F$ from 2000. In our model, any change that makes it easier for the rest of the world to buy U.S. assets is represented by a fall in $\phi_F$ even if this change in buying U.S. assets is facilitated entirely by the foreign government.

Figures 6a and 6b compare our derived measure of the cost of international lending (6a) with an index of financial openness using the raw data provided by Chinn and Ito (2005)(6b). Chinn and Ito’s measure does not correspond directly to ours, as their measure also captures the liberalization of both inward and outward investment. For example, for Thailand, the Chinn and Ito measure captures how easy it is for foreigners to invest in Thailand; as well as how easy it is for Thai residents to invest abroad. Our derived measure of the decline in the cost of international lending only captures the ease of investing in the home (U.S.) country. In our model, the cost to U.S. residents of investing in the foreign country, $\phi_H$, is assumed to be constant.

In addition, while the Chinn and Ito measure is calculated from a careful reading of changes in regulations that made it easier for the rest of the world to invest in the United States, that is, a de jure measure, ours is a de facto measure that captures the actual net capital flows from the rest of the world.

**Figure 5. Index of Output (Data and Model, Basic and Extended)**

Source: Authors’ calculations.

Note: Plots the per capita output feeding in total factor productivity (TFP) in unison with declines in cost of bond trading. The results are plotted for the basic (in Chakraborty and Dekle, 2008), and the extended model where the extended model incorporates exogenous government expenditure and capital adjustment costs.
to the United States, exclusive of the productivity shocks in the two countries.

Despite these differences, the pattern in our measure corresponds with Chinn and Ito’s, especially between 1985 and 1994, when $\phi_F$ or cost of bond trading in our model is falling (or it is becoming easier to invest in U.S. financial markets), and the Chinn and Ito measure is rising (indicating increased financial openness). The Chinn and Ito financial openness measure declines from 1996 to 2000, which corresponds to a rise in our $\phi_F$, which, as mentioned, is related to the Asian financial crisis. However, from 2000 to 2003, the two measures drift apart; while $\phi_F$ falls sharply, the Chinn-Ito index shows less openness. The Chinn and Ito index may be capturing the capital controls instituted by many emerging markets after the Asian currency crisis, while our $\phi_F$, which focuses on capital outflows, cannot capture such capital inflow controls. For the entire period 1980 to 2003, the correlation coefficient between $\phi_F$ and the Chinn-Ito measure at $-0.41$ has the right sign; the two measures are found to be cointegrated by the Johansen test (with a trend and intercept term).

IV. Conclusion

We show that a standard, equilibrium macroeconomic model, augmented to capture changes in financial liberalization in the rest of the world, empirically explains well, the evolution of the U.S. current account from 1980 to 2003. Our results are robust to the inclusion of permanent (random-walk) productivity shocks. Using our procedure, we derive a series for the cost of

![Figure 6a. Cost of International Lending, $\phi_F$](image-url)
buying U.S. assets by foreigners, and compare it with other measures of financial liberalization in the literature. Our series also matches up well with distinct, identifiable episodes in international financial liberalization.

We show that the introduction of capital adjustment shocks and “wasteful” government spending explains partly the slow output growth in the U.S. economy. However the predicted per capita output growth in the model is much higher than actual per capita output growth in the data for the United States.

**DATA APPENDIX**

The world in our model is made up of two regions: the United States (referred to as “home”) and the EU-15 and Japan; or the EU-15, Japan and emerging Asia referred to as “foreign.” The emerging Asian region comprises of China, Hong Kong, Indonesia, Malaysia, Philippines, Singapore, South Korea, Taiwan POC, and Thailand.

For our calculations, we need data on national income, employment, hours worked, population, capital stock, and the current accounts. The national income accounts data along with the data on population, employment, and hours worked is calculated from the “Total Economy Database” and the “Industry Growth Accounting Database” that is maintained by the University of Groningen at the Growth and Development Center. The data on the capital stock are not available from the above data set; and we collect it from the Kiel Institute database on capital stocks in the OECD countries. The data for China are from Dekle and Vandenbroucke (2006) and for Korea, capital data for some of the
latter years are available from *Korea Statistical Yearbooks*. We use annualized data for the period 1980 to 2003. Given below is a description of the variables used and how they are constructed from our annualized data.

\(y_{it}\): The per capita output in the United States is calculated as aggregate GDP divided by the population. For the home country per capita output, we consider the weighted aggregate output of the EU-15 and Japan; and the EU-15, Japan and emerging Asia, where the weights are calculated as the share of a country’s GDP in the aggregate GDP of the rest of the world (ROW). The population of the home country is similarly calculated as the weighted average of the population of countries constituting our ROW, where the weights are calculated as the population of the constituent country divided by the aggregate population of the ROW. The output data are expressed in our data set in millions of 1990 U.S. dollars.

\(k_{it-1}/y_{it}\): The capital output ratio for the home country is calculated as the share of the weighted capital stock to weighted output, where the weights are measured as before as the share of the variable of the constituent country to the aggregate of the ROW.

\(l_{it}\): Labor is calculated as:

\[
l_{it} = \frac{E_{it}}{N_{it}} \times \frac{H_{it}}{(50 \times 100)},
\]

where

\(E_{it}/N_{it}\): Weighted employment to weighted population ratio for home and aggregate employment to aggregate population ratio for the United States.

\(H_{it}(50 \times 100)\): Weighted average of annual hours worked to total hours, where total hours is assumed to be 50 hours per week and there are 100 work-weeks. For the United States, instead of weighted hours, we just take the annualized hours worked.

\(g_{it}\): Per capita government expenditure. For our analysis, we need the time series of government expenditure from the United States that we collect from the U.S. Bureau of Economic Analysis. We arrive at the per capita figures by dividing aggregate government expenditures by the population.

*current account share*: Along with output, we try to match the current account share in output in the United States; where the data come from Bureau of Economic Analysis. The current account balance is collected from the Bureau of Economic Analysis that is in millions of dollars. To get the share of the current account balance in output, we divide the current account balance by GDP. The current account balance data of the United States with individual countries is limited; in particular data for emerging Asia including China are not available before 1999.

*weights*: To get the weights of the countries that comprise our rest of the world group, we take the GDP figures of individual countries from the Groningen data center which is expressed in millions of 1990 U.S. dollars (converted at Geary Khamis PPPs). The Geary-Khamis conversion method is popular in current international comparative studies as it has some desirable properties. The weights are then constructed by dividing the total GDP of each constituent country by the aggregate GDP of the group where the group comprises of EU-15 and Japan; and the EU-15, Japan along with emerging Asia.

*index of financial openness*: The index of financial openness was created by Chinn and Ito (2005)\(^9\) as a proxy for international financial market liberalization. They compiled an index of the degree of capital account openness for 163 countries from 1970 to 2004. The index is calculated on the bases of dummy variables that codify the restrictions on cross-

---

border financial transactions reported in the *Annual Report on Exchange Arrangements and Exchange Restrictions* (AREAER) from the IMF. The dummy variables reflect the four major categories on the restrictions on external accounts: presence of multiple exchange rates; restrictions on current account transactions; restrictions on capital account transactions; and requirements for firms to surrender a fraction of export proceeds. The index is the first standardized principal component of these four variables and it takes higher values for countries that are more open to cross-border capital transactions. For example, the United States that is calculated to be the most open economy in terms of financial openness has an index value of 2.602508 in Chinn and Ito (2005) estimates.

For our measure of financial openness (Figure 6), we take the index from Chinn and Ito (2005) for the group of countries that form our rest of the world, namely, EU-15, Japan, China, and the rest of emerging Asia. We then calculate the weighted average of the index where the weights are given by the share of an individual country’s GDP, in the aggregate GDP of the group.

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