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Northwest of Suez: The 1956 Crisis and the IMF

JAMES M. BOUGHTON*

Egypt's nationalization of the Suez Canal in 1956 and the failed attempt by France, Israel, and the United Kingdom to retake it by force constituted a serious political crisis with significant economic consequences. For the United Kingdom, it engendered a financial crisis as well. That all four of the combatants sought and obtained IMF financial assistance was highly unusual for the time and had a profound effect on the development of the IMF. This case study illustrates the complexities in isolating the current account as the basis for determining a balance of payments "need" and shows that the speculative attack on sterling—and the IMF's response to it—were remarkably similar to financial crises in the 1990s. [JEL F33, F34, N20]

On July 26, 1956, Egypt nationalized the Suez Canal Company and unilaterally assumed control of the canal from the international consortium that had run it for nearly a century. France, Israel, and the United Kingdom almost immediately began planning a joint military action to retake control, while they sought to win international support for a diplomatic solution. When diplomacy failed, Israel invaded the Sinai on October 29, and France and Britain used Egypt's counterattack as an excuse to attack Egypt by air from the Mediterranean two days later. The fighting shut down the canal, which was the major shipping channel between Europe and Asia and a vital link in the transport of petroleum from the Middle East. One week later, however, Britain undercut the operation by accepting a United Nations resolution for a

*This paper was prepared while I was on leave at St. Antony's College, University of Oxford. I would like to thank Shailendra Anjaria, Keith Kyle, Jacques Polak, Gregor Smith, John Toye, and anonymous referees for comments on earlier drafts.
ceasefire. On December 3, the British government announced that it would withdraw its troops over the next few weeks. France and Israel soon also withdrew, and Egypt reopened the canal under its own control the following April.

That this brief flare-up is universally regarded as a crisis is primarily because of the upheavals it engendered in political relations. It successfully climaxed Egypt’s longstanding campaign for full independence from European dominance. It demonstrated Israel’s ability to defend and expand its borders militarily and thus to survive as a nation. It weakened France just as the Algerian war was intensifying. It exposed a rift in relations between Britain and the United States over postcolonial policies at a time when both wanted to counter the rising regional influence of the Soviet Union. In Britain, it brought a sad end to the brief ministry of Anthony Eden and ironically elevated Harold Macmillan in his place. In view of the central place of Suez in the mythology of the British Empire—Eden had once called it “our back door to the East,” and generations had grown up on Kipling’s evocation of an uninhibited life “somewhere east of Suez”—the loss of control over the canal was devastating for those with lingering Victorian aspirations. In the Middle East, it solidified Gamal Abdel Nasser’s budding leadership role and hinted—if only temporarily—at the possibilities for Arab unity. A vast and still-growing literature has analyzed each of these facets in exquisite detail.¹

The economic consequences of Suez were more subtle and temporary and would not by themselves have constituted an international crisis. Notwithstanding the crucial importance of the canal for certain trade flows, the economic impact of its closing was limited by its short duration. By October, Egypt had already proved that it could run the canal safely and efficiently without European assistance. For the six months that the canal was closed, the resulting cost increases, delivery delays, and trade diversion weakened the current account positions of all four of the combatants, but normalcy was largely restored within another six months.²

For the United Kingdom, Suez was also a financial crisis. Throughout 1956 and 1957, Britain had a current account surplus despite the disruptions to its international trade. The value of its currency, however, came under speculative pressure, and the Bank of England was forced to deplete its U.S. dollar reserves to defend the fixed value of the pound sterling against the dollar. Macmillan (then Chancellor of the Exchequer) and Cameron Cobbold (Governor of the Bank of England) put on a brave front in characterizing the Bank’s ability to stave off an attack, but by December the threat of a forced devaluation or float was very real.

¹For a chronology, analysis, and basic bibliography of political events affecting the canal and Anglo-Egyptian relations from the 1850s to the crisis and its ramifications, see Gorst and Johnman (1997). For a full recounting and analysis of the crisis, see Kyle (1991); the “back door” quotation is on p. 7. Farnie (1969) provides a detailed history of the canal and its central place in the British Empire.

²On the general economic aspects of the crisis, see Kunz (1991); Johnman (1989) analyzes the economic effects on the United Kingdom. Also see IMF (1957), pp. 27–30, which notes that the economic effects were less severe than had been anticipated.
NORTHWEST OF SUEZ: THE 1956 CRISIS AND THE IMF

These events unfolded at a time when the IMF was almost totally untested in crisis management. From its first financial operations in 1947 to the onset of the Suez crisis, the IMF had lent only sporadically and in small amounts. The concept of stand-by lending subject to agreed policy conditions was still being developed and had been applied in only a few cases. The little financing that was being provided was mostly either gold tranche drawing (that is, countries were temporarily drawing out the reserve assets that they had deposited on becoming members) or was limited to the first credit tranche (that is, countries were borrowing no more than 25 percent of their quota, an amount equivalent in size to their paid-in gold tranche). In the two years preceding the Suez crisis, the IMF did almost no lending: drawings for 1954–55 totaled $90 million by just five countries (Colombia, Indonesia, Iran, Mexico, and the Philippines). Most of that was in the gold tranche. Only Mexico had a stand-by arrangement, and only the Philippines drew in the upper tranches (that is, borrowed more than the first credit tranche). In these circumstances, it was not obvious that the IMF should play any role at all in the resolution of the economic or financial difficulties of the countries involved in Suez.

A further reason not to anticipate significant IMF involvement was that the country with the gravest financial difficulties—the United Kingdom—did not obviously qualify for IMF assistance in 1956. The IMF’s Articles of Agreement prohibit it from lending to finance a “large and sustained” outflow of capital, which in essence was what Britain faced. That provision was intended to preserve the IMF’s limited financial resources for lending to promote international trade in goods and services. Moreover, if the speculative outflow from sterling was not both large and sustained, the Bank of England had enough resources of its own and enough access to credit that it could fend off the outflow without IMF assistance.

In addition to these factors, the IMF was in an interregnum during the crucial months of October and November and was missing its customary European leadership. Ivar Rooth, a former central bank governor from Sweden, completed his five-year term as Managing Director at the end of September. His

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3 For 1947 through 1955, 14 of the IMF’s 59 members made drawings on the IMF (excluding gold tranche purchases), averaging $46 million per annum (equivalent to 0.06 percent of world imports). For comparison, for 1990–98, 78 of the IMF’s 182 members made drawings on the General Account (excluding those in the reserve tranche), averaging $13,367 million (0.29 percent of world imports). Although the reserve tranche is formally equivalent to the old gold tranche (that is, it represents the member’s paid-in reserves), drawings on it cannot be compared. Gold tranche drawings were treated as temporary and had to be repaid within specified periods; reserve tranche drawings need not be repaid.

4 The first stand-by arrangement in which drawings were made conditional on the country adhering to specified policies was for Peru in 1954. That practice was gradually refined and made more general over the next several years but did not become standard practice until well into the 1960s. See Boughton (2001), Chapter 13, for an elaboration and further references.

5 The “commentary” prepared by the U.S. Treasury in 1944 to explain the purposes and operations of the proposed IMF expressed this point clearly. “In considering the probable attitude of the Fund toward the sale of foreign exchange to facilitate a transfer of capital, it should be borne in mind that the provisions of the Fund proposal are designed to give effect to the general principle that the Fund’s resources should be used primarily for settling international balances on current account.” Horsefield (1969), Vol. 3, p. 167.

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compatriot Per Jacobsson—then a senior official at the Bank for International Settlements (BIS)—was not due to arrive until December. Consequently, the Deputy Managing Director, H. Merle Cochran of the United States, was in charge when the IMF was called upon to respond to a series of politically sensitive financing requests.

Despite these obstacles, the IMF was called upon to help finance the external payments imbalances of all four combatants (Table 1). Over the course of nine months, it lent $858 million to those four countries and committed itself to another $738 million in credits on a stand-by basis. This paper reviews how the IMF came to be so heavily involved and analyzes the implications for its later role as an international crisis manager. Those implications stem primarily from the rescue of the pound sterling from a speculative attack: the first major financial crisis of the postwar era.

I. Conventional Financing: Egypt, Israel, and France

In the space of four weeks in September and October 1956, the IMF received financing requests from three of the four countries engaged in the Suez crisis. Each case was treated as a conventional problem of financing a temporary payments imbalance arising from the current account. None of these three countries had a convertible currency, and speculative pressures were unimportant. Each involved political complications, but ultimately the IMF was able to act upon them without becoming embroiled in the crisis.

Egypt

Egypt's economic difficulties intensified on September 13 when the government was forced to assume operations of the canal. The European boat pilots, on instructions from their former employers, abruptly walked out in an attempt to prove to Egypt that international control was necessary. Egypt, however, was determined to keep the canal open, and it brought in Egyptian pilots to take over. With costs rising and revenues plummeting, Egypt submitted a formal request to the IMF eight days later.

Since joining as an Original Member in 1945, Egypt would be drawing on IMF resources for only the second time. On the first occasion, in April 1949, it had drawn only 5 percent of its quota ($3 million), and it had repaid the credit the following year. Now it was requesting only to draw the full amount of its gold tranche (25 percent of quota, or $15 million). Formally, then, this was a routine request, necessitated by the presumably temporary pressure on Egypt’s balance of payments from the disruption in international trade. The only real question was whether politics would intrude.

By coincidence, the Annual Meetings of IMF and World Bank Governors were just about to begin in Washington when the Egyptian request came in on September 21. The next afternoon, the Executive Board convened in a rare Saturday session, at the Sheraton-Park Hotel where the Governors’ meetings were being held. Egypt’s Executive Director, Ahmed Zaki Saad, presented the case for
<table>
<thead>
<tr>
<th>Date</th>
<th>Type</th>
<th>Egypt quota = 60</th>
<th>France quota = 525</th>
<th>Israel quota = 7.5</th>
<th>United Kingdom quota = 1300</th>
</tr>
</thead>
<tbody>
<tr>
<td>September 22, 1956</td>
<td>gold tranche</td>
<td>15 (25%)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>October 17</td>
<td>standby arrangement covering gold and first credit tranches</td>
<td>262.5 (50%)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>December 10</td>
<td>gold and first credit tranches</td>
<td></td>
<td></td>
<td>561.5 (43%)</td>
<td></td>
</tr>
<tr>
<td></td>
<td>upper tranche</td>
<td></td>
<td></td>
<td></td>
<td>738.5 (57%)</td>
</tr>
<tr>
<td>February 1957</td>
<td>first credit tranche</td>
<td>15 (25%)</td>
<td></td>
<td></td>
<td></td>
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<td></td>
<td>gold tranche drawing on stand-by arrangement</td>
<td>40</td>
<td></td>
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<td></td>
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<td>60</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>April</td>
<td>gold tranche drawing on stand-by arrangement</td>
<td>31.3</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>May</td>
<td>drawing on stand-by arrangement</td>
<td></td>
<td></td>
<td></td>
<td>28.7</td>
</tr>
<tr>
<td></td>
<td>gold and first credit tranches</td>
<td></td>
<td></td>
<td></td>
<td>3.75 (50%)</td>
</tr>
<tr>
<td>June</td>
<td>drawing on stand-by arrangement</td>
<td></td>
<td></td>
<td></td>
<td>42.5</td>
</tr>
</tbody>
</table>
the request. The U.S. Director (Frank Southard) spoke in favor, and the French Director (Jean de Largentaye) consented by noting the absence of any legal basis for objecting. The British Director (William Edward, Lord Harcourt) abstained, as did the Netherlands chair. The official decision stated simply that the IMF “expresses no objection” to the request. At the suggestion of the Canadian Director (Louis Rasmisky), Rooth agreed that no public announcement would be made of this transaction. In this manner, despite the delicacy of the circumstances, the IMF rapidly approved its first Suez-related financing.

Israel

A few days later, the Governor of the Bank of Israel, David Horowitz—also in Washington for the Annual Meetings—made informal inquiries about enlarging and then drawing on Israel’s IMF quota. The request was temporarily rebuffed, for reasons that were partly economic and partly political.

In September, when Horowitz made his request, no one in the IMF had any knowledge about the secret arming of Israel by France or about the emerging invasion plans. The initial difficulty therefore was largely economic. When Israel had joined the IMF in 1954, the staff had expressed concerns that the economy was not yet stable enough to sustain a fixed rate of exchange for its currency. The IMF therefore had discouraged Israel from setting a par value, a step that it viewed as a prerequisite for drawing on IMF resources. For several months now, the Israeli authorities had been pushing the IMF to initiate the process for determining an appropriate value, but the staff had resisted. Israel was maintaining, and was gradually trying to dismantle, a complex system for external transactions. Most transactions took place at an exchange rate of 1.8 Israeli pounds per dollar, but the prevalence of bilateral trade agreements, subsidies, and multiple exchange rates meant that repressed inflation was likely to be a serious problem. In these circumstances, it was difficult for the IMF to assess whether 1.8 was a sustainable rate.

On October 22, the staff position shifted reluctantly from skepticism toward acceptance of 1.8 for Israel’s exchange rate. Although the staff analysis still suggested that Israel would have difficulty holding that rate for very long, the Director of the European Department (Gabriel Ferras) accepted the government’s argument that fixing the rate would provide a nominal anchor for wage negotiations and thus would help stabilize prices. Moreover, Israel clearly could benefit from drawing on IMF resources to supplement its foreign exchange reserves, and setting a par value would help pave the way. The Acting Managing Director, Merle

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6IM/CF, C/Egypt/1710, “Exchange Transaction 1948–59.” The Egyptian request was circulated in the document EBS/56/28, “Use of the Fund’s Resources—Egypt” (September 21, 1956), and the Board’s discussion and decision are in the minutes of Executive Board Meeting 56/48 (September 22).

7A standard clause in resolutions approving new memberships prohibited the member from drawing on IMF resources until after a par value had been approved. That prohibition could be altered or waived, but in this case the chances for approval were not strong. See IM/CF, C/Israel/1000, “Par Values and Exchange Rates 1948–75”; memorandum for file by Roman L. Horne, Secretary of the Fund (October 11, 1956).
Cochran, then placed the matter on the Executive Board agenda for discussion on October 31.8

The timing could not have been worse. After Israel invaded the Sinai on October 29, Cochran sensed that any discussion of IMF assistance to Israel would raise a political firestorm. On the morning of October 30, he went to see the U.S. Executive Director, Frank Southard, who agreed that the request should be pulled from the agenda if possible. Then he spoke with Peter Lieftinck, the Executive Director for the Netherlands who also represented Israel on the Board. Lieftinck was in a ticklish position, because the Israeli authorities wanted him to press to keep their request alive. Nonetheless, he followed his own instincts and agreed that Cochran should cancel the discussion.

The record shows clearly that the postponement of Israel’s request was purely a response to concerns about the political implications. Cochran explained to Southard that his suggestion was a response “to Israel invading Egyptian territory.” Lieftinck’s view was similar, and his preference was “that the matter be kept from coming before the Board, at least until after there might be some outcome of the Israeli-Egyptian question consideration by the United Nations.”9

On economic grounds, the staff also retreated from the earlier endorsement of the exchange rate proposal. The military buildup for the Sinai invasion was seen as likely to crowd out domestic investment and contribute to inflationary pressures, which would eventually force a severe tightening of policies or a devaluation.10 With hindsight, one can see that this concern was overstated, because the buildup was being partly financed in secret by France. The IMF, of course, was not informed of that relationship.11

Despite the widespread international opposition to Israel’s military action, the IMF’s reluctance to act was soon overcome by the restoration of normal leadership in the institution. Almost as soon as Jacobsson took over as Managing Director in December, he overruled the staff and instructed them to put the discussion of Israel’s par value back on the Board agenda. From then on, matters proceeded almost routinely. The staff dutifully produced a favorable recommendation on the proposed exchange rate of 1.8 per dollar, and the Board approved the proposal in mid-March 1957 (one week after Israel withdrew its forces from the Sinai), with Egypt abstaining.12

One other element in this process is worth noting, as it further illustrates the interplay of economic and political factors in the IMF’s response to Suez. Israel’s quota in the IMF, established in 1954, was extremely small and had been set primarily so as to be commensurate with existing quotas for neighboring countries.

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11France, Israel, and Britain officially denied the relationship for many years, until details of the planning finally became known. The full story of the unwinding of secrecy is told in Shlaim (1997). Kyle (1991) details the French military support for Israel, which he likens to lend-lease (p. 268).
12The more favorable view was not unwarranted. Israel maintained its exchange rate at 1.8 until February 1962, when it devalued to 3.0.
James M. Boughton

(Jordan, Lebanon, and the Syrian Arab Republic). At $4.5 million, it was near the bottom of the distribution and well below the range suggested by the standard formulas in use at the time ($9–$16 million). In 1956, in the context of the second Quinquennial Review of quotas, the Israeli authorities requested a substantial increase, to at least $18 million. Although the IMF did not deny that such an increase seemed warranted by the size and structure of Israel’s economy, the staff demurred on the grounds that “the governing factors in this case were overwhelmingly political.” In particular, if the Board agreed to raise Israel’s quota, neighboring countries would almost certainly submit similar requests, and the rest of the distribution would be difficult to maintain. As it happened, the Board had already agreed to raise all small quotas, on request, to a minimum of $7.5 million. Effective in March 1957, that increase was all that Israel could get.

Israel’s desire to borrow from the IMF was thus delayed until after the Suez crisis had ended, and even then its quota and therefore the amount it was entitled to borrow was kept small. Undaunted, Horowitz made a second attempt in May 1957. The staff now recommended approval and noted euphemistically that “developments in the Mediterranean” and the consequent rise in military spending had contributed to a worsening balance of payments position. Other factors also were important, including adverse market conditions for Israel’s exports and the effects of a surge in immigration, especially from Hungary and Poland, where the suppression of uprisings against Soviet control was causing many to flee. As the staff judged that the government was making a reasonable effort to overcome these developments, the use of IMF resources was appropriate.

The Executive Board met on May 15 to consider Israel’s request to draw 50 percent of its quota (the gold tranche and first credit tranche). Three Directors spoke in favor: Lieftinck, Southard, and Julio Gonzalez del Solar (Alternate Director, Guatemala). One abstained (Albert Mansour, the Alternate for Egypt), and the others were not recorded as speaking and were counted in favor. In the end, the handling of the request was essentially routine.

France

A third overture at the 1956 Annual Meetings came from the Governor of the Bank of France, Wilfrid Baumgartner. On Thursday, September 27, Baumgartner called on Rooth to alert him to the possibility that France might soon need to request a sizable stand-by arrangement. Reserves were low and falling rapidly, as the franc

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13Israel’s quota was the same as Lebanon’s, larger than Jordan’s ($3 million), and smaller than Syria’s ($6.5 million). For the political and economic background, see the papers in IMF/CF, C/Israel/1200, “Quota” and memorandum from John M. Stevens (Director of the Economic Department) to Rooth.

14Israel’s quota and attachment (March 16, 1956), in C/Israel/1210, “Quota Adjustment—Increase 1957.”

15See the memorandum cited in footnote 7.

16See Horsfield (1969), pp. 389–91. In the next three general quota increases (1966, 1970, and 1978), Israel obtained increases that were well above average and that finally brought its quota in line with its economy.

was being subjected to "a flight of capital which, however, is not serious at present." Rooth, who had just one week left as Managing Director, made a note for the record but took no further action.  

Two weeks later, the French Executive Director, de Largentaye, followed up on Baumgartner's hint by informing Edward M. Bernstein, the IMF's Director of Research, that France wished to enter into a stand-by arrangement for 50 percent of its quota. Although this commitment would be no larger than the one to Israel in relation to the borrower's quota, it would be the largest financial commitment in the IMF's fledgling history ($262.5 million). Moreover, the Executive Board had declared France to be ineligible to use IMF resources in 1948, and although the declaration had been lifted in 1954, France—like Israel—still lacked an approved par value for its currency. One might therefore expect the request to have been treated with particular care, but in the event it was processed without undue delay. Whether the IMF would have been more reluctant if it had known of the secret plans that France was developing for the invasion of Egypt is difficult to judge. The staff was able to move quickly in recommending approval of the French request, because a mission team had recently returned from Paris where it had reviewed economic and financial developments for the annual consultation report required under Article XIV of the IMF's Articles of Agreement. The mission's analysis of the economy focused—probably correctly—entirely on issues other than Suez.

The dominant adverse influence on the French balance of payments was the war in Algeria. In addition, severe frosts early in the year had sharply reduced agricultural output. Discussions of these and more general issues took place in Paris in July, before the nationalization of the canal, but even in mid-October neither the staff nor the Executive Board saw any need to raise the question of whether the Suez crisis was likely to disrupt the French trade or payments positions. As a corollary, Baumgartner's citation of capital flight as the basis for the request was never raised as an issue. The Board unanimously approved France's request for a stand-by arrangement on October 17.
II. The British Financial Crisis

The fourth Suez-inundated country to turn to the IMF was the United Kingdom, under rather different circumstances from the others. France's current account position was estimated to have deteriorated by $1.1 billion in 1956, from a $409 million surplus to a $700 million deficit, while Israel's deficit had widened by $75 million to $358 million. Egypt's current account balance had not yet deteriorated, but its prospects had been hit hard by the loss of canal revenues and by large internal expenditures by the Suez Canal Company. In contrast, the United Kingdom registered a current account surplus of £159 million in the first half of 1956, and the authorities expected a small surplus in the second half as well. In the event, the surplus turned out to total £245 million in 1956 (slightly higher than the previous year) and a similar magnitude in 1957. Moreover, because the British had the second-largest quota in the IMF, after the United States, they had the potential to place a serious drain on the IMF's stock of usable currencies (lendable funds). If the IMF aimed to conserve its resources for operations in support of current account balances, it would have to look carefully at the justification for extending credit to the United Kingdom.

The British problem, in a nutshell, was speculation that the Bank of England would have to abandon the sterling parity, which had been set at $2.80 in 1949. Maintaining that rate was important for several reasons. The government viewed $2.80 as appropriate for trade purposes, it feared the inflationary consequences of having to pay expensive dollars for oil imports while the canal was closed, and it regarded exchange rate stability as essential for preserving the sterling area as a preferential trade zone and sterling's broader role as a reserve currency. Although the United Kingdom had not yet established full external convertibility (it would do so in 1958), its system of capital controls was fragmented and porous. The pound was widely held as a reserve and investment medium and thus—in contrast to the other currencies affected by Suez—was subject to speculative pressure. The problem was compounded by the transparency of the reserve position, which the government published each month. The official reserve target since the late 1940s was to maintain a minimum balance of $2 billion ($2,000 million). To fall through that floor would be interpreted in financial markets as a signal that devaluation or even floating would have to be seriously considered.


24The Bank of England was obliged by the IMF Articles of Agreement to maintain the value of the pound sterling within 1 percent of the par value, viz. in a range from $2.772 to $2.828. The Bank, however, had undertaken publicly to maintain a narrower band, from $2.78 to $2.82.

25For an overview of the opportunities for speculation against the pound in 1956, which included leads and lags in trade payments, nonresident capital transactions, and capital transactions by U.K. residents within the sterling area, see Klug and Smith (1999), p. 185. For the rationale and origins of the reserve target, see ibid., p. 192.
Without an understanding of the importance that British officials ascribed to maintaining the sterling parity, Britain's total dependence on U.S. support for its Suez operations is difficult to comprehend. If devaluation or floating had been viable, the United Kingdom could have resisted external financial pressures for long enough to wage what likely would have been an effective military campaign against Egypt. Both of those options, however, were categorically rejected by Macmillan, and even more strongly by the Governor of the Bank of England, Cameron F. Cobbold. Significantly, they saw the consequences as spilling over from economics into political and diplomatic relations. In Cobbold's view, devaluation—"only" seven years after the last one, in 1949—"would probably lead to the break-up of the sterling area (possibly even the dissolution of the Commonwealth), the collapse of [the European Payments Union], a reduction in the volume of trade and currency instability at home leading to severe inflation." Consequently, "we should regard a further devaluation of sterling as a disaster to be fought with every weapon at our disposal." Although Macmillan knew full well that they could not hold the pound at $2.80 without U.S. support and that such support was not likely to be offered, he expressed his complete agreement with these views.26

For the United Kingdom, therefore, the need for assistance from the IMF resulted not from economics but from the psychological impact of a political crisis on financial markets. Eden's commitment to oppose Nasser in circumstances where victory was highly uncertain shifted market sentiment against sterling at a time when the Bank of England was known to have only a small cushion of liquid dollar-denominated claims. The U.K. authorities underestimated the threat from market speculation and focused instead on official holdings. For some months before Nasser nationalized the canal, the Bank of England had been preparing to protect sterling against a flare-up in the Middle East by blocking Egyptian accounts. When sterling came under immediate selling pressure in the wake of the nationalization on July 26, 1956, British officials suspected that Egypt was dumping sterling on the market, and they moved quickly to prohibit the transfer of sterling for Egyptian pounds.27 Not surprisingly, this gnat-swatting exercise did little to stem the pressure. Reserves dripped away gradually over the next four months and would have approached the self-imposed floor by end-October except for the serendipitous receipt of $177 million in September from the sale of the Trinidad Oil Company to an American firm (Figure 1).28

26PRO, T236/4188, note from Cobbold to Macmillan (October 17, 1956), with handwritten response from Macmillan (October 24); and minutes of a meeting between the two (November 16). For a detailed study of Anglo-American relations during the Suez crisis, see Lucas (1996).


28Net reserves as plotted in Figure 1 are adjusted for certain extraordinary transactions: balances due and paid through 1958 on U.S. wartime loans, the 1956 IMF drawing, receipts from the sale of the Trinidad Oil Company, and the 1957 U.S. loan. The gap between gross and net at the beginning of 1955 is $698 million, the amount that was due to be paid on U.S. loans over the next four years. The gap at the end of 1958 is $1,092 million, the sum of the IMF and U.S. Export-Import Bank credits, Trinidad Oil receipts, and a 1957 waiver of interest due on U.S. loans. For daily reserve movements during the second half of 1956, see Klug and Smith (1999), Figure 3. Those data also appear to exclude some extraordinary transactions, most notably the IMF drawing.
Britain's first line of defense was intervention in the form of spot purchases of sterling in the foreign exchange market. As Klug and Smith (1999, pp. 199-200) have noted, the Bank of England—"in the Keynesian spirit of the time"—did not believe that the classic monetary response of raising short-term interest rates would do any good. Tightening the budget would help, but not quickly enough to resolve a financial crisis: even in 1956.

It was clear almost immediately to both the Treasury and the Bank that a second line of defense was needed to protect the reserve floor. Despite the U.S. opposition to the European effort to force Egypt to return control of the canal, Macmillan hoped to build on Britain's and his own "special relationship" with the United States (his mother was American) to persuade Washington to help him support sterling. Either they could temporarily waive the interest due on lend-lease credits advanced during the Second World War, or they could provide new loans through the Export-Import Bank. If that line of defense failed, Macmillan expected to be able to count on the apolitical tradition of the IMF to draw the modest amounts to which Britain was virtually entitled. More important, though with less than complete logical consistency, he expected to be able to build on Britain's informally accepted special status in the IMF—as one of the two major founding countries and the second-largest member—to draw much larger amounts and to the fullest possible extent.29

29 As late as 1967, long after performance criteria (required policy adjustments) had become standard practice in IMF stand-by arrangements, it approved arrangements for the United Kingdom without formal policy conditions. Prior to 1956, the United Kingdom had drawn small amounts on four occasions in 1947-48 but had maintained a credit position since November 1953.
Macmillan put the IMF card on the table just two weeks after the nationalization of the canal, when he told a Treasury meeting “that he was proposing to ask Australia to sell us some gold, and was also considering whether we should not withdraw dollars from the International Monetary Fund.” Over the next two weeks, the Treasury considered whether they could try to draw on the IMF even without U.S. support. A simple majority of votes cast in the Executive Board, which was all it would take, was conceivable but very unlikely against U.S. opposition. All seven European Directors, Australia, and Canada were expected to vote in favor of a British drawing. The U.S. and Egyptian Directors and all three from Latin America would probably oppose it. Only if the remaining three—from China, India, and Japan—could all be persuaded at least to abstain would the motion carry. That prospect was quite dim. Even worse, as Macmillan readily acknowledged, once investors knew that the IMF resources were sterling’s last line of defense, even a large stand-by would “not do much to put off the day” when the parity would have to be abandoned. American support was therefore essential.30

The occasion for testing sterling’s defenses was the same one seized by the other three countries for approaching the IMF: the Annual Meetings of the IMF and World Bank Governors in Washington at the end of September. Macmillan, participating as the Governor in the IMF for the United Kingdom, met privately with his American counterpart, Treasury Secretary George M. Humphrey; with the Secretary of State, John Foster Dulles; and with President Eisenhower.31 Discussions with both Humphrey and Dulles touched on the prospects for U.S. financial support, especially through waivers of interest payments on wartime credits. Only with Humphrey did the possibility of getting IMF credits figure prominently.32 Although Humphrey clearly gave Macmillan no promises, the Chancellor interpreted his and the other officials’ comments as an assurance that some form of U.S. financial aid would be forthcoming after the presidential election of November 6. He therefore returned home confident that he could count on his American friends to help him maintain “the strength of sterling.” As he recalled bitterly in his memoirs, “[t]here was no hint, at this time, of any difficulty being put in our way, or of financial backing to Britain not being available in full, whatever the circumstances” (Macmillan, 1971, p. 135).

Macmillan took no further action during October, as the Bank of England continued to sell off its dollar reserves to maintain the $2.80 exchange rate and the Cabinet continued to prepare for war. Through this period, the Bank’s reserves were not so much under attack as merely dripping away. Israel’s invasion of the Sinai on October 29 and the opening of the Franco-British military offensive two

30PRO, T236/4188, note from Norman Brook to Edward Bridges (August 9, 1956), and note from Leslie Rowan to Macmillan (September 21) with handwritten response from Macmillan (September 25). For the Treasury’s expectations on how each Director might vote, see T236/4189, cable from Harold Caccia to the Foreign Office (November 9). Brook, Bridges, and Rowan were senior Treasury officials, and Caccia was the British Ambassador to the United States. For the distribution of voting power in the IMF in 1956, see Horsefield (1969), Vol. 2, p. 353.


32See NARA, Records Group 56, Entry 70A6232, Box 78; William Harcourt to Andrew Overby (Assistant Treasury Secretary), “The Chancellor of the Exchequer’s talk with the Secretary of the Treasury,” September 25, 1956.
days later solidified U.S. and world opposition and greatly accelerated the drain on reserves, to a pace that clearly constituted a speculative attack (Klug and Smith, 1999, p. 191). On November 2, the United Nations General Assembly overwhelmingly approved a U.S. resolution calling for a cease-fire and withdrawal of forces. Four days later, the British Cabinet bowed to the relentless financial and diplomatic pressure and agreed to a cease-fire.

Although Macmillan persisted in denying that financial pressure was the decisive factor, none of the extensive evidence supports a credible alternative explanation. Moreover, Macmillan acknowledged in his memoirs that the Cabinet learned during its crucial meeting on November 6 that the United States would not support its "technical" request to draw on the IMF "until we had agreed to the cease-fire" (Macmillan, 1971, p. 164). Exactly how and when the request was made to the U.S. authorities and who conveyed the negative response has never been established. That ambiguity, however, does not vitiate the conclusion reached by Gorst and Johnman (1997, p. 133), that "it is clear that it was financial pressures that were driving British policy" to accept the cease-fire. The more important ambiguity relates to the policy requirement. As subsequent events demonstrated, the Americans were insisting not just on a cease-fire, as Macmillan claimed, but on full compliance with the UN resolution; that is, on an immediate withdrawal of all troops from Egypt.

With the American obstacle apparently out of the way, attention turned again to how to put together a large enough support package to stop the run on sterling. Cobbold opposed making an immediate approach to the IMF, on the grounds that speculators would interpret it as a sign of weakness unless it was announced as part of a broader package of assistance. Treasury officials were less convinced of the virtues of waiting, since they suspected that the IMF was their only hope. They all agreed, however, that if they were to approach the IMF, they should try to borrow the maximum possible amount, which they expected to be "three tranches," or 75 percent of quota. Although a few small countries (Burma, El

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33"The falling away of reserves was not in itself calamitous.... Although ... I did not conceal the seriousness of the financial situation, this was not the reason for our acceptance of the cease-fire" (Macmillan, 1971, p. 164). Also see Rhodes James (1986), p. 573, which describes financial factors as "decisive" but not "the only factor"; and Fforde (1992), p. 556, which suggests (rather outrageously) that the best way to restore confidence in sterling might have been to forge ahead militarily and capture the canal.

34Macmillan (1971, pp. 163-64) claimed that he called "New York" (presumably the Federal Reserve Bank), which referred him to "Washington" (presumably the Treasury), which then sent a reply. Eden (1960) referred to a telephone call from Humphrey to R. A. Butler (Leader of the House of Commons), in which he "made it clear that the United States would not extend help or support to Britain until a definite statement of withdrawal had been made" (p. 572). Although Eden did not specify the date of the call, the context suggests that it would have been around November 6. Fforde (1992, pp. 556 and 563-64) noted the absence of documentary evidence and the improbable timing of the calls and concluded that Macmillan probably never made them. That would imply that Macmillan had already given up on the possibility of gaining U.S. support and that he also made up the story about receiving a reply during the cabinet meeting.

35PRO, T236/4188; Rowan to Roger Makins (Permanent Secretary) [and through him to the Chancellor], "The Reserves, IMF (a standby or drawing), and the Waiver" (October 26, 1956); and minute by Rowan of a meeting between Macmillan and Cobbold at the House of Commons (October 30); T236/4189, note from Makins to Macmillan (November 9); and record of a meeting in Makin's office (November 12). Kyle (1991), p. 500, misinterprets Cobbold's "three tranches" suggestion as implying a drawing of 100 percent of quota, rather than 75 percent.
Salvador, and the Philippines) had drawn their full quotas, none of the large countries had gone higher than 50 percent. A drawing by the United Kingdom of 75 percent would total nearly $1 billion, almost four times the size of the previous record, the October stand-by for France. Nonetheless, both the Treasury and the Bank were determined to try to get it.

Curiously, the IMF staff played a largely passive role in this developing drama. A staff mission team was in London in late November, but it was there only to conduct the routine annual consultations. The mission does not seem to have questioned the authorities' determination to maintain the exchange rate at $2.80 (though it may have done so in high-level meetings, without written documentation), nor the stance of macroeconomic policies. Nor did it raise formally the question of whether a drawing from the IMF might be appropriate or necessary. The issue, after all, was political rather than economic: the exchange rate level was appropriate if temporary financing could be secured. Although the mission was led by the highly respected Irving Friedman, Director of the Exchange Restrictions Department, none of the internal Treasury correspondence relating to Suez even mentioned the ongoing discussions with the staff. The mission completed its technical work in five days, but the consultations were then suspended until after the matter of financial assistance was resolved in Washington.36

Even if British officials took no notice of the staff, they were anxious to ensure that the IMF's management would be at least neutral if the United States continued to oppose them. They were greatly reassured when the Managing Director-elect, Per Jacobsson, paid a call on Cobbold on November 22, while on his way to take up his duties in Washington. Although Jacobsson seems not to have made any promises, Cobbold was quite pleased with their meeting. "You may like to know," he reported to Macmillan and the Treasury, "that Dr. Jacobsson agrees, both in U.K. interests and in world interests, with our policy, and that he considers that our present exchange rate is fundamentally sound on an international economic basis."37

Jacobsson in fact had been even more persuaded by Cobbold than the Governor realized. Jacobsson's biography, written by his daughter largely on the basis of his diary, conveyed his views in dramatic terms:

When PJ walked into his office at the IMF, on Monday 3 December 1956, he was determined that the financial assistance the UK would receive was to total $1.300 m. The speculation against sterling ... could only be stopped if really large funds were available. The alternative, as he learnt when passing through London, would be a floating sterling rate, the consequences of which would be unpredictable. There would therefore have to be maximum financial assistance. This was the conclusion PJ reached during the boat journey which took him from his job at the BIS to his new post at the IMF.38

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36The mission documentation is in IMF/CF. C/United Kingdom/420.1. "Exchange Restrictions Consultations 1956:"
37PRO. T236/490: Cobbold to Makins (November 22, 1956).
Meanwhile, Macmillan still lacked American support, and he was forced into supplication to get it. (The Prime Minister, Eden, had by now withdrawn to Jamaica in poor health.) Despite the cease-fire, British and French troops remained on the ground in Egypt. The U.S. government wanted them out, and Britain's need for financial assistance gave them the perfect lever to force an immediate withdrawal. Macmillan tried unsuccessfully to arrange a meeting with Humphrey in late November, but he managed to convey to him through William Harcourt and Harold Caccia (Executive Director and Ambassador, respectively) that a failure to support sterling could have catastrophic political consequences, including a triumph for international communism. Such threats doubtless seemed fanciful to Humphrey, who replied simply that the United States would support Britain when the latter was “conforming to rather than defying the United Nations.” Even then, he warned that a drawing on the IMF beyond the first credit tranche (that is, more than $561 million) would be problematic. A larger drawing, Humphrey argued, could cause "a run on the Monetary Fund, which might be as serious as a run on sterling."39

Britain faced a firm deadline for obtaining approval of a financial support package. On December 4, Macmillan would have to announce that a massive loss of reserves in November ($279 million, net) had pushed the balance below the $2 billion floor. Without support, the parity would have to be abandoned. Humphrey was on a short vacation and would not return to Washington until December 3, the same day that Per Jacobsson was to arrive for his first day as Managing Director. If the pound was to be saved, it would have to be saved on December 3.

Left with no alternative, the British Cabinet accepted the second half of the UN resolution and set a deadline of December 22 for a full troop withdrawal. Harcourt and Caccia then called on Humphrey to find out how much financial aid this capitulation had purchased. The extent of bilateral support was still vague but now could be counted upon and publicly announced as forthcoming.40 Of more immediate and concrete concern was the amount of IMF credits to be put at Britain's disposal.

At the beginning of the meeting on December 3, Secretary Humphrey continued to insist that his government could not support a large-scale support operation from the IMF. The U.S. Treasury would have to borrow the dollars to finance the operation, which he feared would put upward pressure on interest rates in the New York money market. Moreover, such a large operation, in his view, would violate the IMF's own operating principles and could lead other countries to submit similar requests. Then, quite suddenly and to the astonishment of his

39PRO, T236/4190, cables from Caccia to the Foreign Office (November 27, 1956). The appeal on grounds of stopping the spread of communism came only after Caccia dissuaded Macmillan from stressing the dangers to the sterling area, which clearly would have made even less of an impression on the Americans.

40Negotiations led quickly to agreement on a $500 million loan from the U.S. Export-Import Bank, which was announced on December 21 and executed the following March. More protracted negotiations led in 1957 to a loosening of the conditions under which Britain could reschedule interest payments due on outstanding lend-lease credits. See Kunz (1991), pp. 162 and 181.
British visitors, he swept aside those worries and reversed course. According to Caccia,

Humphrey, who was clearly thinking aloud, then suddenly switched his point of view. He said that if we went beyond the first two tranches, for which there was precedent, he was not sure that it would not be better for us to go for all four tranches. The object was to demonstrate beyond all doubt to the world that sterling was supported, and had resources sufficient to withstand any attack. Would it not be better to draw the first two tranches and get a standby for the other two?41

The crisis was over. When Macmillan revealed the November reserves losses in the House of Commons the next day, he was able simultaneously to announce that Britain would be making “an immediate approach” to the IMF as part of a broad effort to “fortify” reserves, although he was still circumspect regarding how much of the U.K. quota might be available.42 Speculators against sterling still had a one-way bet, but the odds were now pretty long against it.

All that remained was for the Executive Board to ratify the arrangements that had been agreed bilaterally between the two great powers. Success seemed assured, because—as indicated by the quotation given above—Per Jacobsson shared Humphrey’s view that a $1.3 billion package was needed to stem speculation against the pound. On December 6, as the Board meeting approached, Jacobsson wrote in his diary that “since the confidence factor played such a great role the amounts ought to be high enough to impress the market.” The two men also agreed that this operation should be a special case. Humphrey calculated that no other country would dare ask for such massive support and, even if one did, it was manifest that the IMF could not afford to repeat it very often. As the Managing Director observed in a speech on December 6, the IMF “cannot fuel all the people all the time.”43

Two potential roadblocks remained. First, the United Kingdom was technically in violation of the IMF’s Articles of Agreement for having failed to notify it officially of the imposition of exchange controls against Egypt at the end of July. British officials were aware that Egypt could choose to delay the funding request by raising a legal challenge. In the event, Egypt did not do so, and the required notification was made a few months later (see above, footnote 27).

The second, and potentially more troublesome, obstacle was the prohibition in the Articles against lending to finance a large and sustained capital outflow. On

41 PRO, T236/4190, cable from Caccia to the Foreign Office (December 3, 1956). Frank Southard, the U.S. Executive Director, later wrote that Humphrey had cleared the four-tranche proposal with him before floating it to the British, saying that “he had always believed that if a big job was to be tackled, one should go all out.” See Southard (1979), p. 20.
42 The U.K. quota in the IMF was $1.300 billion, and the country had a credit balance (a gold tranche position) of 18 percent of quota, or $236.5 million. The proposal was to make up a package totaling 100 percent of quota, comprising an immediate drawing of $561.5 million and a 12-month stand-by arrangement for $738.5 million. If fully utilized, the United Kingdom would have had a debit position of 82 percent of quota, but it would have faced an obligation to repay the full 100 percent in credits (see footnote 3).
43 Both quotations are from Jacobsson (1979), p. 285.
December 5, Harcourt met with Jacobsson and several senior IMF staff in the Managing Director's office, to discuss strategy for handling the financing request. They acknowledged that the credits would finance a capital outflow, but Jacobsson suggested (rather stretching the point) that to the extent that these flows were in the form of "leads and lags" in payments, they were linked directly to the financing of the current account. Harcourt then made a more coherent argument: without this financing, it would be difficult for Britain to maintain progress toward establishing full convertibility of sterling for current account transactions. Until this point, the IMF had followed the lead of its founding fathers (John Maynard Keynes and Harry Dexter White) in regarding the current and capital accounts as essentially separable phenomena. Because Britain's role as an international banker made that separation impossible, this rescue operation was about to break the mold.

The staff report that was circulated to Executive Directors two days later stated explicitly that financing was needed for the capital account, not the current account. It also raised the specter of systemic repercussions if the IMF failed to act on the British request:

The Staff believes that the policies designed to restrain internal demand...will be adequate to maintain the United Kingdom's competitive position and to overcome the temporary payments difficulties and trade dislocation which the U.K. economy will experience in the coming months. However, given the status of sterling as an international currency and the United Kingdom's role as banker for a large trading area, its efforts to overcome its balance of payments difficulties and to follow a policy of extending the area of freer trade and payments could be undermined if confidence in sterling were weakened by further sustained losses of monetary reserves. Moreover, the danger would arise of serious repercussions on the volume of world trade and on the progress made in freeing trade and payments from external restrictions.

The Executive Board approved this rationale and approved the request with one abstention (Egypt). Britain immediately made the drawing of $561 million

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44 Minute of the meeting, by the Personal Assistant to the Managing Director (December 6, 1956); in IMF/CF, C/United Kingdom/1760, “Stand-by Arrangements 1952–1960.”
45 See “Use of the Fund’s Resources—United Kingdom” (EBS/56/44, Sup. 1. December 7, 1956). p. 6; in IMF/CF, C/United Kingdom/1760, “Stand-by Arrangements 1952–1960.” A year later, the IMF’s Deputy Director of Research, Jacques J. Polak, stated this rationale even more clearly in an internal memorandum: “The main reason for the U.K. drawing last year and the only reason for the standby was the need to counteract speculative capital movements.... it is obvious that the unprecedented large Fund assistance was given to counteract and, insofar as possible by its announcement effect to forestall, a flight of capital from sterling” (memorandum dated November 29, 1957; op. cit.).
46 The Executive Director from Egypt, Ahmed Zaki Saad, abstained himself from the meeting. The constituency's abstention was announced by his Alternate, Albert Mansour (also from Egypt), who noted for the record that 3 of the 12 countries in the diverse constituency (Lebanon, Pakistan, and the Philippines) supported the British request. (IMF rules do not permit the splitting of votes within a constituency.) See IMF/CF: minutes of Executive Board Meeting 56/59 (December 10, 1956).
to replenish its reserves and announced that it had another $739 million available on stand-by.

Both the U.K. and the U.S. authorities attached great importance to the goal of not drawing on the stand-by arrangement. It was important to Britain for generating confidence in sterling and to the U.S. authorities for limiting the amount of dollar financing that they had to arrange. At a meeting in Paris the day after the Executive Board approved the arrangement, Humphrey stressed to Macmillan that Britain could not “afford to draw a single penny of the standby” without damaging confidence. Macmillan agreed and hoped that they had put enough money “into the shop window” to make a further drawing unnecessary. In fact, they had. The announcement soon stopped—though it did not reverse—the reserve outflow, and for the next several months British reserves fluctuated above the $2 billion floor (Figure 1). New concerns about loose macroeconomic policies led to a resumption of reserve outflows in the second half of 1957 and induced the government to negotiate a renewal of the stand-by arrangement in December, but at no time did they have to draw on it. 47

III. Consequences

The most obvious consequence of the IMF’s involvement in the Suez crisis is that it put it on the map as an episodic international lender. 48 For the first time, the IMF had played a significant role in helping countries cope with an international crisis. Subsequently, it was called upon repeatedly to deal with other shocks to the financial system, notably the sterling crises and the gold pool crisis of the 1960s, the oil shocks of the 1970s, the developing country debt crisis of the 1980s, and the financial crises in Mexico, Russia, and Asia in the 1990s (Figure 2). Although the IMF also began to lend regularly to help countries cope with the temporary payments effects of economic imbalances, that ongoing activity was quite small in amount relative to the occasional spurts occasioned by financial crises.

What has been lost in the previous discussions of these events is the striking modernity of the 1956 sterling crisis. When the “tequila” crisis hit Mexico in 1995, the IMF’s Managing Director, Michel Camdessus, called it the “first financial crisis of the twenty-first century.” Virtually all of the elements of that situation that made it seem newly complex were, however, present in the Suez crisis 40 years earlier.

First, what the United Kingdom faced in 1956 was almost purely a speculative attack on a stable currency against a backdrop of reasonably sound economic policies. That is, it was a financial and not an economic crisis, and its primary effect was on the capital account of the balance of payments. Mexico in 1995 had a current account deficit and faced a much greater need for economic adjustment, and it therefore lacked the financial resources to preserve its fixed exchange rate.

47See NARA, Records Group 56, Entry 70A6232. Box 81: “Secret Record of a Discussion at the Hotel Talleyrand, Paris, at 3 p.m. on Tuesday, December 11th, 1956.” The $2.80 rate held for another decade, until the 1967 devaluation to $2.40.

even with $40 billion in financial commitments from the IMF and other external creditors. In both cases, however, the most pressing requirement for resolving the crisis was to stem the speculative attack.

Second, in both cases the crisis was precipitated by a clash of policy goals, between maintaining a stable exchange rate and simultaneously establishing open markets for the currency. The United Kingdom in 1956 was on a path toward establishing full convertibility for sterling. Mexico in 1995 had already done so but was in danger of being forced to reestablish exchange controls to avoid a disastrous depreciation of the peso exchange rate. What was new in the 1990s was the relevance of this conflict for a country with an emerging market and a nonreserve currency. In the 1950s, even the potential for market speculation was relevant only for a very few countries.

Third, a rapid response was essential. Despite the limited convertibility of sterling in 1956, Britain began losing reserves rapidly after the United Nations condemned its invasion of Egypt in early November. The IMF had to respond by early December if Britain was to avoid floating the pound. The length of time between the onset of the attack and approval of the financial package was almost the same as in the peso crisis of 1995.49

Fourth, the key in both cases was to post a large enough number to impress financial markets, convince speculators that a bet against the currency could not be won, and persuade investors to keep their money in the country. In the tequila...

49 Mexico initially devalued the peso on December 20, 1994. It applied for an IMF stand-by arrangement on January 5, and the request was approved on February 1. The two circumstances, however, were not fully comparable in that the U.K. arrangement did not require the negotiation of policy conditions.
crisis, the IMF was forced to increase its commitment by $10 billion after the U.S. Congress refused a request from the Clinton administration for a package of loan guarantees. Markets had come to expect a total multilateral package of $40 billion, and the only way to avoid a resumption of panic selling was to assemble commensurate financing. In 1956, a similar psychological floor was created by the known commitment of the British government to maintain reserves of at least $2 billion. A commitment by the IMF had to be large enough to protect that floor if it was to do any good at all. The 1956 commitment was unprecedented in absolute size; the 1995 package was unprecedented both in size and in relation to the borrower's quota. In both cases, the required size of the package was determined by market psychology, not economics.

Fifth, in neither case could the return of private sector investors be assured, but in both cases it eventually emerged spontaneously. Britain was unable to restore the level of foreign exchange reserves, net of drawings from the IMF, until well into 1958. Similar delays characterized the financial crises of the second half of the 1990s, as the initial effect of multilateral rescues was to replace rather than to restore international private investment. These experiences contrast with the more tightly controlled handling of the Latin American debt crises of the 1980s, when “concerted lending,” multiyear reschedulings, and complex “menus” for restructuring existing debts were commonly invoked. With a longer focus, recent experience appears to be both more normal and more viable.

Sixth, the IMF’s involvement in both cases was necessitated by the unwillingness of the United States to provide sufficient resources bilaterally, despite its acknowledged self-interest in a successful resolution of the crisis. In 1956, even after the United Kingdom had acquiesced to American political demands, all that the U.S. government could promise in the short run was a modest loan from the Export-Import Bank. In 1995, all that it could promise to Mexico was $20 billion in short-term credits through the Exchange Stabilization Fund. In both cases, a much larger multilateral package would have to be assembled to end the crisis, and the IMF was the institution that was best placed to do so.

Because no one recognized these parallels in 1995, the Mexican case appeared to be a much more radical departure from past practice than it was. When the IMF made an even more rapid and large-scale commitment to Korea in the midst of the Asian financial crisis in 1997, it was building on a tradition that extended back not 2 years, but more than 40.

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Redistribution Through Public Employment: The Case of Italy

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This paper examines the regional distribution of public employment in Italy. It documents two facts. The first is that public employment is used as a subsidy from the North to the less wealthy South. About half of the wage bill in the South of Italy can be identified as a subsidy. Both the size of public employment and the level of wages are used as a redistributive device. The second fact concerns the effects of subsidized public employment on individuals' attitudes toward job search, education, "risk taking" activities, and so on. Public employment discourages the development of market activities in the South. [JEL H53, J31, J64]

Two key roles of government are to provide public goods and to redistribute income across individuals and regions. Often these two functions overlap since public goods provision may also be used to compensate for geographical income imbalances. Public employment, in particular, can be used to support poorer regions or those with higher unemployment. This paper documents the size of this type of redistribution between the North and South of Italy and attempts to evaluate the efficiency of this type of policy. Italy is an especially...
interesting case because of the large income disparity between North and South and because of the large size of the public employment sector.

In this paper we first document the amount of geographical imbalance in the allocation of public jobs. Using survey evidence collected by the Bank of Italy, we then highlight various cultural and social consequences of an extensive reliance on public employment as a source of jobs and income. Third, we evaluate the amount of redistributive flows achieved with public employment.

Our results are striking. We conclude that about half of the public wage bill in the South of Italy can be defined as a “subsidy.” This effect is due to a combination of the size of public employment and of the wage premium for public employees relative to alternative occupations. We also show that the reliance on public jobs as a redistributive channel implies sizable and possibly undesirable sociological effects. Since public jobs in the South are more attractive and available than private sector jobs, educational and attitudinal choices are tilted toward the public sector. Also, individuals do not want to exit the public sector unless they are forced to, and this creates path dependence and rigidities.

In a nutshell, the argument is the following. The two “regions” of Italy (North and South) are bound by a unitary fiscal system, which implies that public wages are almost identical in nominal terms between the North and South. Since the cost of living is much lower in the South, real public wages are lower in the North than in the South. Also, opportunities in the private sector are better in the North, so public employment is comparatively more attractive in the South, relative to alternative opportunities. As a result, residents in the South seek more public employment in order to take advantage of a large income premium and a greater job security. Over time the South is caught in an equilibrium of dependency in which public jobs are a critical source of disposable income and in which private opportunities do not materialize. This creates a culture that discourages private activities and entrepreneurship and that becomes self-fulfilling: the less individuals are prepared to “face the market,” the more they prefer public jobs.

But, if this is the case, why is this redistributive system chosen? One answer may be that this is simply a by-product of a centralized fiscal system and centralized union bargaining, which fixes equal nominal wages for the entire country. However, the lack of any attempt to diversify public wages between the North and South suggests that the implied redistribution might be politically desirable. The reason may be that redistribution through public employment is less visible than direct transfers, therefore it is politically less costly and may be more effective at creating patronage for local politicians. In fact, a model by Coate and Morris (1995), slightly modified by Alesina, Baqir, and Easterly (2000), clarifies this politico-economic argument. The idea is simple: suppose that a proposal that introduces a tax in region 1 (North) to finance a direct subsidy to region 2 (South)

See Raffa and Zullo (1993) for a discussion of the difficulties of small private innovative business ventures in the South.
would not pass because it is opposed by voters in the North. Further assume that the government wants to redistribute toward the South and assume that, say, several new teachers are hired and disproportionately placed in the South. This second redistributive policy is less transparent (although perhaps less efficient) and may win approval even in the North because of the uncertainty about the real needs of the public school system.

Public employment may also be used to correct labor market imperfections. When labor markets do not produce full employment, say, because of tax distortions and rigidities, it is politically rewarding to offer public sector jobs. This is particularly the case when the welfare system (as in Italy) is distorted and ineffective at protecting the temporarily unemployed. In fact, Rostagno and Utili (1997) and Boeri (2000) describe the shortcomings of the Italian system of social protection and conclude that the Italian “welfare state” is very skewed in favor of retirees and does not protect efficiently the temporarily unemployed. Obviously, while a temporary unemployment subsidy may create incentives for job search, a permanent employment in the public sector does not.2

Public bureaucracies, once established, become a major political force. In many countries, and certainly in Italy, public sector unions are particularly strong and capable of protecting job security, if not the level of real wages.3 This protection generates hysteresis: once public employment increases, it takes a long time to be reduced.

This is not the first paper that argues that public employment is used as a redistributive device. To begin with, there is an immense literature on public sector employment, most of which is focused on the United States. We refer the reader to the two excellent surveys by Ehrenberg and Schwarz (1986), and Gregory and Borland (1999). For our purposes, the latter paper, which focuses not only on the United States but on the evidence available for other member countries of the Organization for Economic Cooperation and Development (OECD) as well, concludes that “public sector employees generally have higher average earnings than private sector employees.” Furthermore, the authors write, “in most countries, some part of this difference is also attributable to higher rates of pay or rents for public sector employees.” Particularly interesting are the results of Borjas (1986), who examines wage variations in U.S. state public employment and attributes three-fourths of the interstate variation to political variables reflecting the demand of different constituencies. Also, Katz and Krueger (1991) find that in the United States, while local and state governments are responsive to local economic conditions, the market for federal employees is set outside the regional context.

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2A related problem concerns the use of disability pensions in Italy. These pensions have been largely used especially in the South as permanent unemployment subsidy, with the obvious distortionary effects on incentives. See Boeri (2000).

3For a review of the literature on public unions, see Gregory and Borland (1999) and Freeman and Ichniowski (1988).
I. The Distribution of Public Employment in Italy

The Data

As a source for macroeconomic data on regional differences, we draw on various Italian government statistics. Data on regional production, population, and employment are taken from publications of Istituto Nazionale di Statistica (ISTAT), Italy’s national statistical institute (ISTAT. 1996a and 1996b). Figures on the regional distribution of public employment are taken from Il Conto Annuale (Italian Treasury, 1995), an annual publication of the Italian Treasury. Our data for postal and railroad employees have been provided by the Italian Treasury.

The main data source for our empirical microanalysis is the Bank of Italy survey on Household Income and Wealth (BIW). The BIW is a biannual household survey that covers all regions in Italy and contains a broad range of information on individual characteristics and economic performance. We use data from surveys in 1993 and 1995 that contain detailed information on socioeconomic factors relevant to our study.

The 1995 (1993) BIW survey provides information on 23,924 (24,013) individuals covering a total of 8,135 (8,089) households. A special feature of this survey is that it contains information on parents and children of the head of the household. This allows us to track intergenerational links (family ties) and relate them to public sector employment. In most of our analysis we restrict the sample to respondents between age 15 and 62 for men and 57 for women, the traditional standard age of retirement. Note that the BIW survey oversamples government employees by a factor of two, an issue that we discuss below.

Table 1 lists all the variables used in this paper and their sources. Table 2 provides sample statistics for some of the variables used in our empirical analysis of the 1995 BIW survey.

Imbalance in the Distribution of Public Jobs

For the purposes of discussion in this paper, we have divided Italy into three regions: North, Center, and South. As Table 3 shows, Italy has a pronounced mismatch between regional economic output and the use of its public resources. About 55 percent of total output is produced in the North, while only 44 percent of the total population resides there. Also, the South of Italy has considerably fewer labor force participants (51.5 percent compared with 62.5 percent in the North). The unemployment rate in the South (21.0 percent) is more than double
<table>
<thead>
<tr>
<th>Variable</th>
<th>Description</th>
<th>Year</th>
<th>Source</th>
</tr>
</thead>
<tbody>
<tr>
<td>Public employees</td>
<td>Total number of government employees including national and local employees</td>
<td>1995</td>
<td>Italian Treasury</td>
</tr>
<tr>
<td>Postal workers</td>
<td>Total number of postal workers</td>
<td>1995</td>
<td>Italian Treasury</td>
</tr>
<tr>
<td>Railroad workers</td>
<td>Total number of railroad workers</td>
<td>1997</td>
<td>Italian Treasury</td>
</tr>
<tr>
<td>Police</td>
<td>Total number of police employees</td>
<td>1995</td>
<td>Italian Treasury</td>
</tr>
<tr>
<td>Tax inspectors</td>
<td>Total number of tax inspectors</td>
<td>1996</td>
<td>Italian Treasury</td>
</tr>
<tr>
<td>Regional product</td>
<td>Regional state product</td>
<td>1995</td>
<td>ISTAT</td>
</tr>
<tr>
<td>Regional unemployment rate</td>
<td>Regional unemployment rate</td>
<td>1995</td>
<td>ISTAT</td>
</tr>
<tr>
<td>Regional public employment rate</td>
<td>Fraction of public employees in the regional labor force (excludes military, postal and railroad workers)</td>
<td>1995</td>
<td>ISTAT, Italian Treasury</td>
</tr>
<tr>
<td>Class size</td>
<td>Number of students per session</td>
<td>1995</td>
<td>ISTAT</td>
</tr>
<tr>
<td>Log hourly wages</td>
<td>Log of hourly disposable labor income</td>
<td>1995</td>
<td>BIW</td>
</tr>
<tr>
<td>High school degree</td>
<td>Highest degree: high school</td>
<td>1995</td>
<td>BIW</td>
</tr>
<tr>
<td>College degree</td>
<td>Highest degree: college</td>
<td>1995</td>
<td>BIW</td>
</tr>
<tr>
<td>Parent schooling</td>
<td>Years of schooling: head of household</td>
<td>1995</td>
<td>BIW</td>
</tr>
<tr>
<td>Business degree</td>
<td>Dummy: holding a business-type degree (for a definition, see section 4.3)</td>
<td>1995</td>
<td>BIW</td>
</tr>
<tr>
<td>Years work experience</td>
<td>Years of reported work experience</td>
<td>1995</td>
<td>BIW</td>
</tr>
<tr>
<td>Firm size: 20-99 employees</td>
<td>Dummy: reported number of employees</td>
<td>1995</td>
<td>BIW</td>
</tr>
<tr>
<td>Firm size: 100-499 employees</td>
<td>Dummy: reported number of employees</td>
<td>1995</td>
<td>BIW</td>
</tr>
<tr>
<td>Firm size: more than 500 employees</td>
<td>Dummy: reported number of employees</td>
<td>1995</td>
<td>BIW</td>
</tr>
<tr>
<td>White collar</td>
<td>Self-described employment type</td>
<td>1995</td>
<td>BIW</td>
</tr>
<tr>
<td>Teacher</td>
<td>Self-described employment type</td>
<td>1995</td>
<td>BIW</td>
</tr>
</tbody>
</table>
that in the Center (10.3 percent) and about three times higher than that of the North (6.7 percent).6

The regional differences in the distribution of public jobs are large. Public civilian employment per capita is higher in the South than in the North (about 61 public employees per thousand population in the South versus 51 in the North). As a share of total employment the difference is even more staggering: 12 percent of the employed in the North are in the public sector against 21 percent in the South. The comparison with the Center is clouded by the presence of the national capital in the Lazio region. Including this region, public employment is artificially high in the Center. For this reason we focus mostly on North-South comparisons.

Table 3 underestimates the differences between North and South for two reasons. First, it does not include employees of public and semipublic enterprises. Second, Wagner's Law implies that the size of government (and thus the number of public employees) increases with income per capita. Since the South is poorer than the North, Wagner's Law predicts a smaller government sector in this region.

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6Regional differences are so large that it seems surprising that there is no significant labor mobility from South to North. Cannari, Nucci, and Sestito (2000) show that mobility costs (i.e., housing cost of relocation) are very large and make geographical relocation too costly despite large differences in income.
Table 2. Descriptive Statistics of BIW

<table>
<thead>
<tr>
<th>Variable</th>
<th>Mean</th>
<th>Standard Deviation</th>
<th>Minimum</th>
<th>Maximum</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Region</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>North</td>
<td>0.448</td>
<td>0.497</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td>Center</td>
<td>0.196</td>
<td>0.397</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td>South</td>
<td>0.354</td>
<td>0.478</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td><strong>Household structure</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Parents (fraction)</td>
<td>0.640</td>
<td>0.479</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td>Children (fraction)</td>
<td>0.337</td>
<td>0.473</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td><strong>Demographics</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Age</td>
<td>36.73</td>
<td>13.01</td>
<td>15</td>
<td>62</td>
</tr>
<tr>
<td>Male</td>
<td>0.520</td>
<td>0.499</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td>Married(^1)</td>
<td>0.589</td>
<td>0.491</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td>Lives in city &gt; 500,000</td>
<td>0.136</td>
<td>0.343</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td>School</td>
<td>9.737</td>
<td>3.805</td>
<td>3</td>
<td>20</td>
</tr>
<tr>
<td>High school</td>
<td>0.363</td>
<td>0.480</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td>College</td>
<td>0.071</td>
<td>0.257</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td><strong>Employment status</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Unemployed</td>
<td>0.165</td>
<td>0.371</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td>Retired</td>
<td>0.227</td>
<td>0.418</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td>Labor force participation</td>
<td>0.634</td>
<td>0.481</td>
<td>0</td>
<td>1</td>
</tr>
</tbody>
</table>

Source: BIW (Bank of Italy, 1995).

\(^1\)Also includes unmarried people living with a partner.

Differences in the age structure of the population in the North and the South may account for different levels of employment in two large sectors: education and health. In fact, the fraction of the population below age 14 is higher in the South than in the North (12.4 percent in the North versus 18.8 percent in the South). On the contrary, the share of the population above age 65 is higher in the North than in the South (18.2 percent versus 13.8 percent). This implies that one should expect more health care employees in the North and more teachers in the South. As Table 3 shows, health care employees are just slightly more evident in the North while teachers are far more numerous in the South. Note that the large number of teachers is able to keep class size as low as in the North. This is a form of redistribution, since poorer regions with more children receive the same number of teachers per capita than wealthier regions with fewer children. Also, employment in public universities is higher in absolute numbers (and, a fortiori, in per capita terms) in the South than in the North.

In all the other categories, such as federal and regional administration, public employment per capita is higher in the South. Note that for some of the points we make below, on the effects of public employment on labor market structure and social attitudes, what matters most is the share of public employment in the labor force or relative to private employment. Evaluated in this way, public employment in the South is much higher than in the North in all the categories of employment.

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### Table 3. Regional Economic Performance and Public Employment

<table>
<thead>
<tr>
<th></th>
<th>North</th>
<th>Center</th>
<th>Center Without Lazio</th>
<th>South</th>
</tr>
</thead>
<tbody>
<tr>
<td>Regional product over national product (GDP)</td>
<td>55.1</td>
<td>20.5</td>
<td>10.6</td>
<td>24.3</td>
</tr>
<tr>
<td>Regional population over total population</td>
<td>44.4</td>
<td>19.3</td>
<td>10.1</td>
<td>36.4</td>
</tr>
<tr>
<td>Participation rate</td>
<td>62.5</td>
<td>59.7</td>
<td>62.3</td>
<td>51.5</td>
</tr>
<tr>
<td>Unemployment rate</td>
<td>6.7</td>
<td>10.3</td>
<td>8.2</td>
<td>21.0</td>
</tr>
<tr>
<td>Public employees per 100 residents</td>
<td>5.1</td>
<td>6.9</td>
<td>6.1</td>
<td>6.1</td>
</tr>
<tr>
<td>Public administration</td>
<td>0.64</td>
<td>1.53</td>
<td>0.88</td>
<td>0.87</td>
</tr>
<tr>
<td>Education and research</td>
<td>1.67</td>
<td>2.14</td>
<td>2.02</td>
<td>2.43</td>
</tr>
<tr>
<td>Regional administration</td>
<td>1.12</td>
<td>1.27</td>
<td>1.38</td>
<td>1.30</td>
</tr>
<tr>
<td>Health care</td>
<td>1.31</td>
<td>1.22</td>
<td>1.39</td>
<td>1.08</td>
</tr>
<tr>
<td>Other</td>
<td>0.35</td>
<td>0.75</td>
<td>0.41</td>
<td>0.42</td>
</tr>
<tr>
<td>Public employees per 100 employed</td>
<td>12.4</td>
<td>18.6</td>
<td>15.4</td>
<td>22.1</td>
</tr>
<tr>
<td>Public employees per unit of regional product²</td>
<td>124.0</td>
<td>194.4</td>
<td>155.0</td>
<td>275.1</td>
</tr>
<tr>
<td>Police officers per crime age population (15–65)</td>
<td>0.07</td>
<td>0.12</td>
<td>0.09</td>
<td>0.11</td>
</tr>
<tr>
<td>Police officers per 1000 crimes denounced³</td>
<td>7.3</td>
<td>12.3</td>
<td>13.4</td>
<td>10.7</td>
</tr>
<tr>
<td>Tax inspectors per unit of regional tax yield²</td>
<td>11.6</td>
<td>14.2</td>
<td>...</td>
<td>59.9</td>
</tr>
<tr>
<td>Postal workers per 100,000 units of correspondence⁴</td>
<td>72.9</td>
<td>122.5</td>
<td>134.1</td>
<td>108.3</td>
</tr>
<tr>
<td>Railways workers per 100,000 tons of goods shipped³</td>
<td>179.3</td>
<td>566.2</td>
<td>...</td>
<td>1782.7</td>
</tr>
<tr>
<td>Age structure: 15 and younger in population</td>
<td>12.4</td>
<td>13.2</td>
<td>12.1</td>
<td>16.1</td>
</tr>
<tr>
<td>Age structure: 65 and older in population</td>
<td>18.2</td>
<td>18.4</td>
<td>20.8</td>
<td>13.8</td>
</tr>
<tr>
<td>Class size⁶ (primary school)</td>
<td>16.2</td>
<td>16.9</td>
<td>15.9</td>
<td>18.0</td>
</tr>
<tr>
<td>Class size⁶ (secondary school)</td>
<td>20.7</td>
<td>20.5</td>
<td>20.5</td>
<td>21.0</td>
</tr>
</tbody>
</table>

Note: All data refer to 1995, unless otherwise indicated.

¹Employed and unemployed as a fraction of population between 15 and 65.
²Regional product and regional tax yield in Lit 100 billion. Taxes (collected in 1996) include value-added taxes, personal and corporate income tax, the so-called local tax on incomes (ILOR, abolished in 1997), and customs duties.
³Police officers in 1996 per 1,000 crimes denounced by the police in 1995.
⁴Number of post office employees per 100,000 letters and parcels sent in 1997.
⁵Railways workers in 1997 per 100,000 tons of goods shipped in 1996.
⁶Class size defined as students per session.
Other Factors: Productivity of Public Jobs and Quality of Public Service

Although it is difficult to measure the productivity of public employees, evidence suggests that the productivity of public employees in the South is lower than in the North.

Tax administration presents a striking picture. In 1996, about 25,000 tax inspectors in the North collected and administered Lit 213 trillion in taxes accruing to the central administration. While the number of staff devoted to the same tasks in the South was not significantly lower, taxes collected there amounted to only Lit 34 trillion. Hence, the average productivity of the staff employed in tax administration in the Northern regions was six times higher than in the South. Some of this striking difference can be explained by the fact that income per capita in the North is higher than in the South, so tax collected per number of taxpayers is higher. However, every indicator of tax evasion suggests that tax compliance is lower in the South, despite the large number of tax collectors.

Similarly, the regional concentration of personnel within the national post office and the railways cannot easily be attributed to differences in the demand for postal services and transportation. In the former sector, a Northern worker "produces" in a year ten times the annual output of her representative Southern colleague. In the transportation sector, the productivity gap—measured in manpower per units of goods shipped—while less extreme, is still large. We use goods rather than passengers because it is difficult to evaluate the role of transit passengers, traveling from a region to another through many other regions. Given the difficulties in measuring productivity in the public sector, care is needed in interpreting these data.

The indices of concentration of Italy's police per macroeconomic area reported on Table 3 are rather inconclusive. The higher density of officers charged with law enforcement in the Southern regions—with generally poorer records in terms of safety maintenance—reflects the government's objective to prevent criminal acts. The Southern ratio of police officers relative to the criminal age population is 51 percent higher than in the North. This higher ratio should therefore have a more significant deterrence effect on crime. The reported difference of law enforcement officers relative to reported crime, however, is only 47 percent larger in the South. This raises serious doubts about the effectiveness of a larger police force on crime deterrence. Also, aggregate ratios conceal remarkable disparities among single regions within the South that are nevertheless not easy to justify.

A different way of looking at productivity of public good provisions is to consider users' satisfaction. In the 1993 survey of the BIW, the heads of house-

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7We measure production of postal services in terms of the number of letters and parcels sent locally. If we were to include the number of withdrawals from and payments into postal checking and savings accounts, the productivity differentials would be even larger.

8The post office and the railways used to be administrations with the general government. The railway company became a stock company in 1992 and the post office was turned into an independent public agency in 1994. As a result the employees of both these entities are no longer employees of the general government.
holds were asked to report the use of local public services and to provide a qualitative evaluation of the services available. Table 4 compares the amount of public services used across regions. The residents in the North indicate a higher use of public transportation and health services. The Southerners, on the other hand, use more education and childcare facilities, which is consistent with the different age distribution in the two regions. Overall there appears to be no stronger reliance on public services in the South than in the North.

Table 5 reports the results from individual evaluations of different types of public services. In all the public functions (transportation, health services, education, and municipal services), residents in the North are more satisfied with the quality of local services. Obviously, these results should be taken cautiously given their qualitative nature. However, they are consistent with the evidence of Putnam (1993), who looks at several different measures of efficiency in different regions of Italy. For example, Putnam assessed the responsiveness and effectiveness of local bureaucracy in different regions of Italy by measuring processing time and quality in response to three specific information requests. In the most efficient regions (Emilia-Romagna and Valle d’Aosta, both in the North) two of three requests received thorough replies within a week. In the least efficient regions (Calabria, Campania, and Sardinia, all in the South), none of the requests received any attention and only direct inquiry and personal visits led to a response. A variety of other tests performed by this author reached similar conclusions. In fact, this widely cited book is entirely devoted to documenting and explaining the remarkable differences in public sector efficiency between the North and the South of Italy.

In summary, this evidence suggests that public employment is skewed in favor of the South without any benefit in terms of greater satisfaction for the public services provided or more frequent use of public services.

II. Socioeconomic Consequences of the Distribution of Public Employment

Wage Differentials

We begin by testing whether the public sector has a more equal payment structure across regions than the private sector. Data on earnings are taken from the B/W (Bank of Italy, 1995) and are based on reported monthly after-tax income. An important caveat is that since earnings are measured after tax, and given the progressivity of the tax system, this could understimate the North-South wage differential. An additional potential problem with income data is underreported income from nonmarket activities. Italy has a rather large gray economy that primarily supplements income of households in the South. The omission of this income source leads to overestimation of the North-South income gap and could bias the

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9 These questions were not asked in 1995.
Redistribution through public employment: the case of Italy

Table 4. Recent Use of Public Services

<table>
<thead>
<tr>
<th>Use of Public Services</th>
<th>North</th>
<th>Center</th>
<th>South</th>
</tr>
</thead>
<tbody>
<tr>
<td>Use of public transportation services</td>
<td>0.56</td>
<td>0.53</td>
<td>0.39*</td>
</tr>
<tr>
<td>Use of public health services</td>
<td>0.22</td>
<td>0.21</td>
<td>0.20</td>
</tr>
<tr>
<td>Medical tests in public laboratories</td>
<td>0.62</td>
<td>0.67</td>
<td>0.49*</td>
</tr>
<tr>
<td>Medical examinations (public)</td>
<td>0.52</td>
<td>0.48</td>
<td>0.41*</td>
</tr>
<tr>
<td>Use of medicines</td>
<td>0.81</td>
<td>0.84</td>
<td>0.79*</td>
</tr>
<tr>
<td>Nursery school attending</td>
<td>0.05</td>
<td>0.06</td>
<td>0.08*</td>
</tr>
<tr>
<td>Public primary, secondary school attending</td>
<td>0.20</td>
<td>0.27</td>
<td>0.31*</td>
</tr>
<tr>
<td>Public university attending</td>
<td>0.09</td>
<td>0.11</td>
<td>0.12*</td>
</tr>
</tbody>
</table>

Source: B/W (Bank of Italy, 1993).
Note: * indicates statistically significant differences of group means at 1 percent level.

Table 5. Quality of Public Services

<table>
<thead>
<tr>
<th>Quality Assessment</th>
<th>North</th>
<th>Center</th>
<th>South</th>
</tr>
</thead>
<tbody>
<tr>
<td>Public transportation functioning</td>
<td>6.09</td>
<td>5.45</td>
<td>4.52*</td>
</tr>
<tr>
<td>Health services functioning</td>
<td>6.03</td>
<td>5.21</td>
<td>4.00*</td>
</tr>
<tr>
<td>University functioning</td>
<td>6.31</td>
<td>5.79</td>
<td>4.76*</td>
</tr>
<tr>
<td>Municipality offices functioning</td>
<td>6.27</td>
<td>5.57</td>
<td>4.60*</td>
</tr>
<tr>
<td>Municipality street cleaning</td>
<td>6.20</td>
<td>5.70</td>
<td>4.52*</td>
</tr>
<tr>
<td>Public parks and gardens availability</td>
<td>6.11</td>
<td>5.53</td>
<td>3.68*</td>
</tr>
<tr>
<td>Public water quality</td>
<td>5.01</td>
<td>4.54</td>
<td>3.91*</td>
</tr>
<tr>
<td>Safety and crime control</td>
<td>5.91</td>
<td>5.70</td>
<td>4.02*</td>
</tr>
<tr>
<td>Nursery school functioning</td>
<td>7.16</td>
<td>6.76</td>
<td>5.38*</td>
</tr>
<tr>
<td>Primary and secondary school functioning</td>
<td>6.97</td>
<td>6.68</td>
<td>5.65*</td>
</tr>
</tbody>
</table>

Source: B/W (Bank of Italy, 1993).
Note: * indicates statistically significant differences of group means at 1 percent level.

Public-private income comparison. The latter problem may actually lead to an understatement of the public sector wage premium if public employees are more active in the gray market. This may be the case since reduced work hours and relaxed enforcement in public offices allow much time for second jobs in the gray economy.

In Table 6, column 1, we report estimates from standard wage regressions for public employees. In column 3 we run comparable regressions for the private sector. The dependent variable is the log of hourly earnings of fully employed workers and excludes self-employed workers (column 2). Hourly wages are obtained by dividing monthly earnings by 4.35, the average number of workweeks...
### Table 6. Wage Regressions for the Private and Public Sector  
(dependent variable: log hourly earnings from full-time employment)

<table>
<thead>
<tr>
<th></th>
<th>Public sector</th>
<th>Private sector</th>
<th>Private sector</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
</tr>
<tr>
<td><strong>Constant</strong></td>
<td>4.480</td>
<td>4.163</td>
<td>4.180</td>
</tr>
<tr>
<td></td>
<td>(105.2)</td>
<td>(159.23)</td>
<td>(156.18)</td>
</tr>
<tr>
<td><strong>High school degree</strong></td>
<td>0.063</td>
<td>0.095</td>
<td>0.091</td>
</tr>
<tr>
<td></td>
<td>(3.50)</td>
<td>(6.73)</td>
<td>(6.85)</td>
</tr>
<tr>
<td><strong>College degree</strong></td>
<td>0.246</td>
<td>0.252</td>
<td>0.242</td>
</tr>
<tr>
<td></td>
<td>(10.3)</td>
<td>(8.23)</td>
<td>(7.93)</td>
</tr>
<tr>
<td><strong>Years work experience</strong></td>
<td>0.037</td>
<td>0.042</td>
<td>0.041</td>
</tr>
<tr>
<td></td>
<td>(8.37)</td>
<td>(13.11)</td>
<td>(12.97)</td>
</tr>
<tr>
<td><strong>Years work experience</strong></td>
<td>-0.001</td>
<td>-0.001</td>
<td>-0.001</td>
</tr>
<tr>
<td></td>
<td>(-5.96)</td>
<td>(-9.41)</td>
<td>(-9.24)</td>
</tr>
<tr>
<td><strong>Female</strong></td>
<td>-0.105</td>
<td>-0.115</td>
<td>-0.106</td>
</tr>
<tr>
<td></td>
<td>(-7.66)</td>
<td>(-9.56)</td>
<td>(-8.52)</td>
</tr>
<tr>
<td><strong>Married</strong></td>
<td>0.064</td>
<td>0.100</td>
<td>0.101</td>
</tr>
<tr>
<td></td>
<td>(4.04)</td>
<td>(7.60)</td>
<td>(7.70)</td>
</tr>
<tr>
<td><strong>Center</strong></td>
<td>0.011</td>
<td>-0.070</td>
<td>-0.072</td>
</tr>
<tr>
<td></td>
<td>(0.66)</td>
<td>(-4.99)</td>
<td>(-5.21)</td>
</tr>
<tr>
<td><strong>South</strong></td>
<td>-0.014</td>
<td>-0.189</td>
<td>-0.192</td>
</tr>
<tr>
<td></td>
<td>(-0.99)</td>
<td>(-13.79)</td>
<td>(-14.00)</td>
</tr>
<tr>
<td><strong>White collar</strong></td>
<td>0.032</td>
<td>0.165</td>
<td>0.149</td>
</tr>
<tr>
<td></td>
<td>(1.63)</td>
<td>(11.00)</td>
<td>(9.49)</td>
</tr>
<tr>
<td><strong>Teacher</strong></td>
<td>0.355</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(13.94)</td>
<td></td>
<td></td>
</tr>
<tr>
<td><strong>Mid-management</strong></td>
<td>0.116</td>
<td>0.288</td>
<td>0.259</td>
</tr>
<tr>
<td></td>
<td>(4.00)</td>
<td>(11.53)</td>
<td>(10.14)</td>
</tr>
<tr>
<td><strong>Top management</strong></td>
<td>0.292</td>
<td>0.616</td>
<td>0.588</td>
</tr>
<tr>
<td></td>
<td>(7.21)</td>
<td>(13.32)</td>
<td>(12.69)</td>
</tr>
<tr>
<td><strong>Firm size: 20–99 employees</strong></td>
<td>0.114</td>
<td></td>
<td>0.110</td>
</tr>
<tr>
<td></td>
<td>(7.94)</td>
<td></td>
<td>(7.54)</td>
</tr>
<tr>
<td><strong>Firm size: 100–499 employees</strong></td>
<td>0.190</td>
<td></td>
<td>0.177</td>
</tr>
<tr>
<td></td>
<td>(11.36)</td>
<td></td>
<td>(10.26)</td>
</tr>
<tr>
<td><strong>Firm size: less than 500 employees</strong></td>
<td>0.275</td>
<td></td>
<td>0.247</td>
</tr>
<tr>
<td></td>
<td>(17.64)</td>
<td></td>
<td>(15.03)</td>
</tr>
<tr>
<td><strong>Industry dummies</strong></td>
<td>No</td>
<td>No</td>
<td>Yes</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td><strong>Adjusted R²</strong></td>
<td>40.4</td>
<td>47.0</td>
<td>47.8</td>
</tr>
</tbody>
</table>

Source: BIW (Bank of Italy, 1995).

Note: t-statistics in parentheses. Excluded category for work qualification is blue collar workers. Excluded category for industry dummies is manufacturing. The additional controls included in the regressions are the following dummies: invalid worker, sick worker, and big city—all statistically insignificant.
a month. We then divide this number by reported weekly hours including overtime.10

Focusing on the regional effects (the category left out is North), we find that public sector wages are not statistically different between the South and the North. On the contrary the results for the private sector are quite different. In column three we estimate the same wage regression for private employees. We focus again on the regional factors. Southern residents earn on average about 18.9 percent less than their Northern counterparts. This result is robust even after we take worker qualifications and industry structure into account. The other controls in the regression appear quite reasonable. Education implies a wage premium; years of work experience increase wages but at a decreasing rate. Females receive a lower wage even controlling for education and years of experience, and being married implies a wage premium.11

We now proceed to a more direct evaluation of the public sector wage premium in the North and South. Given the findings in Table 6, we expect that public employees in the South earn a sizable wage premium over private sector jobholders. Table 7 reports results from pooled (public and private) wage regressions of fully employed workers. Again we focus first on regional wage effects. Income from labor in the South is 13.6 percent lower than in the North. Also the first column of this table shows that at a national level public employment pays 19.0 percent more than the average private sector job.12 We now examine whether this premium differs by region.

In columns 2 and 3, we decompose this effect by estimating the public sector premium for the North and South. The public employment premium in the North is still positive but considerably smaller at 12.5 percent. By contrast, in the South of Italy we observe a public employment premium in excess of 26.0 percent over local private sector employment. A direct comparison of these two figures suggests that the Southern public sector wage premium is 13.5 percent ( = 26 - 12.5) higher in the South. One explanation is that the regional public wage premium may reflect a discrepancy in cost of living adjustments by the public and the private sector. While the private sector at least partly compensates for the lower cost of living, the public sector does not because of its nominal wage policy. Alternatively, the regional wage discrepancy between the private and public sector may be attributable to industry composition effects: the Southern private sector may predominantly operate in relatively low wage industries or in industries that do not negotiate wages at the national level. Part III of this paper contains a more detailed empirical analysis of the regional public wage premium. The key finding is that the public sector provides a substantial wage premium in the South, which is likely to have distortionary effects on the Southern labor market. The total effect is a combination of these two effects.

10The comparison between private and public sector wages may be slightly affected by the fact that overtime may be more widespread in the private sector.
11A wage premium on being married is commonly found in the labor literature; see, for instance, Polachek and Siebert (1993). The labor literature has discussed various alternative explanations of this finding.
12This result is consistent with the findings of Gregory and Borland (1999).
Table 7. Pooled Wage Regression: Private and Public Sector
(dependent variable: log hourly earnings from full-time employment)

<table>
<thead>
<tr>
<th></th>
<th>All Regions (1)</th>
<th>North (2) (1)</th>
<th>South (3)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>4.168</td>
<td>4.238</td>
<td>3.85</td>
</tr>
<tr>
<td></td>
<td>(182.1)</td>
<td>(147.22)</td>
<td>(78.5)</td>
</tr>
<tr>
<td>High school degree</td>
<td>0.213</td>
<td>0.189</td>
<td>0.271</td>
</tr>
<tr>
<td></td>
<td>(21.2)</td>
<td>(14.35)</td>
<td>(12.9)</td>
</tr>
<tr>
<td>College degree</td>
<td>0.507</td>
<td>0.443</td>
<td>0.631</td>
</tr>
<tr>
<td></td>
<td>(30.9)</td>
<td>(19.54)</td>
<td>(20.62)</td>
</tr>
<tr>
<td>Years work experience</td>
<td>0.048</td>
<td>0.046</td>
<td>0.055</td>
</tr>
<tr>
<td></td>
<td>(18.0)</td>
<td>(13.12)</td>
<td>(9.63)</td>
</tr>
<tr>
<td>Years work experience</td>
<td>-0.001</td>
<td>-0.001</td>
<td>0.001</td>
</tr>
<tr>
<td></td>
<td>(-11.9)</td>
<td>(-8.24)</td>
<td>(-6.67)</td>
</tr>
<tr>
<td>Female</td>
<td>-0.099</td>
<td>-0.088</td>
<td>0.091</td>
</tr>
<tr>
<td></td>
<td>(-10.3)</td>
<td>(-7.06)</td>
<td>(-4.44)</td>
</tr>
<tr>
<td>Married</td>
<td>0.108</td>
<td>0.074</td>
<td>0.183</td>
</tr>
<tr>
<td></td>
<td>(9.9)</td>
<td>(5.26)</td>
<td>(7.99)</td>
</tr>
<tr>
<td>Mid-management</td>
<td>0.094</td>
<td>0.119</td>
<td>0.057</td>
</tr>
<tr>
<td></td>
<td>(5.3)</td>
<td>(5.18)</td>
<td>(1.46)</td>
</tr>
<tr>
<td>Top management</td>
<td>0.21</td>
<td>0.290</td>
<td>0.137</td>
</tr>
<tr>
<td></td>
<td>(7.2)</td>
<td>(7.56)</td>
<td>(2.16)</td>
</tr>
<tr>
<td>Center</td>
<td>-0.073</td>
<td></td>
<td>(-6.2)</td>
</tr>
<tr>
<td>South</td>
<td>-13.1</td>
<td></td>
<td>(-12.2)</td>
</tr>
<tr>
<td>Public sector</td>
<td>0.190</td>
<td>0.125</td>
<td>0.260</td>
</tr>
<tr>
<td></td>
<td>(18.1)</td>
<td>(8.69)</td>
<td>(12.56)</td>
</tr>
<tr>
<td>Adjusted R²</td>
<td>44.3</td>
<td>41.6</td>
<td>50.9</td>
</tr>
</tbody>
</table>

Source: B/I/W (Bank of Italy, 1995).
Note: t-statistics are in parentheses. Additional controls included in the regressions are the following dummies: invalid worker, sick worker, and big city worker—all statistically insignificant.

Family Persistence of Public Sector Jobs

We now examine whether there is a tendency for members of the same family to be in the public sector: that is, we ask whether family ties to the public sector matter. This is interesting for two reasons. First, if family ties matter, they may indicate that a “culture” of public jobs is diffused in a family. A child raised in a culture of public job security may aspire to the same type of career. Furthermore, if these cultural effects are important, they may spill beyond the immediate family to a network of connected individuals. Second, if family connection matters, it could mean that it is easier to obtain a public job if a family member can help you get one through personal contacts, inside information, recommendations, or favors.

We begin by exploring the influence of the employment history of other family members on the likelihood of public sector employment. We compare the frequency of public sector employment between two groups of individuals:
workers with ties to the public sector, and workers without ties. Table 8 considers two types of family ties: between spouses and child-parent ties. In the latter category we can distinguish between two types of children-parents ties: ties between the head of the household and his or her parents, and the head of the household and his or her children. The spousal tie is affected by a serious problem of reverse causation and may therefore be biased; in fact, individuals may meet in the workplace and then marry.

Table 8 reveals how important family ties are for public employment. Children of public sector employees are almost twice as likely to end up in the public sector, relative to the others. The effect of spousal ties is strong, but perhaps not easily interpretable. Note that the effects of family ties are prevalent in the North, South, and Center.

In addition to being less prone to participate in the labor force in general, children of civil servants appear to have longer unemployment spells if unable to find a public position in the first place (see Table 9). The conditional probability of household members—aged 26 to 40—remaining unemployed if not hired by government tends to rise (by 5.9 percent13) with the number of close family relatives—parents and grandparents—who serve, or have served, as bureaucrats. In an earlier version of this paper, we also investigated school choices of public employees (and their children) relative to those employed in the private sector. We found weak evidence that children of public employees make choices that appear less “business oriented” than their counterparts in the private sector.

**Job Search**

Our hypothesis is that job searches in the South are mainly directed toward the public sector, but we cannot observe the direction of worker’s job search efforts. We can, however, observe the on-the-job search effort of different worker types, namely, public and private sector employees.

Table 10 (columns 1 and 3) reports Logit estimates of job search efforts controlling for the employing sector and regional level of the unemployment rate. The dependent is a dummy variable indicating whether a person has been looking for a job in the recent past.14 As shown in column 1, holding a job in the public sector (variable Public) significantly reduces the search effort. This suggests that public sector jobs are secure and provide a very high level of satisfaction. Several explanations for this finding are possible. The most obvious is that the high wage premium for public jobs in the South discourages anyone from looking elsewhere. Also, the workload may be even lower in the public sector than in comparable private sector jobs.

Finally, column 3 shows whether the Southern residents search less than Northerners when they hold a public sector job. The interaction variable Public x South is borderline (in)significant at the 10 percent level. This may suggest that Southerners are searching less on the job than their Northern counterparts. This

---

13This calculation is based on average marginal effect derived from Logit estimates.
14The sample is restricted to fully employed workers.
Table 8. Family Ties in the Public Sector: Frequency of Public Employment with Family Ties to the Public Sector  
(in percent of all employees)

<table>
<thead>
<tr>
<th>Family ties</th>
<th>Child-Parent Ties</th>
<th></th>
<th>Spousal Ties</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Head of Household</td>
<td></td>
<td>Head of Household</td>
</tr>
<tr>
<td></td>
<td>Yes</td>
<td>No</td>
<td>Yes</td>
</tr>
<tr>
<td>North</td>
<td>35.6</td>
<td>17.9</td>
<td>13.1</td>
</tr>
<tr>
<td>Center</td>
<td>42.8</td>
<td>25.1</td>
<td>21.9</td>
</tr>
<tr>
<td>South</td>
<td>48.8</td>
<td>23.9</td>
<td>33.9</td>
</tr>
</tbody>
</table>

Source: BIW (Bank of Italy, 1995).

Family ties to the public sector are defined as having one or more immediate family member (parent or spouse) who holds or held a job in the public sector. Sample weights applied. All group mean differences are statistically significant.

Evidence, although not very strong, is consistent with the result that public sector jobs are particularly valued in the South.15

Entrepreneurship

Does a large dependence on public employment deter the development of entrepreneurial activity? This issue is particularly important for Italy, since its economy relies more than other OECD countries on small business activities.16 We identify four categories of entrepreneurial activity: (1) professionals; (2) business owners; (3) independent workers or craftsmen; and (4) owners or assistants in family businesses.

Table 11 reports the empirical results from Logit estimation where the dependent variable is a binary indicator, equal to 1 if a respondent pursues an entrepreneurial activity (as defined above) and zero otherwise.17 In regression 1, we estimate entrepreneurship as a function of the level of schooling, work experience, and two regional variables: the regional unemployment rate and the fraction of public employees in the labor force. We find that education increases the likelihood of entrepreneurship, while the regional variables have no significant impact. The level of public employment has a negative sign but its effect is insignificant. In column 2 we add information on the regional economic performance (regional output over regional population). We find that residents of a highly productive region are less likely to undertake an entrepreneurial activity. Probably, low economic activity encourages self-employment, owing to a lack of

15We have also explored whether the type of education of the workers—business or more business—affected their search effort. We did not find significant effects.
16An OECD study (1995) reports that Italy has a large number of small- and medium-sized firms in its core industries. About 36.8 percent of all employees in Italy work in firms with less than 200 employees compared with 20.8 percent in Germany, 25.8 percent in France, and 34.1 percent in Japan.
17The sample is limited to heads of households older than age 20.
alternative employment opportunities. More interesting in this model is the public employment effect. The estimated coefficient on public employment is negative and significant at the 5 percent level. Thus, even though a lower productivity level on average would tend to encourage entrepreneurship, the presence of a large public sector tends to offset this effect. This result is robust to controlling for the family background of the respondent, as shown in the last column of Table 11.

III. The Size of Regional Redistribution Through Public Employment

We now evaluate the size of the regional redistribution obtained through public employment. We distinguish between two components: the quantity effect \( Q \), namely, the “excess” number of public jobs; and the price effect \( P \), that is, the “wage premium” paid to public employees in the South, to be defined below. Our
Table 10. Job Search Activity in Private and Public Sector (Logit)
(dependent variable: job search: 1 = yes, 0 = no)

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>-0.955</td>
<td>-1.250</td>
<td>-1.286</td>
</tr>
<tr>
<td></td>
<td>(-4.27)</td>
<td>(-5.00)</td>
<td>(-5.12)</td>
</tr>
<tr>
<td>Years of schooling</td>
<td>0.005</td>
<td>0.005</td>
<td>0.006</td>
</tr>
<tr>
<td></td>
<td>(0.34)</td>
<td>(0.31)</td>
<td>(0.37)</td>
</tr>
<tr>
<td>Years work experience</td>
<td>-0.062</td>
<td>-0.063</td>
<td>-0.062</td>
</tr>
<tr>
<td></td>
<td>(-10.68)</td>
<td>(-10.77)</td>
<td>(-10.72)</td>
</tr>
<tr>
<td>Female</td>
<td>-0.093</td>
<td>-0.079</td>
<td>-0.077</td>
</tr>
<tr>
<td></td>
<td>(-0.87)</td>
<td>(-0.74)</td>
<td>(-0.72)</td>
</tr>
<tr>
<td>Public</td>
<td>-1.406</td>
<td>-1.421</td>
<td>-1.244</td>
</tr>
<tr>
<td></td>
<td>(-8.72)</td>
<td>(-8.80)</td>
<td>(-6.67)</td>
</tr>
<tr>
<td>Public x South</td>
<td>-0.531</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(-1.65)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Center</td>
<td>-0.002</td>
<td>-0.108</td>
<td>-0.115</td>
</tr>
<tr>
<td></td>
<td>(-0.01)</td>
<td>(-0.77)</td>
<td>(-0.82)</td>
</tr>
<tr>
<td>South</td>
<td>0.193</td>
<td>-0.393</td>
<td>-0.312</td>
</tr>
<tr>
<td></td>
<td>(1.60)</td>
<td>(-1.52)</td>
<td></td>
</tr>
<tr>
<td>Regional unemployment rate</td>
<td>4.442</td>
<td>4.451</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(2.62)</td>
<td>(2.63)</td>
<td></td>
</tr>
<tr>
<td>Log Lik</td>
<td>-1408.1</td>
<td>-1404.5</td>
<td>-1403.1</td>
</tr>
</tbody>
</table>

Source: B/W (Bank of Italy, 1995).
Note: r-statistics are in parentheses. Model 4 restricted to workers with at least a high school degree.

An estimate of the implicit interregional transfer (TR) through public employment is given by:

\[
TR = (E_C - E_B)W_C + E_B(W_C - W_B) = QW_C E_C + PW_C E_B, \tag{1}
\]

where, \( Q = (E_C - E_B)/E_C \) and \( P = (W_C - W_B)/W_C \).

In equation (1) TR is the implicit monetary value of the interregional transfer, \( E_C \) is the current number of public employees in the South, \( E_B \) a numerical benchmark to be specified below, \( W_C \) the average wage rate currently paid to public employees in the South, and \( W_B \) a benchmark wage to be defined below.\(^{18}\) We call the expression \( Q = (E_C - E_B)/E_C \) the quantity effect and \( P = (W_C - W_B)/W_C \) the...
price effect. After rearranging of terms and substituting for $E_B$, we obtain the following expression for $TR$:

$$TR = W_c E_c Q + (1 - Q) P$$ \hspace{1cm} (2)$$

The main task in determining the quantity effect is to construct a baseline rule for the level of public employment, which is to identify $E_B$. Since it is not obvious how to do this, we offer several different estimates for $E_B$. Also, to highlight the North-South comparison, we use only the characteristics of the North as the determinants for the baseline scenario for the South.

In order to estimate the price effect, $P$, we need to compute $W_B$, the “benchmark” salary for public employees in the South. This can be tackled from two angles. First, one can define the benchmark in terms of the wage rate payable if the public wage policy were to conform to the norm of equalizing regional public compensations in real terms. Second, one can construct an institutional counter-
factual and ask which wage rate would be paid in the South if Italy were not a unified country. A natural candidate for \( W_R \) would be, in this case, a measure of the nominal wage prevailing in the private sector to remunerate labor of comparable quality. We pursue both strategies below.

**The Quantity Effect**

**One-dimensional baseline estimates**

The simplest approach for the calculation of \( Q \) is to assume that the South should have the same level of public employment per unit of a particular characteristic, that is, regional attribute, as the North. We present four alternative attributes: the size of the labor force, employment, regional output, and the regional level of consumption. The baseline estimates for the North are then calculated as the ratio of northern public employment over the specific regional characteristics. It is not obvious what is the "best" measure, and therefore we present a menu of them. For example, from the point of view of economic efficiency alone, public employment should be roughly a constant fraction of the population across regions; in this case, the population weight is the correct one. If, instead, economic efficiency is seen to suggest that the public sector should be a certain fraction of GDP, then output weights or employment weights are preferable.

In Table 12 (first row) we report several different baseline estimates derived from northern regional observations. The baseline estimate in column 3, for example, is obtained by dividing the Northern public employment by total Northern employment, and it implies a baseline fraction of 12.1 percent of the employed population.

All Southern regions have excess public employment according to all four baseline rules. In the second row of Table 12 we report estimates of excess public employment for the South as a whole. Predicted excessive employment is measured as a fraction of the respective Southern regional public employment. The smaller estimates of excess employment are based on the population. Estimates are higher when we take the full economic disparities into account. The employment- and output-based rules imply that 43 percent to 55 percent of the public employment in the South is above the baseline limits. Two regions in South—Molise and Basilicata—have the highest “excessive” public employment.

**Multidimensional estimates: Wagner regressions**

The estimates just discussed do not take into account that different characteristics of the economies in the Northern and Southern regions may "require" different levels of public employment, purely for economic reasons. Any attempt to determine the "optimal size" of public employment on a regional base should be undertaken with caution. Here we estimate a regional model of public employment in the spirit of Wagner’s Law.\(^{19}\) We construct a provincial data set derived from residential informa-

\(^{19}\)There is a rich empirical literature testing the time-series implications of this proposition. Relatively little research has been conducted using cross sectional information. One exception is Eberts and Gronberg (1992).
tion in the Bank of Italy survey (1993 and 1995). Each individual in the survey can be identified by his or her province of residence and reported sector of employment (that is, private versus public). This information allows us to develop provincial attributes by calculating provincial population averages. In total we obtain information on 99 provinces that can be mapped to the 20 main regions in Italy. To maximize the degrees of freedom, we expand our data set and merge data from the 1993 survey with the 1995 survey. We end up with an average of 484 observations for each province: the lowest number of observations is 36, and the maximum is 3,135. All provincial characteristics are then derived by calculating weighted means of individual observations using population weights from the survey. Income variables are expressed in 1993 Lire. Given the data problems and the way we have to construct these provincial data, the results of these Wagner regressions should be taken as suggestive and indicative.

In Table 13, we report the regression results for the Northern provinces, where the dependent variable is the fraction of public employment over total employment. The model with the best fit is the more narrowly defined Northern model. All the control variables point in the expected direction even though standard errors are high. For instance, higher levels of income lead—albeit with weak evidence—to more public employment. As additional determinants of public employment, we used information on the employment structure of a province, the fraction of old (older than 65) and young residents (younger than 15), and the degree of urbanization. We also use as a control the fraction of employment in the service industries. In an earlier version of this paper, we explored alternative specifications, including restricting the definition of what is included in the Northern regions. The results do not change much; in fact, they often improve.

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20 Among these provinces, 43 are part of the North and 35 are part of the South. The number of provinces in each region varies from 1 to 10.

21 Since the 1993 and the 1995 surveys contain the same questions, we do not need to modify the variables of interest.

---

**Table 12. One-Dimensional Baseline Rules and Excess Public Employment in the South**

<table>
<thead>
<tr>
<th>Baseline Rules</th>
<th>Population (1)</th>
<th>Labor force (2)</th>
<th>Employment (3)</th>
<th>Regional product (4)</th>
<th>Regional consumption (5)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Baseline rule</td>
<td>0.05</td>
<td>0.113</td>
<td>0.121</td>
<td>1.636</td>
<td>2.331</td>
</tr>
<tr>
<td>Predicted excess public employment in South (percent)</td>
<td>19.6</td>
<td>32.2</td>
<td>43.5</td>
<td>54.9</td>
<td>37.6</td>
</tr>
</tbody>
</table>

Sources: Italian Treasury (1996); and ISTAT (1996a).

1 Measured in billions of Lire.
Table 13. Wagner Regression of Public Employment for Northern Provinces

<table>
<thead>
<tr>
<th>Variable</th>
<th>Coefficient</th>
<th>t-statistic</th>
</tr>
</thead>
<tbody>
<tr>
<td>Log of total income</td>
<td>-1.929</td>
<td>(-1.68)</td>
</tr>
<tr>
<td>Service sector employment among private sector</td>
<td>0.2102</td>
<td>(0.92)</td>
</tr>
<tr>
<td>Dependency rate</td>
<td>0.3323</td>
<td>(1.36)</td>
</tr>
<tr>
<td>Old people living without family members</td>
<td>0.3055</td>
<td>(1.14)</td>
</tr>
<tr>
<td>Urbanization rate</td>
<td>0.0564</td>
<td>(1.8)</td>
</tr>
<tr>
<td>Adj. R²</td>
<td>0.18</td>
<td></td>
</tr>
</tbody>
</table>

Number of observations: 43

Source: B/W (Bank of Italy, 1993 and 1995) plus regional data from ISTAT.
Note: t-statistics are in parentheses. For definition of regions see text.

We now turn to the estimation of public employment in the South relative to the benchmark. We obtain predicted levels of public employment by using the estimates from Table 13. Before we can derive these estimates, however, we need to tackle two issues. First, the level of income in the South has not been adjusted for differences in the cost of living. If we evaluate the Wagner model for the North at the nominal level of income in the South we would overestimate the real income differences and thus overestimate the level of excessive public employment. This is because the price level in the South is lower, and therefore the nominal income of the South underestimates the real income of the South relative to the North. We correct for this difference by increasing the level of income by 15 percent and, more extremely, by 25 percent.²² Note that these corrections will lead to lower estimates of excessive public employment. A second issue is the size difference of the Southern provinces. To obtain a combined estimate for total Southern exces-

²²In the next section we provide estimates of the difference in cost of living, which are consistent with the range of these adjustments.
sive public employment, we need to weight the provincial predictions before we can add them up. We do this by applying provincial weights derived from population weights from the BIW 1993 and 1995.

Table 14 summarizes the results of this exercise. The multivariate model predicts an excessive employment rate between 38 percent and 43 percent. The regions with the highest levels of excessive public employment according to this measure are Campania, Puglia, and Calabria, and they differ from the one-dimensional estimates. Finally, regional per capita income and the rate of excessive public employment are negatively correlated with a coefficient of $(-0.21)$. This observation hints at the use of public employment as a redistributive device.

To sum up, one-dimensional estimates of excessive public employment lie between 20 percent–55 percent of total public employment in the South. The multivariate estimates have a smaller range and lie between 38 percent and 43 percent.

The Price Effect

**Price Effect 1: cost of living adjustment**

First we estimate by how much the public sector would have to reduce wages in the South in order to equalize pay in real terms, between the North and the South. This is a simple measure of the implicit subsidy that is due to the higher real wage for public employees in the South.

Our estimates for regional price differences are derived from cost of living data for Italian cities. As mentioned above, ISTAT, the Italian statistical agency, does not provide price level indices for different regions. We therefore use data of city price deflators for the period 1947–95 to calculate the cumulative price divergence between the North and the South. We assume that the cost of living difference between the North and the South was small at the beginning of the period. We use data from six northern and seven southern cities. The accumulated difference of the average price index amounts to 14.3 percent by 1995. We also derived alternative measures of real income differences from wage regressions of the private sector. Our estimates are very similar and range between 15–18 percent. These regression results are available from the authors upon request. This measure of price level difference probably underestimates the extent of the higher cost of living in the North. Cannari, Nucci, and Sestito (2000), looking at real estate prices, suggest much larger differences between the North and South.

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23Cost of living data are available for the following cities: North: Turin (Piedmont), Genoa (Liguria), Trento (Trentino), Triest (Friuli), Bologna (Emilia-Romagna), and Venezia (Veneto); South: Campo Basso (Molise), Napoli (Campania), Bari (Puglia), Potenza (Basilicata), Reggio Calabria (Calabria), Palermo, and Catania (Sicily).
Price Effect 2: adjustment to public-private pay structure

An alternative way to calculate the price effect is to ask which wage rate would be paid in the South region if Italy were not a unified country. A natural determinant for the Southern baseline wage, $w_B$, would be the wage rate that generates the same public-private sector pay structure as in the North. This comparison would not only account for differences in the cost of living but also take into account regional differences in productivity.

We run two types of wage regressions, one for the North in order to determine the base public-private sector wage structure, and one for the South (see Table 15 for the results). We assume that public sector work is similar to the service sector, and therefore that the public-private sector wage comparison should focus on the service sector. We use the following service sector industries: banking and insurance, real estate, and personal services. However, since wages in the banking and insurance sector are to a large extent set on the national level, the Southern public-private sector comparison is somewhat biased. For this reason we run a separate set of regressions specifically controlling for the banking and insurance sector from the private service sector.

To determine the Northern pay structure, we run a pooled (public and private) wage regression for fully employed Northern residents only. We use the same control variables as above but also include a dummy for employment in the private service sector. We exclude public sector employment. We find that in the North the private service sector pays on average 7.2 percent (column 1) less than the public sector. When we leave out the banking sector the differential is substantially larger at 18.0 percent. We can now compare this finding with estimates from the South (Table 12, columns 3 and 4). The wage differential between public and private service sector employees is much larger. On average, public employees earn about 24.9 percent more than their private sector counterparts if we include the banking sector, and a stunning 40.5 percent more if we exclude the banking sector.

By what amount would the Southern wage have to be adjusted to achieve the Northern pay structure? If the Northern public-private pay structure were to
## Table 15. Wage Regressions for the North and South: Public-Private Pay Structure
(dependent variable: log hourly earnings from full-time employment)

<table>
<thead>
<tr>
<th></th>
<th>Log hourly wages</th>
<th>North (1)</th>
<th>South (2)</th>
<th>North (3)</th>
<th>South (4)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td></td>
<td>4.373</td>
<td>4.382</td>
<td>4.109</td>
<td>4.133</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(128.4)</td>
<td>(126.6)</td>
<td>(75.3)</td>
<td>(75.4)</td>
</tr>
<tr>
<td>High school degree</td>
<td></td>
<td>0.184</td>
<td>0.174</td>
<td>0.269</td>
<td>0.245</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(13.9)</td>
<td>(12.9)</td>
<td>(12.8)</td>
<td>(11.5)</td>
</tr>
<tr>
<td>College degree</td>
<td></td>
<td>0.438</td>
<td>0.427</td>
<td>0.629</td>
<td>0.602</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(19.3)</td>
<td>(18.3)</td>
<td>(20.5)</td>
<td>(19.2)</td>
</tr>
<tr>
<td>Years work experience</td>
<td></td>
<td>0.045</td>
<td>0.045</td>
<td>0.054</td>
<td>0.052</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(13.1)</td>
<td>(12.9)</td>
<td>(9.6)</td>
<td>(9.2)</td>
</tr>
<tr>
<td>Years work experience</td>
<td></td>
<td>-0.001</td>
<td>-0.001</td>
<td>-0.001</td>
<td>-0.00</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(-8.3)</td>
<td>(-8.1)</td>
<td>(-6.6)</td>
<td>(-6.3)</td>
</tr>
<tr>
<td>Female</td>
<td></td>
<td>-0.094</td>
<td>-0.081</td>
<td>-0.092</td>
<td>-0.074</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(-7.5)</td>
<td>(-6.3)</td>
<td>(-4.4)</td>
<td>(-3.6)</td>
</tr>
<tr>
<td>Married</td>
<td></td>
<td>0.075</td>
<td>0.072</td>
<td>0.183</td>
<td>0.185</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(5.3)</td>
<td>(5.0)</td>
<td>(7.9)</td>
<td>(8.0)</td>
</tr>
<tr>
<td>Mid-management</td>
<td></td>
<td>0.111</td>
<td>0.099</td>
<td>0.055</td>
<td>0.025</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(4.8)</td>
<td>(4.0)</td>
<td>(1.4)</td>
<td>(0.6)</td>
</tr>
<tr>
<td>Top management</td>
<td></td>
<td>0.283</td>
<td>0.265</td>
<td>0.136</td>
<td>0.135</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(7.4)</td>
<td>(6.6)</td>
<td>(2.1)</td>
<td>(2.0)</td>
</tr>
<tr>
<td>Nonservice sector</td>
<td></td>
<td>-0.139</td>
<td>-0.141</td>
<td>-0.263</td>
<td>-0.270</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(-9.3)</td>
<td>(-9.4)</td>
<td>(-11.7)</td>
<td>(-12.1)</td>
</tr>
<tr>
<td>Service sector</td>
<td></td>
<td>-0.072</td>
<td>-0.249</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td>(-3.4)</td>
<td>(-7.9)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Service sector without banks and insurance</td>
<td></td>
<td>-0.180</td>
<td></td>
<td>-0.405</td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td>(-7.2)</td>
<td>(-10.8)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>R²</td>
<td></td>
<td>41.9</td>
<td>41.3</td>
<td>50.9</td>
<td>52.4</td>
</tr>
</tbody>
</table>

Source: B/W (Bank of Italy, 1995).

Notes: t-statistics are in parentheses. Excluded industry category is public employment. The service sector consists of banking and insurance, real estate, and personal services. The nonservice sector consists of agriculture, manufacturing, telecommunications, construction, and transportation. Additional controls included in the regressions: invalid worker, sick worker, and big city worker—all statistically insignificant.
prevail in the South, our estimates indicate that public wages would have to adjust downward by 17.7 percent. These estimates are slightly lower than the cost of living estimates. If we drop the banking sector from our comparison, the wage gap increases to 15.6 percent. Again we find that the estimated wage adjustments are similar to our previous results.

The Total Cost of Excessive Public Employment

We are now able to provide an estimate of the cost of excessive public employment. We recall from earlier that interregional transfer cost is defined as:

\[ TR = W_C E_C (Q + (1 - Q)P), \]  

where \( W_C E_C \) is the current expenditure on public employment in the South and \( Q \) and \( P \) are the quantity and price effects, respectively. The earlier results indicate that the excessive rate of public employment in the South, \( Q \), lies between 20 to 55 percent with a more narrow range of 38-44 percent from the Wagner estimates. On the other hand, the price effect, \( P \), which measures the excessive payment levels, ranged roughly from 11 to 18 percent. We can use these two pieces of information and calculate the combined effect as described in equation (3) above. The total effect ranges between a minimum at 30 percent and a maximum at 65 percent of the public sector wage bill for the South. Taking the middle range of the Wagner estimates and the middle range for the price effect, we get a value of almost exactly 0.5.

IV. Conclusion

The allocation of public employment in Italy is an important source of geographical redistribution between regions, in particular between the North and the South. About half of the wage bill of the South can be thought of as redistributive, that is, in excess of what it “should be” relative to various ways of calculating a benchmark. This amount is the result of a quantity and a price effect. The former is due to the fact that there are many more public employees in the South relative to the North; the second arises because, while public wages are very similar across regions, the price level instead is lower in the South, so that real wages are higher in the South.

The heavy reliance on attractive public jobs in the South leads to a vicious circle in which private sector jobs are not sought after. This also implies that for private entrepreneurs it is expensive to offer jobs as attractive as those offered by the public sector. The result is that the economy in the South is overly dependent on public jobs that are of the nature of permanent welfare. The problem is compounded by the use (and misuse) of disability pensions, which are also concentrated in the South and are in many cases another source of permanent unemployment compensation.
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How Does U.S. Monetary Policy Influence Sovereign Spreads in Emerging Markets?

VIVEK ARORA and MARTIN CERISOLA*

This paper quantifies the impact of changes in U.S. monetary policy on sovereign bond spreads in emerging market countries. Specifically, the paper explores empirically how country risk, as proxied by sovereign bond spreads, is influenced by U.S. monetary policy, country-specific fundamentals, and conditions in global capital markets. While country-specific fundamentals are important in explaining fluctuations in country risk, the stance and predictability of U.S. monetary policy are also important for stabilizing capital flows and capital market conditions in emerging markets. [JEL E43, F36, G15]

The increased globalization of the world economy over the past decade has been reflected in the increased dependence of emerging markets on developments in the U.S. economy. While the dramatic rise in capital flows to emerging markets has been induced primarily by the implementation of sound macroeconomic policies and wide structural reforms in these countries, it has also been driven by changing conditions in industrial countries that have encouraged investors to diversify their portfolios into developing country assets. In particular, Calvo, Leiderman, and Reinhart (1993) have emphasized the role of economic conditions—particularly interest rates—in industrial

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countries,\(^1\) while others have also pointed to structural changes in institutional portfolios in industrial countries, which led to a permanent increase in their exposure to developing countries.

The resumption of capital flows to developing countries during the 1990s was accompanied by a dramatic decline in interest rate spreads but increased countries’ vulnerability to sudden reversals in investors’ confidence and increased turbulence. Some past episodes of market turbulence occurred at the same time that the stance of U.S. monetary policy was being changed considerably (for example, in 1994) or even precipitated changes in U.S. monetary policy (for example, during the second half of 1998). Given the integration of global capital markets, changes in U.S. monetary policy have been felt by developing countries through effects on the cost and availability of funds, and on their creditworthiness.

In addition to the direct impact of changes in U.S. interest rates on rates in developing countries, interest rate spreads (the differences between yields on sovereign bonds of developing countries and U.S. treasury securities of comparable maturities), which are a proxy for country risk, have tended to move in the same direction as the changes in U.S. interest rates. This effect on developing country spreads was seen clearly in 1994 when a tightening of U.S. monetary policy was reflected in a substantial widening of spreads, and in 1998, when an easing of U.S. monetary policy in response to the flight to quality and the concerns about a U.S. credit crunch associated with the Russian default and the near demise of Long-Term Capital Management (LTCM) helped to restore global liquidity conditions and to reduce sovereign spreads somewhat.

This paper presents empirical evidence on how changes in U.S. monetary policy influence country risk in several developing countries in Latin America, Asia, and Eastern Europe. In particular, we examine empirically how country risk, as proxied by sovereign bond spreads, is influenced by U.S. monetary policy, country-specific fundamentals, and by conditions in world capital markets.

I. What Drives Sovereign Spreads in Emerging Markets?

From a theoretical perspective, a rise in U.S. policy interest rates could lead to an increase in emerging market spreads for several reasons.\(^2\) To the extent that emerging market bonds are risky (there is a probability of default), the yield on emerging market bonds would have to rise by more than any rise in the risk-free rate. To illustrate, if \(r\) and \(i\) represent the interest rate on the risk-free asset and

---

\(^1\) Calvo, Leiderman, and Reinhart (1993) note that flows to Latin America and developing countries in general, during the early 1990s were triggered by "...falling interest rates, a continuing recession, and balance of payments developments in the United States..."

\(^2\) See Kamin and von Kleist (1999) for further discussion.
the risky asset, respectively, and \( p \) is the probability of repayment on the risky asset, then the equilibrium condition is:

\[
(1 + r) = p \cdot (1 + i) + (1 - p) \cdot 0.
\]

(1)

The interest rate spread, \( S \), defined as the difference between the rate on the risky asset and on the risk-free asset, in equilibrium is then:

\[
S = \frac{(1 + r)(1 - p)}{p}.
\]

(2)

and its derivative with respect to \( r \) is \((1 - p)/p\), which is positive as long as \( p < 1 \). This says that as long as there is some risk of default, the rate on the risky asset will have to rise by more than any rise in the risk-free rate in order to compensate investors for the risk.

A rise in U.S. rates could also raise emerging market spreads through its effects on the ability of debtor countries to repay loans. A rise in U.S. rates would tend to increase debt-service burdens in borrowing countries, which would reduce their ability to repay loans. In addition, as noted by Kamin and Kleist (1999), a rise in U.S. rates could reduce investors' appetite for risk, leading them to reduce their exposure in risky markets, in turn reducing available financial resources in borrowing countries. In terms of the above illustration, if the probability of repayment is a negative function of the risk-free rate \( (p = p(r), \text{ with } p' < 0) \), then the first derivative of \( S \) with respect to \( r \) is:

\[
dS/dr = \left[\frac{(1 - p)}{p}\right] - \left[\frac{(1 + r)p'}{p^2}\right],
\]

(3)

which is positive (since \( p < 1 \) and \( p' < 0 \)). This says that a rise in the risk-free rate raises the spread both because of the risk of default (the first term) and because that risk rises as the risk-free rate goes up (the second term).

A number of relatively recent papers have explored the question of how emerging market spreads are determined, including the role of macroeconomic fundamentals and changes in market sentiment. Notwithstanding the straightforward theoretical prediction and ample anecdotal evidence, the empirical literature on how U.S. monetary policy has affected emerging market spreads is less conclusive. Most of these analyses have tended to explore the role of global liquidity conditions, as proxied by a specific yield on a U.S. treasury security, on sovereign bond spreads. For example, in a study of 11 emerging market countries, Cline and Barnes (1997) found a positive but statistically insignificant effect of U.S. treasury yields on sovereign spreads during the mid-1990s. Kamin and von Kleist (1999) found no
statistically significant relationship between U.S. treasury rates and spreads for selected emerging market countries, with the correlation being negative in some cases. Eichengreen and Mody (1998a and 1998b) took these analyses further by explicitly analyzing demand and supply factors in the market for emerging market bonds. They found, for a sample of Latin American and East Asian countries during the early 1990s, that a rise in U.S. treasury interest rates tended to reduce spreads, and at the same time reduce the probability of a bond issue. The interpretation was that a rise in U.S. rates deterred emerging country issuers from coming to the market; with fewer issuers (who were likely to be of higher quality), prices rose and spreads fell. These results, however, may be sensitive to the nature of the underlying data used in the analyses. These studies focused on sovereign spreads for new bond issues (so-called launch spreads) rather than on spreads for bonds actively traded in secondary markets. Also, some of the analyses cover a subperiod (1991–93) when the market for sovereign bonds was developing, and another one (1994–95) when shocks seriously restricted access to the market for lower quality issuers.

This paper adds to the existing literature in three dimensions. First, rather than examining spreads on new issues, we examine secondary market sovereign spreads, which is the concept that is most common in public discussion and which, as Eichengreen and Mody (1998a) note, can behave differently than launch spreads, as they tend to be actively traded based on current and expected developments. Second, we isolate the impact of U.S. monetary policy by explicitly incorporating the U.S. federal funds target rate as an explanatory variable instead of the yield on a U.S. treasury security. Most of the specifications adopted so far have been somewhat simplistic, proxying U.S. monetary policy by the yield on U.S. treasury securities. However, shocks to U.S. treasury yields are not necessarily the result of changes in U.S. monetary policy. As seen in Figure 1, while the yield on the three-month U.S. treasury bill has in general fluctuated in tandem with the U.S. federal funds target rate, there have been many instances when these two rates have departed from each other. A recent instance was the so-called flight to quality during the Asian crisis, when U.S. treasury bill yields fluctuated dramatically even in the absence of changes in U.S. monetary policy. The level of the U.S. federal funds target rate is thus a more direct measure than the yield on U.S. treasury securities. Third, as discussed below, we explicitly analyze the role of market volatility, including uncertainty about U.S. monetary policy actions. In doing so, we present an autoregressive conditional heteroskedasticity (ARCH)–based measure of volatility that escapes criticisms that apply to more commonly used measures.

Clearly, from a theoretical and empirical point of view, changes in U.S. interest rates, or likewise in global liquidity conditions, would be expected to influence positively country risk and sovereign spreads in developing countries.

\[3\text{Cline and Barnes (1997) pointed out in addition that falling U.S. interest rates are generally associated with an abundance of capital in international markets, which tends to drive down yields.}\]

\[4\text{Earlier analyses based on secondary market developments include Dooley, Fernandez-Arias, and Kletzer (1996) and Calvo, Leiderman, and Reinhart (1996). These papers found a significant negative impact of industrial-country interest rates on secondary market prices of emerging market debt.}\]
While Eichengreen and Mody (1998a and 1998b) found that a tightening of global liquidity conditions (as proxied by the 10-year U.S. treasury bond yield) tended to reduce sovereign spreads in emerging markets, as noted earlier, their result may be largely explained by the nature of the underlying data used in their analyses and by the sample period (1991–95), which covers the 1991–93 subperiod when the market for sovereign bonds was at a very early stage of development and when shocks seriously restricted access to the market for lower quality issuers.

We started by replicating the methodology of earlier studies, but using secondary market data. We estimated the following model individually for a group of emerging markets. The model was estimated for Argentina, Brazil, Bulgaria, Colombia, Indonesia, Korea, Mexico, Panama, the Philippines, Poland, and Thailand for the period 1994–99 (with a few exceptions due to data limitations). We adopted the following standard linear relationship:

\[
\log(\text{spread},) = \alpha \log(\text{ustnote},) + \eta Z, + \omega, \quad (4)
\]

which aims at explaining fluctuations in the logarithm of sovereign spreads as a function of the log of the level of the yield on the 10-year U.S. treasury bond (ustnote) and country-specific macroeconomic variables (Z), where \( \alpha \) and \( \eta \) are parameters to be estimated and \( \omega \) is the error term. As for country-specific fundamentals, we selected a set of macroeconomic variables that have traditionally been used in the literature exploring fluctuations in sovereign spreads. In particular, the variables chosen were the fiscal balance, the net foreign asset position of the
banking system, central government external debt, and total external debt (all expressed as a ratio to GDP), the debt-service ratio, and the ratio of gross international reserves to imports. However, more recent studies, such as Kaminsky, Lizondo, and Reinhart (1997) and Kaminsky and Reinhart (1998), have emphasized the need to identify key macroeconomic and financial variables that may provide some early warning signals of banking and currency crises, and the role of other fundamental factors in driving banking and currency crises. Kaminsky, Lizondo, and Reinhart (1997) propose a set of variables that track more effectively the emergence of a crisis, such as deviations of the real exchange rate from trend, equity prices, and the ratio of broad money to gross international reserves. These additional variables, which are not part of our estimation, are worth exploring in future research.

The results, presented in Table 1, confirm that there is a positive relationship between sovereign spreads in secondary markets and the yield on U.S. treasury securities. Evidently, the main difference from Eichengreen and Mody's study seems to be related to the use of secondary rather than primary sovereign bond spreads. In addition, the results show that the level of the 10-year U.S. treasury note yield has a significant positive effect on sovereign spreads, with a mean group elasticity estimated at 0.78 and a mean standard error of 0.36.

While the above analysis suggests that the U.S. 10-year treasury note yield—a proxy for global liquidity conditions—tends to influence positively secondary market sovereign bond spreads, more direct measures of U.S. monetary policy—such as the U.S. federal funds rate—and a model-driven proxy for market volatility may help to explain better fluctuations in sovereign bond spreads.

What Is Market Turbulence and How Do We Proxy It?

Several authors have emphasized that, in addition to country-specific fundamentals, changes in market sentiment have been important in driving fluctuations in emerging market sovereign spreads (see, for example, Cantor and Packer (1996), Eichengreen and Mody (1998a and 1998b), and Kamin and von Kleist (1999)). These changes in market sentiment have often been sudden and abrupt, and have led many authors to argue that these changes in sentiment have been manifested by some form of market turbulence or “contagion” of shocks from one country to another, which has driven down sovereign debt prices or widened spreads. Baig and Goldfajn (2001) have analyzed the contagion from Russia to Brazil during 1998, while Edwards and Susmel (2000) have explored how changes in financial volatility, particularly interest rate volatility, have affected countries that supposedly have experienced market turbulence or contagion.

While some authors have argued that these episodes of market turbulence have to some extent reflected evidence of “irrational investor behavior,” others have tried to explain these episodes primarily as “liquidity events.” In particular, Valdés

5These regressions are based on the 10-year U.S. treasury note yield, as used by Eichengreen and Mody. Similar results were obtained when using the yield on the three-month U.S. treasury bill.

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Table 1. Determinants of Sovereign Bond Spreads for Selected Emerging Markets

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<td>0.96 (0.00)</td>
<td>0.84 (0.00)</td>
<td>0.21 (2.08)</td>
<td>0.54 (0.00)</td>
<td>1.01 (0.00)</td>
<td>1.08 (0.00)</td>
<td>0.54 (0.02)</td>
<td>0.54 (0.00)</td>
<td>1.54 (0.00)</td>
<td>0.83 (0.00)</td>
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<td>-0.09 (0.00)</td>
<td>-0.47 (0.01)</td>
<td>-0.32 (0.00)</td>
<td>0.00 (4.93)</td>
<td>...</td>
<td>-0.01 (0.00)</td>
<td>-0.08 (0.00)</td>
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<td>-0.09 (0.00)</td>
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<td>...</td>
<td>...</td>
<td>...</td>
<td>...</td>
<td>-1.01 (0.00)</td>
<td></td>
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<td>-0.39 (0.708)</td>
<td>...</td>
<td>...</td>
<td>-2.25 (0.050)</td>
<td>...</td>
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<td>...</td>
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<tr>
<td>Debt-service ratio</td>
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<tr>
<td>Central government debt (in percent of GDP)</td>
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<td>0.10 (0.00)</td>
<td>0.06 (0.00)</td>
<td>0.21 (0.00)</td>
<td>0.11 (0.03)</td>
<td>0.09 (0.00)</td>
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<td>0.11 (0.06)</td>
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<td>Total external debt (in percent of GDP)</td>
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<tr>
<td>Adjusted R-squared</td>
<td>0.53</td>
<td>0.31</td>
<td>0.24</td>
<td>0.60</td>
<td>0.76</td>
<td>0.50</td>
<td>0.57</td>
<td>0.29</td>
<td>0.77</td>
<td>0.75</td>
<td>0.71</td>
</tr>
</tbody>
</table>

Source: IMF staff estimates.

Note: Probability values for the null hypothesis of a coefficient equal to zero, are reported in parentheses. Standard errors have been adjusted following the Newey-West procedure.

1Refers to net debt.

2A dummy was included to allow for the effects associated with the introduction of a currency board in Bulgaria.
(1997) and Kaminsky and Reinhart (2000) have emphasized the financial aspects of contagion. Kaminsky and Reinhart noted the role of international bank lending and cross-market hedging as sources of contagion based on fundamentals. Valdés argued that contagion results primarily from the interaction of investors who face liquidity constraints and who have invested in emerging market assets that are potentially highly illiquid. When facing liquidity needs in one particular class of asset or country, such investors would tend to withdraw liquidity from some other class or country.

In the same vein, others have emphasized the importance of liquidity effects on capital flows and asset prices in emerging markets, which on certain occasions may have been associated with sudden and unexpected changes in U.S. monetary policy. For example, during the first half of 1994, the U.S. Federal Reserve raised the federal funds rate by 125 basis points, which precipitated a sharp unwinding of highly leveraged positions by hedge funds, proprietary traders, and institutional investors, which had been financing purchases of long-term treasury securities with short-term borrowing. This unwinding of positions contributed to and exacerbated a steep correction in emerging sovereign bond markets. In sum, what these authors suggest is that a need for liquidity, precipitated by a rise in U.S. interest rates or other exogenous shock, becomes one of the main transmission vehicles of financial turmoil across assets and countries.

In terms of how to model turbulence or volatility, several approaches have been tested in the literature. Most approaches have used statistical measures based on standard errors for a certain variable that was considered as relevant in capturing the market turbulence or contagion. For example, the work of Hardouvelis (1989) in exploring the link between the level of margin requirements and stock market volatility in the United States was based on a moving average representation for volatility of real stock returns. More recent studies (Hardouvelis, Pericli, and Theodossiou, 1997) have proxied market volatility by computing the standard deviation of daily returns during a month. This volatility measure, which is based on daily data, is constructed in a way that tends to avoid data overlapping, and its associated problems, by being sampled every month. Other more advanced techniques have aimed at estimating conditional volatility and have been based on Schwert’s (1989) procedure and on the ARCH model developed by Engle (1982). With these methods in mind, we proxied market volatility by computing different statistical and econometric measures on the spread between the yield on the three-month U.S. treasury bill and the U.S. federal funds target rate. In principle, the yield on the three-month U.S. treasury bill can be considered a key short-term risk-free rate that usually serves as a benchmark for pricing other high-yield assets in world capital markets, and that would most likely reflect changes in global liquidity and economic conditions. More important, changes in the spread between the three-month treasury bill yield and the

---

6See IMF (1995a, 1995b, and 1996) for a more detailed analysis and account of events.
U.S. federal funds target rate may capture heightened uncertainty about the expected stance of U.S. monetary policy, as in the first half of 1994. All the different proxies for market volatility tend to show increased market turbulence during 1994 and in the second half of 1998 (Figure 2). A scatter plot shows that a proxy based on a six-month moving average of standard deviations for the spread between the three-month yield on the U.S. treasury bill and the federal funds target rate was highly statistically significant in explaining fluctuations in sovereign spreads across countries (Figure 3). The validity of this proxy for volatility, however, has been questioned in the empirical literature by Hsieh and Miller (1990), who argue that it induces a spurious correlation between variables due to its high serial correlation. The construction of this proxy using moving averages leads to strong autocorrelation, which leads to highly problematic statistical inference. Therefore, regressing a highly autocorrelated series, such as the proxy for market turbulence, on other variables can produce a significant coefficient, even when no true relationship exists. The clustering of observations in Figure 3 would suggest that, were these observations independent, a strong direct relationship would be found between sovereign spreads and the proxy for volatility. The $R^2$ is quite high, at close to 30 percent. In fact, these observations are far from independent, and the high positive correlation between sovereign spreads and the proxy for market volatility is primarily the result of the way the proxy was constructed. Nevertheless, in our empirical estimates, we used these proxies for market volatility in estimating the model, and the results are clearly sensitive to the chosen proxy. Given the constraints and limitations of the first two proxies for market turbulence noted above, however, we decided to use the fitted values for the conditional standard error from an ARCH model for the spread between the three-month yield on the U.S. treasury bill and the federal funds target rate. As is well established in the literature, ARCH models are useful in analyzing financial data because they capture the persistence in volatility that is observed in many financial time series. In particular, large shocks tend to be followed by large shocks of unpredictable sign, suggesting that there is persistence in market volatility and that it tends to vary over time. As seen in Figure 4, the positive relationship between spreads and market volatility looks significantly different from the one presented in Figure 3 once one allows for a proxy that minimizes data overlapping and serial correlation. In fact, there is less of a positive correlation between the variables, as the $R^2$ declines to only 8 percent.

7It is also evident, however, that changes in this spread may not necessarily fully reflect expected changes in the stance of U.S. monetary policy, as was demonstrated during the Asian crisis and, to some extent, during the events associated with the default by Russia and the near demise of LTCM during the second half of 1998. Cline and Barnes (1997) and Kamin and von Kleist (1999) use short-term interest rates as a proxy for global liquidity conditions. Eichengreen and Mody (1998a and 1998b) use the yield on the 10-year U.S. treasury bond as a proxy for global economic conditions.

8The alternative proxy suggested by Hardouvelis, Pericli, and Theodossiou (1997), the standard deviation of the daily spread within a month, is not presented in the paper because it was not statistically significant in most equations, except in those for Argentina, Bulgaria, and Indonesia.

9Notwithstanding these shortcomings, the autocorrelation coefficient is not highly persistent, as it declines to almost zero at the fourth lag.

10When using the six-month moving average proxy for market volatility, the econometric estimates show that this variable is highly significant across countries.
U.S. MONETARY POLICY AND EMERGING MARKET SPREADS

Figure 2. Alternative Proxies for Market Volatility

- Within-Month-based (right scale)
- Six-month moving average (right scale)
- ARCH-based (left scale)

Figure 3. Sovereign Spreads and Moving-Average-Based Market Volatility

R² = 0.28
Econometric Evidence

The econometric model for sovereign bond spreads presented in the previous section was modified by explicitly including the U.S. federal funds target rate and a proxy for market volatility. The model, which was estimated individually for the same group of countries, is as follows:11

\[ \log(\text{spread}_t) = \rho \log(\text{ff}_t) + \lambda \text{mktvol}_t + \theta Z_t + e_t. \]  

The model aims at explaining fluctuations in the logarithm of sovereign spreads as a function of the level of the U.S. federal funds target rate (\(\text{ff}_t\)), the proxy for market volatility (\(\text{mktvol}\)) derived from an ARCH model, and country-specific macroeconomic variables (\(Z\)), where \(\rho\), \(\lambda\), and \(\theta\) are parameters to be estimated, and \(e\) is the error term.12 As explained before, the proxy for market volatility is

---

11 We did not believe panel data estimation would have been more efficient than the chosen procedure. As the results show, homogeneity in the estimated parameters is highly rejected, as parameters differ significantly across countries and even within regions. With a relatively small number of countries and a large number of observations, it is more efficient to estimate the model for each country separately rather than impose some form of homogeneity through panel data estimation. In addition, panel data estimation would have severely restricted the sample period, given that data for most Asian countries were available starting only in 1997.

12 An alternative proxy for U.S. monetary policy is the federal funds futures rate. In using the target (spot) rate, we thought that market expectations of the federal funds rate would be reflected in the spot yield on the three-month treasury bill, and as a result our proxy for market volatility would indirectly capture expectations about U.S. monetary policy.
intended to capture changes in investor sentiment which may be related to expected changes in U.S. monetary policy. It may also pick up the effects of other market-related events, such as the flight to quality effects during the Asian crisis.

In line with the previous model, the results show that the level of the U.S. federal funds target rate has significant positive effects on emerging market spreads, with the mean group elasticity estimated at 0.82 (Table 2). The estimated elasticities vary considerably across countries (the standard error for the mean group estimate is 0.35): the estimates for Argentina, Colombia, Panama, and the Philippines are smaller than the average; the estimates for Brazil, Mexico, and Bulgaria are close to 1; and those for Korea and Poland appear to be very high given their past macroeconomic performance and low indebtedness. Nevertheless, the results suggest that the estimated impact of changes in the U.S. federal funds rate on sovereign spreads is slightly higher (but economically not significant) than the one estimated for the yield on the 10-year U.S. treasury bond.

The model also supports the view that increased market volatility, which may be related to heightened uncertainty about the expected path of U.S. monetary policy, has significant positive effects on spreads across countries and regions. However, a significant proportion of fluctuations in emerging market spreads is driven by country-specific fundamentals. In particular, the results suggest that improved macroeconomic fundamentals, such as higher net foreign assets (in terms of GDP or imports), lower fiscal deficits, and lower ratios of debt service to exports and debt to GDP, help to lower sovereign spreads. For example, a higher net foreign asset position contributed to lower spreads in many Latin American and Asian countries—particularly those that had in place fixed exchange rate regimes and where lender-of-last-resort considerations seemed particularly important—such as Argentina, Panama, Thailand, and Korea. Foreign indebtedness appears to contribute positively to sovereign spreads, especially in Latin America (particularly Argentina, Mexico, Brazil, and Panama), the Philippines, and to some extent Poland, all countries that underwent comprehensive debt reschedulings in the past.

The model presented in Table 2 explains fluctuations in emerging market sovereign spreads relatively well for most countries (see Figure A1 in the appendix). In particular, the model explains roughly between half and three-quarters of the fluctuations in spreads for most countries, and for most countries (9 out of 11) the adjusted R² increases significantly. In addition, using the Phillips-Perron (1988) test, we do not reject the hypothesis that sovereign spreads are cointegrated with the chosen country-specific fundamentals, the U.S. federal funds rate, and the

---

13Needless to say, the rise in the level of emerging market interest rates will not necessarily be as large as the sum of the rise in spreads and the rise in the U.S. federal funds rate. In the United States, the yield curve tends to flatten as monetary policy is tightened, so that a rise in short-term interest rates is usually not fully passed through to longer-term rates.

14The results for Korea, Thailand, and Indonesia, especially the size of the U.S. interest rate elasticity, should be interpreted with some caution due to the relatively small sample size and the fact that the estimation mainly covers the period of an IMF arrangement. In the case of Poland, the model did not include any measure of indebtedness due to the lack of a time series from 1994, and as a result, may be biasing upwards the coefficient of the U.S. federal funds rate.
Table 2. Determinants of Sovereign Bond Spreads for Selected Emerging Markets

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<td>U.S. federal funds rate</td>
<td>0.54 (.000)</td>
<td>0.95 (.000)</td>
<td>0.93 (.000)</td>
<td>0.26 (.052)</td>
<td>0.54 (.000)</td>
<td>1.26 (.000)</td>
<td>1.09 (.000)</td>
<td>0.57 (.000)</td>
<td>0.63 (.001)</td>
<td>1.45 (.000)</td>
<td>0.78 (.000)</td>
</tr>
<tr>
<td>Markel volatility(^1)</td>
<td>0.08 (.003)</td>
<td>0.05 (.000)</td>
<td>0.07 (.013)</td>
<td>-0.01 (.411)</td>
<td>0.05 (.001)</td>
<td>0.03 (.150)</td>
<td>0.04 (.093)</td>
<td>0.02 (.022)</td>
<td>-0.01 (.576)</td>
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<td>Net foreign assets (in percent of GDP)</td>
<td>-0.05 (.000)</td>
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<td>-0.38 (.000)</td>
<td>-0.29 (.036)</td>
<td>-0.01 (.056)</td>
<td>... ...</td>
<td>-0.01 (.000)</td>
<td>-0.07 (.000)</td>
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<td>-0.01 (.010)</td>
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<td>-0.08 (.000)</td>
<td>... ...</td>
<td>... ...</td>
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<td>-0.91 (.000)</td>
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<td>Gross reserves to imports</td>
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<td>... -2.68 (.001)</td>
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<td>Debt-service ratio</td>
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<td>0.08 (.160)</td>
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Table 2. (concluded)

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<tr>
<td>(in percent of GDP)</td>
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<td>Dummy¹</td>
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<td>-0.40</td>
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<td>(in percent of GDP)</td>
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<tr>
<td>Adjusted R-squared</td>
<td>0.51</td>
<td>0.55</td>
<td>0.45</td>
<td>0.62</td>
<td>0.81</td>
<td>0.54</td>
<td>0.60</td>
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</table>

Source: IMF staff estimates.

Notes: Probability values, for the null hypothesis of a coefficient equal to zero, are reported in parentheses. One and two asterisks imply rejection of the null hypothesis of no cointegration at the 90 and 95 percent level of significance. Standard errors have been adjusted following the Newey-West procedure.

¹Based on the fitted conditional standard error from an ARCH model for the spread between the three-month treasury bill and the federal funds rate.

²Refers to net debt.

³A dummy was included to allow for the effects associated with the introduction of a currency board in Bulgaria.

proxy for market volatility in 8 out of 11 countries, with the models for Brazil, Poland, and the Philippines rejecting the hypothesis of cointegration. This may be partly related to the finding that the model is subject to a structural break in late 1995 in several countries (see Figure A2 in the appendix). Specifically, in the cases of Argentina, Brazil, Bulgaria, Mexico, the Philippines, and Poland, the model fails to fully account for the sharp narrowing of spreads that took place during the period leading up to the Asian crisis. The narrowing of sovereign spreads between the first half of 1996 and mid-1997 was particularly pronounced in these countries, and may have been associated more with changes in market access and with global portfolio shifts by institutional investors than with country-specific fundamentals. These results seem to suggest that some form of "contagion" may have also contributed to narrowing rather than widening sovereign spreads for a group of developing countries during this period.

Global Liquidity Conditions and Other Factors at Work

Following the Mexican financial crisis of 1994–95, there was a large compression of emerging market sovereign spreads, which declined from a peak close to 1,600 basis points in March 1995 to about 325 basis points in July 1997 (see Figure 1). In fact, as noted by some analysts, the international bond market experienced, between end-1994 and early 1996, one of the greatest rallies in its recent history. Such a compression in sovereign spreads for U.S. dollar-denominated bonds was driven by supply as well as demand factors. On the supply side, Andrews and Ishii (1995) noted that developing countries shifted the currency denomination of bond issues away from the U.S. dollar to issues denominated in deutsche marks and Japanese yen. In fact, Argentina, Brazil, and Mexico became very active in issuing yen-denominated bonds in the Euro-yen and Japanese markets (Figure 5). Access by developing countries to the alternative currency issues was eased by the deregulation of the yen-denominated market, which eliminated restrictions on the sale of sovereign yen-denominated Eurobond issues to Japanese investors in 1994, and reduced the minimum credit rating requirement in 1996, from investment to noninvestment grade for any sovereign issuer of Samurai bonds. On the demand side, interest rates in industrial countries declined markedly and were at extremely low levels in Japan, Germany, and France for a considerable period of time (Figure 6), while several of the Latin American countries, particularly Mexico, faced a rapid recovery in macroeconomic fundamentals. All these factors may have contributed to restoring investors' confidence rapidly, boosting global liquidity, and renewing the demand for new bond issues by developing countries.

It is difficult to assess whether the failure of our model to fully account for the sharp compression in spreads in Argentina, Brazil, Mexico, Bulgaria, Poland, and the Philippines, particularly between mid-1996 and mid-1997, reflects the omission or inadequate account of country-specific fundamentals rather than the inability to capture global changes (including global liquidity conditions, portfolio shifts, or momentum strategies by institutional investors). Nevertheless, we

suspect that global liquidity factors may have been at work given that the failure to predict such narrowing of spreads is primarily confined to a group of developing countries that have usually been treated by institutional investors as one group in
an asset class.\textsuperscript{16} To capture some of these effects, particularly the structural changes associated with the liberalization of the yen-denominated bond market, we extended the model by including the Hodrick-Prescott cyclical component of the number of yen-denominated sovereign bond issues by emerging-market countries during this period. A significant (but very small) negative effect was found for some of those countries, particularly Argentina and Mexico, while the rest of the results remained largely unchanged.\textsuperscript{17}

II. Conclusions

This paper presented empirical evidence on how U.S. monetary policy has influenced country risk in several developing countries in Latin America, Asia, and Eastern Europe. In contrast to previous results in the literature, but consistent with what we might anticipate from theory, our results suggest that the level of U.S. interest rates has direct positive effects on sovereign bond spreads. In addition, the econometric evidence supports the view that, while country-specific fundamentals are extremely important in determining country risk, so is the stance and predictability of U.S. monetary policy.

An approach to U.S. monetary policy that provides financial markets with clear indications of policymakers' views about the balance of inflationary risks and intentions is likely to reduce the negative impact of a rise in U.S. interest rates on country risk in developing countries. More important, policymakers in developing countries still enjoy a significant degree of freedom to influence country risk and economic growth. Country-specific macroeconomic fundamentals, such as a sound and sustainable fiscal policy and low indebtedness, are extremely important in reducing country risk and domestic interest rates, factors that are highly conducive to fostering sustainable economic growth.

The search for the best proxies for U.S. monetary policy, market volatility, and country-specific fundamentals is a complicated task and we do not claim to have found the true underlying model. Evidently, several different options are available to model any of the fundamental factors determining country risk. In particular, future research could explore the role of the U.S. federal funds futures, rather than the target level, as a proxy for U.S. monetary policy, while some of the early warning indicators of currency crises can be included in the set of country-specific fundamentals.

\textsuperscript{16}See Aitken (1998), and Borensztein and Gelos (2000) for empirical evidence on the role played by shifts in institutional investors' sentiment in determining asset prices in developing countries.

\textsuperscript{17}The results are available upon request.
APPENDIX

Data Description

Data on sovereign bond spreads for each country were obtained from Merrill Lynch and are based on its iGOV Index. The U.S. target federal funds rate and the three-month U.S. treasury bill rate were obtained from the U.S. Federal Reserve system.

Country-specific data were based on information provided by national authorities. Several data series were available on a monthly basis, but some (including GDP) were available only on a quarterly basis, and a few only on an annual basis. Quarterly and annual data were converted to a monthly basis using a cubic spline interpolation.

Data definitions are as follows:

- **Net foreign assets (NFA)**: NFA of the banking system, in percent of GDP.
- **Fiscal balance**: Budget balance of the central or federal government, defined in percent of GDP.
- **Gross reserves to imports**: Gross international reserves as a percent of imports of goods and nonfactor services.
- **Debt-service ratio**: External debt service as a percent of exports of goods and nonfactor services.
- **Central government debt**: External debt of the central or federal government, in percent of GDP.
- **Total external debt**: External debt of the private and public sectors, in percent of GDP.

Details on the estimation period and data availability for individual countries are as follows:

- **Argentina**

- **Brazil**

- **Bulgaria**

- **Colombia**

- **Indonesia**
Korea

Mexico

Panama

Philippines

Poland

Thailand
Figure A1. Sovereign Spreads in Selected Emerging Markets, Actual versus Fitted Values (in logarithm)

Argentina

Brazil

Bulgaria

Colombia

Indonesia

Korea
Figure A1. (concluded)

Sources: Merrill Lynch; and authors' estimates.
U.S. MONETARY POLICY AND EMERGING MARKET SPREADS

Figure A2. Stability Tests

Argentina

Brazil

Bulgaria

Colombia

Indonesia

Korea

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Figure A2. (concluded)

Mexico

Panama

Philippines

Poland

Thailand

Based on the cumulative sum of squared residuals statistic. Confidence bands for a 95 percent level of significance.
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Vivek Arora and Martin Cerisola


Modeling the IMF's Statistical Discrepancy in the Global Current Account

JAIME MARQUEZ and LISA WORKMAN*

This paper offers a framework for judging when the discrepancy embodied in current account forecasts is large. The first step in implementing this framework involves developing an econometric model explaining the components of the aggregate discrepancy, estimating the associated parameters, and generating the aggregate discrepancy's conditional expectation. The second step is to compare this model-based forecast with the discrepancy embodied in countries' current-account forecasts. If the gap in discrepancies is below a critical value, then the discrepancy embodied in the countries' current account forecasts is not large. Otherwise, the discrepancy is large and calls for a careful reexamination of the associated current account forecasts.

For projections of global external imbalances to be useful, they must be internally consistent: external surpluses and deficits across countries must add up

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to zero. That this adding-up condition does not hold in the data need not, by itself, undermine the usefulness of these projections: reporting mistakes, sampling errors, and recording asymmetries across countries are a fact of life and, when combined, give rise to a statistical discrepancy. But when this discrepancy reaches the level of the current account surplus of Japan in 1990 ($100 billion), and is projected to reach almost $300 billion by 2001 (Figure 1), one cannot avoid questioning the usefulness of such projections.

Sustained discrepancies of this magnitude are worrisome because they undermine the credibility of global current account forecasts. Specifically, if the discrepancy stems from recording practices of a few countries, then their forecasts for growth could be systematically mismeasured with obvious implications for the global consistency of individual country forecasts of the current account. Such consistency is central to studying, for example, which countries will absorb the reduction of the U.S. external deficit.

One tempting response to this statistical discrepancy is to develop a rule to allocate it across countries. Years of work by the IMF suggest, however, that reliance on such rules might further distort the countries' current account forecasts. Indeed, the evolution of this discrepancy has no obvious statistical pattern and, thus, an allocation rule suitable for one year need not work in other years. Moreover, movements in the statistical discrepancy reflect large and often mutually offsetting movements in its components. This property creates the impression of improved accuracy when just the opposite is true (Figure 2). Thus using a rule to allocate the discrepancy over 1993–97 would reallocate small amounts—precisely the opposite of what is needed.

Facing sustained and large discrepancies in the global current account with no reliable allocation rule leaves practitioners with two courses of action: either ignore the discrepancy and the internal consistency of global forecasts or reexamine the associated current account forecasts if the implied discrepancy is, in some sense, large. In other words, if we denote $C_i$ as the current account forecast for the $i$th country, then $D^c = \Sigma \delta C_i$ is the current account discrepancy associated with those forecasts and the question is whether $D^c$ is large enough to merit a revision of the underlying $C_i$s.

Determining whether $D^c$ is large involves specifying a reference value and this paper offers a practical approach to determining it. Specifically, as reference value we choose the expected value of the discrepancy. This choice allows us to define a discrepancy as large if it is significantly different from its conditional expectation. Implementing this choice involves developing an econometric model to generate the distribution of the discrepancy and Sections II–IV of this paper document the associated modeling aspects: level of disaggregation, functional form, explanatory variables, and estimation method.

1The chief study on the global current account discrepancy is the 1987 Report on the World Current Account Discrepancy (IMF, 1987), which focuses on the 1983 discrepancy; its findings have been confirmed in IMF (1996, 1999a, 1999b). The appendix uses a hypothetical example to highlight the potential pitfalls in using fixed rules to allocate the global current account discrepancy across countries.
Figure 1. World Current-Account Discrepancy


Figure 2. World Current-Account Discrepancy by Categories (percent of World Imports)
To evaluate the usefulness of the model, Section III asks whether it can detect large discrepancies when large discrepancies are known to occur: not being able to detect a known significant change would question the model’s usefulness. As a test case we use the switch in Europe’s trade methodology in 1993, which is acknowledged to be responsible for large discrepancies (IMF, 1997, p. 9). We then generate ex ante forecasts of the discrepancy through 2001 and compare them against those reported in the IMF’s World Economic Outlook for May 2000. The model identifies the discrepancy in that report as large and calls for a rethinking of the country-based forecasts.

I. Model Design

Aggregation

The model explains the global current account discrepancy, $D$, as the sum of the discrepancies in four global accounts:

$$D = D_g(q) + D_i(q) + D_s(q) + D_u(q),$$

where $q$ is a vector of explanatory variables, $D_g$ is the global discrepancy in the trade account, $D_i$ is the global discrepancy in the investment-income account, $D_s$ is the global discrepancy in the service account, and $D_u$ is the global discrepancy in the unrequited-transfers account.

The alternative to explaining the accounts’ discrepancies is to develop a single-equation model for the overall discrepancy, as in Sheets (1998). A single equation is appealing because of its simplicity but it suffers from aggregation pitfalls. Specifically, the global discrepancy might be zero not because of accurate recordings but because discrepancies in various accounts are mutually offsetting (see Figure 2). Coefficient estimates of a single equation would then reflect the happenstance of inaccurate recordings whereas coefficient estimates for separate equations would avoid them.

Analytical Framework

We differentiate between actual and recorded transactions. Let actual global credits in a given account be $X_a$ and actual global debits be $M_a$, where $a = g, i, s, u$. Whereas $X_a = M_a$, the corresponding recorded measures need not be identical:

$$X_a' = X_a \cdot (1 + e_{xa}),$$

$$M_a' = M_a \cdot (1 + e_{ma}), a = g, i, s, u,$$
MODELING THE IMF’S STATISTICAL DISCREPANCY

where the ' indicates a recorded magnitude, \( e_{xa} \) is the error in credits, and \( e_{md} \) is the error in debits. The global discrepancy in that account is:

\[
\mathcal{F}'_a = X'_a - M'_a = X'_a - M'_a + X'_a \cdot e_{xa} - M'_a \cdot e_{md} = X'_a \cdot (e_{xa} - e_{md}) \quad a = g, i, s, u.
\]

To translate this accounting identity into a statistical model, we postulate that \( e_{xa} = e_{xa}(q, u_a) \) and \( e_{md} = e_{md}(q, v_a) \) where \( u_a \) and \( v_a \) are random variables. Thus:

\[
\mathcal{F}'_a = X'_a \cdot [e_{xa}(q, u_a) - e_{md}(q, v_a)], \quad a = g, i, s, u.
\]

We now assume that:

\[
X_a = \theta_a \cdot z_a(q, w_a) \cdot M_w', \quad a = g, i, s, u,
\]

where \( M_w' \) is recorded world imports and \( w_a \) is a random variable. With this assumption, we model the importance of a statistical discrepancy as:

\[
\mathcal{F}'_a / M_w' \equiv D_a = \left[ \theta_a \cdot z_a(q, w_a) \right] \cdot \left[ e_{xa}(q, u_a) - e_{md}(q, v_a) \right], \quad a = g, i, s, u,
\]

which is nonlinear in the variables included in \( q \).

To illustrate the key features of our approach, assume the simplest formulation:

\[
e_{xa} - e_{md} = (e_{xa} - e_{md}) \cdot (\gamma_{a0} + \gamma_{ai} \cdot q_i + u_{ar}) \cdot u_{at} - N(0, \sigma_2^2)
\]

\[
z_a(q, w_a) = \pi_{a0} \cdot q_i \quad \text{for } a = g, i, s, u.
\]

Then:

\[
D_a = (e_{xa} - e_{md}) \cdot \left[ (\theta_a \cdot \pi_{a0}) \cdot (\gamma_{a0} \cdot q_i^2 + \gamma_{ai} \cdot q_i^2 + u_{at} \cdot q_{it}) \right] = (e_{xa} - e_{md}) \cdot \left[ \ell_{a0} \cdot q_i^2 + \ell_{ai} \cdot q_i^2 + \ell_{at} \cdot u_{at} \cdot q_{it} \right], \quad a = g, i, s, u.
\]
The term \((e_{xa} - e_{ma})\) embodies the factors that give rise to a statistical discrepancy in the first place: reporting mistakes and sampling errors. In the absence of these factors, the actual and the recorded transactions would be equal to each other and there would be no statistical discrepancy. The term in square brackets embodies the factors that account for movements in the (scaled) discrepancy. For example, if \(q_{it}\) is a variable capturing behavioral incentives to misreport by one of the transactors, then changes in those incentives will induce nonlinear changes in the account's statistical discrepancy.\(^2\)

**Statistical Framework**

Our modeling recognizes the roles of simultaneity and dynamics. Simultaneity considerations arise because international transactions are recorded using the principle of double entry. This principle requires recording two accounts simultaneously and, thus, discrepancies in one account could reflect mismeasurements from another.\(^3\)

Dynamic considerations might arise because faulty recording practices are institutionalized and fixing them takes time. One formulation capturing these two features is:

\[
A_0 \cdot \Lambda_t = A_1 \cdot \Lambda_{t-1} + B \cdot Q_t + U_t \tag{2}
\]

\[
D_t = 1' \cdot \Lambda_t \tag{3}
\]

where \(A_0\) is a \(4 \times 4\) matrix of coefficients recognizing the role of simultaneity; \(A_1\) is the \(4 \times 1\) vector of discrepancies to be modeled \((\Lambda' = (D_g, D_s, D_c, D_d))\); \(A_1\) is a \(4 \times 4\) matrix of coefficients capturing the importance of dynamic considerations; \(B\) is a \(4 \times n\) matrix of coefficients; \(Q_t\) is the vector of explanatory variables consisting of the entries in \(q_t\), as well as nonlinear terms (more below); \(U_t\) is the vector of disturbances distributed as \(N(0, \Omega_u, 1)\); and \(1'\) is a vector of ones.

The reduced form implied by equation (2) is:

\[
\Lambda_t = \Pi_d \cdot \Lambda_{t-1} + \Pi_q \cdot Q_t + V_t \tag{4}
\]

\[
D_t = 1' \cdot \Lambda_t \tag{5}
\]

\(^2\)Our working paper documents how we apply this framework to deriving estimating equations for each account's discrepancy; see http://www.federalreserve.gov/pubs/ifdp/2000/678/ifdp678.pdf.

\(^3\)For example, a donor country may record the value of a transfer as both a credit (such as merchandise exports for aid) and as a debit (unrequited transfer). The recipient country might debit the trade account (e.g., merchandise imports) and credit the capital account (e.g., capital inflow) instead of crediting unrequited transfers.
where $\Pi_d = A_0^{-1} \cdot A_t$, $\Pi_q = A_0^{-1} \cdot B$, and $V_t = A_0^{-1} \cdot U_t \sim N(0, \Omega_v)$. Section III uses the Full Information Maximum Likelihood method (FIML) for parameter estimation and implements dynamic simulations to estimate the expected global discrepancy at time $t$ as:

$$\hat{E}(D_t) = \hat{\Pi}_d \cdot \hat{\lambda}_{t-1} + \hat{\Pi}_q \cdot Q_t + V_t.$$  \hspace{1cm} (6)

where a circumflex denotes an estimated magnitude. Note that $\hat{E}(D_t)$ is conditioning on the model’s own generated values for the lagged endogenous variables and not on historical values. The estimate of the variance of the discrepancy at time $t$ is:

$$\text{vär}(D_t) = \text{var}(\hat{\Pi}_d \cdot \hat{\lambda}_{t-1} + \hat{\Pi}_q \cdot Q_t + V_t).$$  \hspace{1cm} (7)

Note that $\text{vär}(D_t)$ varies with changes in the explanatory variables. Furthermore, reliance on FIML allows for the correlations across the residuals of the model to affect $\text{vär}(D_t)$.

One can use these equations to test whether the discrepancy embodied in countries’ current account forecasts, $D^c_t$, is large. Specifically, the null and alternative hypotheses are:

$$H_0: E(D_t) = D^c_t$$

and

$$H_1: E(D_t) \neq D^c_t,$$

and the test statistic is:

$$\tau = \frac{D^c_t - \hat{E}(D_t)}{\sqrt{\text{vär}(D_t)}}.$$  

If $V_t \sim N(0, \Omega_v)$, then finding that $|\tau| > 2$ means that $D^c_t$ is statistically different from its expected value at the 5 percent significance level. We interpret such a finding as suggesting that $D^c_t$ is large.
II. Model Assembly

Selection of Explanatory Variables

Looking to economic theory for what variables to include in $q$ is not fruitful here because there is no economic theory of current account discrepancies as such. Thus to select the explanatory variables, we identify the forces responsible for each account's discrepancy and then translate those forces into a list of macroeconomic variables. This approach yields too many variables to consider and thus, to discriminate among them, we invoke additional criteria. First, the data must be available on a timely basis. Arguing that a variable should be included in a model because its coefficient is highly significant loses its force if the associated data are available with a long delay. Second, given the annual frequency of observations, the number of explanatory variables should be as small as possible. Third, generating a forecast of $q$ should not be more difficult than generating forecasts of the discrepancies directly. Finally, the estimation results cannot violate the maintained assumptions for the residuals. These assumptions are central to the definition of a large discrepancy.

Discrepancies in trade

The factors responsible for a discrepancy in the global trade account are transportation delays, asymmetric valuations, and quality differentials in recording practices. Transportation delays in shipping merchandise imply that recorded increases in export credits are not accompanied by simultaneous recorded increases in import debits. To translate the role of these shipment delays into an explanatory variable, we assume that fluctuations in world trade are driven by fluctuations in economic activity. Thus faster world growth raises recorded exports ahead of recorded imports and raises net credits in the global trade balance.

Asymmetric valuations arise whenever different prices are used to value the same transaction. For example, recipients of oil subsidies from the Organization of the Petroleum Exporting Countries (OPEC) could record oil imports (debits) at the subsidized price whereas OPEC could record the corresponding oil exports (credits) at the market price. In that case, debits would increase less than credits thus inducing a discrepancy in the global trade account. Another example involves the use of different exchange rates to value the same transaction by at least one of the reporting countries. To model these valuation asymmetries we use the price of oil as a proxy for commodity prices, and the U.S. federal funds rate as a proxy for

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4Idiosyncratic recording practices can also induce discrepancies. For example, Bermuda, the Cayman Islands, and Hong Kong SAR combine direct-investment income with other income measures. Similarly, Middle Eastern oil exporters do not report cross-border investment income of private nonbanks (IMF, 1987, pp. 57–58). By their nature, idiosyncratic factors lack an obvious representation in terms of a macroeconomic variable.
exchange rates. The alternative of including in $q$ the numerous exchange rates would exhaust the degrees of freedom.

Differentials in recording quality across countries might help explain movements in the discrepancies. Specifically, if countries with high-quality data increase their share of world trade, then one would expect a reduction in the existing trade discrepancy. To model this possibility, we assume that the United States is the high-quality data country and postulate that if the U.S. share of world imports increases, then there would be a reduction of net credits in the trade discrepancy, all else given. A more concrete example of the role of quality involves the change in methodologies for collecting intra-European Union (EU) trade data by the EU in 1993: a switch from custom records to value-added tax records (IMF, 1997, p. 9). The IMF estimates that this switch induced an excess of credits over debits of $40$ billion a year (IMF. 1999a, p. 4). We model this effect with a dummy variable.\(^5\)

**Discrepancies in investment income**

Discrepancies in this account reflect misrecordings in portfolio-investment income and direct-investment income. These discrepancies stem from incentives to understate capital outflows, the growth of offshore financial centers, and recording idiosyncrasies.\(^6\) The incentive to underreport capital outflows arises from tax avoidance on the corresponding income. This tax-evasion incentive leads to an underreporting of investment income because accounting practices use cumulated capital outflows to estimate the corresponding stock of claims on foreigners, which is then used to compute investment income. Thus understating capital outflows translates into understating the associated income. To model this factor, we assume that an increase in the U.S. federal funds rate accentuates the incentive to understate capital outflows, which results in an underestimate of the stock of claims on foreigners and the resulting investment income.

The growth of offshore financial centers is contributing to the discrepancy in investment income by undermining the ability of statistical agencies to track financial transactions. Specifically, such centers are largely unrelated to domestic activities of the host country and typically do not have to report to the host's statistical agencies.\(^7\) Also, the associated transactions involve securitization with numerous

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\(^5\)To investigate this possibility further, we use the IMF's *Direction of Trade* that reports bilateral trade data. We then compute intra-EU credits and debits and find a sharp increase in net credits starting in 1993. As a fraction of EU imports from their members, the discrepancy increases from about 1 percent prior to 1993 to 7.3 percent in 1993. This gap reaches a maximum of 9.3 percent in 1996.

\(^6\)Portfolio income includes interest payments/receipts among banks, interest and dividends on securities, commercial paper, mortgages, and supplier credits. Direct-investment income includes earnings of foreign subsidiaries, earnings of unincorporated business in foreign countries, and interest of foreign-incorporated affiliates and branches. The bulk of these discrepancies stems from discrepancies in portfolio-investment income.

\(^7\)The new offshore financial centers are located in Hong Kong SAR, Singapore, Bahrain, the Bahamas, Cayman Islands, and Panama; Bahrain's operations are recorded in that country's statistics. These centers offer unregulated operations and tax advantages not offered by traditional centers (e.g., New York, London, and Zurich).
participants, not all of whom report to any national compilers. Expecting financial innovation to grow over time, we use a trend to capture the effect of this process on the discrepancy.

Recording idiosyncrasies come in two flavors: misclassifications and asymmetries. Misclassifications arise from the ambiguity of the term *lasting interest*, which is the criterion used for classifying foreign direct investment. This ambiguity has led some countries to record reinvested earnings as a capital inflow from the parent company (IMF, 1987, p. 36) and not as investment income.9

Asymmetries arise from cross-country differences in recording practices for a given transaction. (See footnote 3.) Given that both misclassifications and asymmetries reflect institutional practices, we model the persistence of errors they induce by including the lag of the discrepancy of the investment income account as an explanatory variable.

**Discrepancy in services**

The discrepancy in services arises from misrecordings in travel expenses, shipping, and other transportation services. This discrepancy is declining but that trend conceals the growing importance of errors in shipping (Table 1).10

Given that most countries have good records of their payments to foreign shippers, the discrepancy arises from underreporting of revenue by ship operators. Indeed, ship operators of the world’s largest fleets claim Greece, Hong Kong SAR, and countries in Eastern Europe as residence, but these economies do not report such earnings to the IMF (IMF, 1996, p. 146).11

To model this discrepancy we assume (i) that shippers seek to avoid income taxes and thus underreport their shipping revenues and (ii) that their propensity to underreport is directly related to the price of oil. We use this price because oil is an important commodity in maritime transportation and because the price of oil is correlated with the prices of other raw materials.12

Counteracting the effects of underreporting credits is the adoption of alternative modes of transportation. Specifically, declines in the physical weight of products allow their transportation using the growing air fleet. The associated tight security procedures virtually guarantee that all the items transported are accounted

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8Securitization takes place when direct loans by banks are replaced by underwriting commitments that do not appear on banks’ balance sheets.

9The IMF’s approach to measuring reinvested earnings requires direct questioning of multinational enterprises, which is difficult given the resources available. Because of these difficulties, the IMF recommends dropping this account from the global sum of investment-income accounts (IMF, 1987, p. 43).

10Shipment debits include the cost of freight, insurance, and those distribution services paid by the importer. Shipment credits include gross revenue on freight earned by vessels operated by residents of the reporting country regardless of the flag registry of the vessel.

11Moreover, the Greek balance of payments excludes the operations of the Greek fleet because the owners of that fleet do not reside in Greece and, as far as the IMF is aware, they are not residents in other countries either (IMF, 1987, p. 90).

12We also considered the IMF’s Commodity Price Index, but we did not find it to exert a significant influence on the behavior of discrepancies.
for, leaving little room for misreporting. To the extent that the decline in physical weight of products will continue, we use a time trend to capture how substitution away from maritime shipping reduces the scope for underreporting export services.

**Discrepancies in unrequited transfers**

Discrepancies in transfers arise from two sources: the recording asymmetries of workers' remittances and the exclusion of the intermediation by international donor organizations from balance of payments accounts. Asymmetries in recording remittances arise when (i) the host country treats temporary workers as residents, recording their remittances as unrequited transfers; and (ii) the country of origin also treats these workers as residents, recording their remittances not as unrequited transfers but as service exports (IMF, 1987, p. 104).

Intermediation by international donor agencies are excluded from countries' balance of payments because these agencies are not considered residents of any country; some of these agencies report to the IMF. As long as this intermediation operates without delays, the transaction by itself does not generate a statistical discrepancy. Over a given horizon, however, these institutions receive contributions for assistance in excess of their disbursements. The shortfall in disbursements is not recorded because these international institutions do not conform to the principle of residency (IMF, 1987, p. 103), giving rise to the account's discrepancy.

A convenient way of modeling this feature is to recognize that if the share of intermediation by international institutions declines, then so will the discrepancy it induces. The share of intermediation declines when donor countries provide their assistance directly to recipient countries and avoid the side effects of the intermediary role of international institutions (delays and residency). To this end we assume that OPEC members, who have been important donors in the past, are more likely to be donor countries the higher the price of oil. An increase in that price would, if we are correct, translate into greater assistance from OPEC, a reduction in the intermediation from international institutions, and a decline in the excess of debits over credits in transfers.

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**Table 1. Discrepancies in Services Account**

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<td>-55.7</td>
<td>-66.3</td>
<td>-63.6</td>
<td>-64.8</td>
<td>-60.9</td>
</tr>
</tbody>
</table>

Source: IMF's Balance of Payments Statistics.
Summary

Table 2 lists the factors we have identified to explain movements in the various discrepancies. This list is what we denoted as \( q' \):

\[
q' = (dY/Y, P_o, R, M_{us}, dum, trend)
\]

where \( dY/Y \) is world growth, \( P_o \) is the nominal price of oil, \( R \) is the nominal U.S. federal funds rate, \( M_{us} \) is the U.S. share in world imports, and \( dum \) is a dummy for European trade. To recognize the role of nonlinearities, we expand this list to include the squares of the price of oil and the interest rate as well as the interactions between these two variables and the trend. The resulting list of explanatory variables is denoted as \( Q' \):

\[
Q' = (\text{Intercept}, q', P_o^2, R^2, P_o \cdot \text{trend}, R \cdot \text{trend})
\]

We want to emphasize that Table 2 does not offer a list of zero restrictions on the coefficients of the reduced form. The simultaneous character of the model allows all exogenous variables to affect the discrepancies in all of the accounts and our statistical analysis allows for that possibility.

Parameter Estimation

Based on annual data from 1972 to 1998, Table 3 shows the least-squares estimates for the unrestricted reduced form, equation (4). The results reveal numerous \( t \)-ratios below the 5 percent critical value, which is not surprising given the relatively small number of degrees of freedom. Second, the maintained assumptions for the residuals are supported empirically.\(^1\)

Using a log-likelihood ratio test, we eliminate variables that are not jointly significant and reestimate the parameters of the restricted reduced form using FIML. According to the estimates, the data support the maintained assumptions for the residuals (Table 4) but persistence effects are small and limited to discrepancies in transfers.

Also, the dummy for the switch in European data methods is positive and significant, and suggests that the switch of recording practices of Europe raised the trade discrepancy by about 0.9 percentage points of world imports. The coefficient

\(^{1}\)We test for joint normality using the Jarque-Bera test: the statistic is distributed as \( \chi^2(n/2) \) where \( n \) is the number of equations. We test for joint serial independence with an F-test for the hypothesis that the coefficients for a VAR(1) of the estimation residuals are jointly equal to zero. We applied an ARCH test to each equation separately and the results cannot reject (not shown) the hypothesis of homoskedasticity. See Hendry and Doornik (1996) for details.
for the trend is positive and significant for services: the underreporting of credits due to maritime shipping is ameliorated by the growing role played by the alternatives to maritime shipping.

Nonlinearities (interactions and squared terms) have statistically significant coefficients, which call for model simulations to evaluate the effects of changes in the remaining variables; we use one-year shocks evaluated in 1998. Based on these simulations, a 10 percent increase in the price of oil lowers net credits in the global discrepancy by 0.03 percent of world imports (Table 5). This small effect reflects offsetting responses from the various accounts. An increase of the federal funds rate by 1 percentage point lowers net credits in the discrepancies for investment income and services; net credits in the overall discrepancy decline by 0.24 percent of world imports or about $13 billion. Raising the world’s growth rate by 1 percentage point increases net credits in the trade discrepancy; faster growth accelerates trade and accentuates the extent to which trade credits are recorded ahead of trade debits. The overall discrepancy experiences an increase in net credits of 0.64 percentage points of world imports or about $35 billion.

An increase of the U.S. share of world imports by 1 percentage point reduces net credits in the trade discrepancy by 0.25 percent of world imports ($14 billion) given that a greater fraction of world trade is being recorded by the country with the high-quality data. Higher U.S. imports also affect the investment income discrepancy given that financing an increase of U.S. imports involves an increase in foreign capital outflows. The underreporting of these outflows accentuates the understating of claims of the rest of the world on the United States and the associated interest receipts. Finally, to the extent that a fraction of the increase in U.S. imports is transported by the U.S. fleet, shipping credits that would not have been

<table>
<thead>
<tr>
<th>Variables: $q$</th>
<th>Factor Modeled</th>
<th>Account</th>
</tr>
</thead>
<tbody>
<tr>
<td>World growth: $dY/dY$</td>
<td>Recording delays</td>
<td>Trade</td>
</tr>
<tr>
<td>Oil prices: $P_o$</td>
<td>Valuation asymmetries, Tax evasion, Disbursement delays</td>
<td>Trade, Services, Transfers</td>
</tr>
<tr>
<td>U.S. interest rate: $R$</td>
<td>Tax evasion, Valuation asymmetries</td>
<td>Investment income, Trade</td>
</tr>
<tr>
<td>U.S. share of world imports: $M_w$</td>
<td>Quality differentials in data</td>
<td>Trade</td>
</tr>
<tr>
<td>Trend</td>
<td>Financial globalization, Transportation technology</td>
<td>Investment income, Services</td>
</tr>
<tr>
<td>Dummy Europe: $dum$</td>
<td>Methodological changes</td>
<td>Trade</td>
</tr>
</tbody>
</table>
Table 3. Estimates of Unrestricted Reduced Form—OLS: 1972–98

<table>
<thead>
<tr>
<th></th>
<th>Trade $D_{it}$</th>
<th>Services $D_{it}$</th>
<th>Investment $D_{it}$</th>
<th>Transfers $D_{it}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$D_{it-1}$</td>
<td>0.15</td>
<td>-0.26</td>
<td>-0.19</td>
<td>0.04</td>
</tr>
<tr>
<td>$D_{it-2}$</td>
<td>0.02</td>
<td>-0.06</td>
<td>0.08</td>
<td>0.46</td>
</tr>
<tr>
<td>$D_{it-3}$</td>
<td>-0.12</td>
<td>0.54</td>
<td>0.24</td>
<td>-0.12</td>
</tr>
<tr>
<td>$D_{it-4}$</td>
<td>-0.29</td>
<td>-0.24</td>
<td>0.31</td>
<td>1.05*</td>
</tr>
<tr>
<td>$dY/Y$</td>
<td>0.37*</td>
<td>-0.03</td>
<td>0.26*</td>
<td>0.01</td>
</tr>
<tr>
<td>$P_o$</td>
<td>-0.24*</td>
<td>-0.11</td>
<td>0.15</td>
<td>0.03</td>
</tr>
<tr>
<td>$R$</td>
<td>-0.49</td>
<td>0.50</td>
<td>0.57</td>
<td>0.09</td>
</tr>
<tr>
<td>$M_{us}$</td>
<td>-0.19</td>
<td>0.17</td>
<td>-0.34*</td>
<td>-0.04</td>
</tr>
<tr>
<td>Trend</td>
<td>-0.07</td>
<td>0.18</td>
<td>0.12</td>
<td>-0.04</td>
</tr>
<tr>
<td>dum</td>
<td>0.63</td>
<td>0.06</td>
<td>0.58</td>
<td>0.17</td>
</tr>
<tr>
<td>$P_o^2$</td>
<td>0.01*</td>
<td>0.00</td>
<td>-0.004</td>
<td>0.003</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.02*</td>
<td>-0.01</td>
<td>-0.012</td>
<td>-0.01</td>
</tr>
<tr>
<td>$P_o \cdot$ trend</td>
<td>0.00</td>
<td>0.003</td>
<td>-0.003</td>
<td>0.00</td>
</tr>
<tr>
<td>$R \cdot$ trend</td>
<td>0.01</td>
<td>-0.02</td>
<td>-0.02</td>
<td>-0.002</td>
</tr>
<tr>
<td>Intercept</td>
<td>6.67*</td>
<td>-5.70*</td>
<td>0.14</td>
<td>0.71</td>
</tr>
<tr>
<td>SER</td>
<td>0.297</td>
<td>0.354</td>
<td>0.337</td>
<td>0.181</td>
</tr>
</tbody>
</table>

*t-ratio above the 5 percent level. $P_o$ = oil price; $R$ = federal funds rate; $dY/Y$ = world growth; dum = dummy for European trade; and $M_{us}$ = U.S. share in world imports.

Hypothesis Testing

<table>
<thead>
<tr>
<th>Null Hypothesis</th>
<th>Test Statistic</th>
<th>Result (p-level)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Residuals are jointly normal</td>
<td>$\chi^2(8)$</td>
<td>11.845 (0.16)</td>
</tr>
<tr>
<td>Residuals are jointly serially independent</td>
<td>$F(16, 12)$</td>
<td>1.54 (0.23)</td>
</tr>
</tbody>
</table>

recorded are now being reported with an increase in net credits of the service discrepancy. Figure 3 compares historical values against model predictions.14

Judging by the mean absolute errors (MAE), either as a percentage of world imports and in U.S. dollars, the predictions of the model are close to historical values and the residuals are not one-sided. The exception is the transfer equation, which shows systematic deviations during the 1990s.

III. Model Applications

For the model developed here to be useful, it should detect large discrepancies when large discrepancies are known to occur. According to the IMF

14The predicted value is $\hat{\lambda}_t = \hat{\Pi}_t \cdot \hat{\lambda}_{t-1} + \hat{\Pi}_t \cdot Q_t$, and thus we use the model’s own predictions for the lagged endogenous variables instead of the historical values. We also examined one-step, out of sample, predictions over 1996–98, the results are comparable to those reported here.
Table 4. Estimates of Restricted Reduced Form—FIML: 1972–98

<table>
<thead>
<tr>
<th></th>
<th>Trade $D_{gt}$</th>
<th>Services $D_{st}$</th>
<th>Investment $D_{it}$</th>
<th>Transfers $D_{at}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$D_{g,t-1}$</td>
<td>—</td>
<td>—</td>
<td>—</td>
<td>0.35 (0.08)</td>
</tr>
<tr>
<td>$D_{t,t-1}$</td>
<td>—</td>
<td>—</td>
<td>—</td>
<td>0.25 (0.11)</td>
</tr>
<tr>
<td>$D_{h,t-1}$</td>
<td>—</td>
<td>0.25 (0.11)</td>
<td>—</td>
<td>0.84 (0.1)</td>
</tr>
<tr>
<td>$D_{u,t-1}$</td>
<td>—</td>
<td>—</td>
<td>—</td>
<td>—</td>
</tr>
<tr>
<td>$dY/Y$</td>
<td>0.40 (0.03)</td>
<td>—</td>
<td>0.23 (0.06)</td>
<td>—</td>
</tr>
<tr>
<td>$P_o$</td>
<td>-0.20 (0.03)</td>
<td>-0.11 (0.02)</td>
<td>0.11 (0.04)</td>
<td>—</td>
</tr>
<tr>
<td>$R$</td>
<td>-0.07 (0.03)</td>
<td>0.14 (0.06)</td>
<td>0.23 (0.03)</td>
<td>—</td>
</tr>
<tr>
<td>$M_{M}$</td>
<td>-0.25 (0.04)</td>
<td>0.13 (0.06)</td>
<td>-0.30 (0.06)</td>
<td>—</td>
</tr>
<tr>
<td>Trend</td>
<td>—</td>
<td>0.09 (0.03)</td>
<td>—</td>
<td>-0.02 (0.01)</td>
</tr>
<tr>
<td>dum</td>
<td>0.86 (0.03)</td>
<td>0.58 (0.18)</td>
<td>—</td>
<td>—</td>
</tr>
<tr>
<td>$P_o^2$</td>
<td>0.005 (0.0006)</td>
<td>—</td>
<td>-0.003 (0.001)</td>
<td>0.0007</td>
</tr>
<tr>
<td>$R^2$</td>
<td>—</td>
<td>—</td>
<td>—</td>
<td>—</td>
</tr>
<tr>
<td>$P_o$ trend</td>
<td>—</td>
<td>0.005 (0.001)</td>
<td>—</td>
<td>—</td>
</tr>
<tr>
<td>$R \cdot trend$</td>
<td>0.004 (0.001)</td>
<td>-0.01 (0.004)</td>
<td>-0.01 (0.002)</td>
<td>—</td>
</tr>
<tr>
<td>Intercept</td>
<td>5.56 (0.62)</td>
<td>-3.4 (1.15)</td>
<td>1.11 (0.70)</td>
<td>0.34 (0.25)</td>
</tr>
<tr>
<td>SER</td>
<td>0.281</td>
<td>0.339</td>
<td>0.329</td>
<td>0.173</td>
</tr>
</tbody>
</table>

Entries in parentheses are heteroskedasticity-corrected standard errors. $P_o =$ oil price; $R =$ federal funds rate; $dY/Y =$ world growth; dum = dummy for European trade; and $M_{M} =$ U.S. share in world imports.

Hypothesis Testing

<table>
<thead>
<tr>
<th>Null Hypothesis</th>
<th>Test Statistic</th>
<th>Result (p-level)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Residuals are jointly normal</td>
<td>$\chi^2 (8)$</td>
<td>3.57 (0.89)</td>
</tr>
<tr>
<td>Residuals are jointly serially independent</td>
<td>$F (16, 37)$</td>
<td>0.84 (0.64)</td>
</tr>
<tr>
<td>Overidentifying restrictions hold</td>
<td>$\chi^2 (31)$</td>
<td>34.57 (0.30)</td>
</tr>
</tbody>
</table>

(IMF, 1997, p. 9), the change in European methodology for collecting trade data induced a major increase in the trade discrepancy, and the question is whether the model detects it as such. Thus, the null and alternative hypotheses are:

$$H_0: E(D_i) = D_i^c$$

and
where $D_t^e$ is the observed value for the discrepancy at date $t$, $t = 1993–98$. A rejection of the null hypothesis means that the model identifies as large the change in European methods to collect data.

Being able to identify statistically large discrepancies is a necessary but not a sufficient condition to judge the model's usefulness. Specifically, this switch is credited with an increase in net credits of the trade discrepancy. Thus, model predictions that do not use post-1992 data for parameter estimation should understate net credits in the trade discrepancy. Also, the recorded increase in the trade discrepancy tended to offset the discrepancy in investment income (see Figure 2) and induced a seemingly small recorded global discrepancy. Thus, model predictions that exclude the switch should show a worsening of the global discrepancy. Finally, we need evidence of stability in the model's parameters to avoid confusing the effects of parameter instability with the effects of changes in data collection methods.

To implement the test, we start by estimating the model’s parameters with data through 1992, which excludes post-switch observations. Comparing the estimation results with those based on the full sample reveals that both sets of parameter estimates are virtually identical (Table 6). With one exception, neither sign nor statistical significance of the estimates change, as a result of using the shorter sample.\footnote{The exception is the coefficient for interaction of trend and oil prices in the service equation: significant with the full sample and insignificant otherwise.} Also, the maintained assumptions for the residuals are supported empirically. Moreover, relative to the parameters of the unrestricted reduced form estimated with data through 1992, the log-likelihood ratio test does not reject the same set of zero restrictions. Overall, this evidence rules out parameter instability as a factor in a finding of large discrepancies.\footnote{We examined the sensitivity of the estimated parameters (unconstrained and restricted reduced forms) to using the London interbank offered rate (LIBOR) on six-month instruments instead of the U.S. federal funds rate. We find that the point estimates are quite robust but the zero restrictions for the shorter sample are rejected for the model based on the LIBOR. Thus one cannot determine whether ex post tests from the LIBOR-based model are due to parameter instability.}
Given the coefficient estimates, we use dynamic simulations to generate predictions for 1993–98, with 1992 as the initial condition:

\[
\hat{\Lambda}_t = \hat{\Pi}_{s_1} \cdot \hat{\Lambda}_{t-1} + \hat{\Pi}_{s_2} \cdot Q_t
\]

and

\[
\hat{E}(D_t) = 1' \cdot \hat{\Lambda}_t \quad \text{for } t = 1993–98,
\]

where the subscript \(s\) denotes that the estimates use the short sample (1972–92). If
Jaime Marquez and Lisa Workman

Table 6. Parameter Estimates with FIML, 1972-98 and 1972-92

<table>
<thead>
<tr>
<th></th>
<th>Trade $D_{it}$</th>
<th>Services $D_{it}$</th>
<th>Investment $D_{it}$</th>
<th>Transfers $D_{it}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$D_{it-1}$</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$D_{it-1}$</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$D_{it-1}$</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$D_{it-1}$</td>
<td>0.25*</td>
<td>0.33*</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$R \cdot \text{trend}$</td>
<td>0.004*</td>
<td>0.004*</td>
<td>-0.01*</td>
<td>-0.01*</td>
</tr>
<tr>
<td>Trend</td>
<td>-0.01*</td>
<td>0.09*</td>
<td>0.17*</td>
<td>-0.02*</td>
</tr>
<tr>
<td>$P_0$</td>
<td>-0.20*</td>
<td>-0.25*</td>
<td>-0.11*</td>
<td>-0.07*</td>
</tr>
<tr>
<td>$P_0^2$</td>
<td>0.005*</td>
<td>0.01*</td>
<td></td>
<td>-0.003*</td>
</tr>
<tr>
<td>$R$</td>
<td>-0.07*</td>
<td>-0.10*</td>
<td>0.14*</td>
<td>0.16*</td>
</tr>
<tr>
<td>$dY/Y$</td>
<td>0.40*</td>
<td>0.36*</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$P_0 \cdot \text{trend}$</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$M_{as}$</td>
<td>-0.25*</td>
<td>-0.21*</td>
<td>0.13*</td>
<td>0.16*</td>
</tr>
<tr>
<td>Intercept</td>
<td>5.56*</td>
<td>5.83*</td>
<td>-3.4*</td>
<td>-4.44*</td>
</tr>
<tr>
<td></td>
<td>0.34</td>
<td>0.36</td>
<td></td>
<td></td>
</tr>
<tr>
<td>SER</td>
<td>0.281</td>
<td>0.294</td>
<td>0.339</td>
<td>0.356</td>
</tr>
</tbody>
</table>

*Statistical significance at the 5 percent level.

Hypothesis Testing with Observations Ending in 1992

<table>
<thead>
<tr>
<th>Null Hypothesis</th>
<th>Test Statistic</th>
<th>Result ($p$-level)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Vector of residuals is normal</td>
<td>$\chi^2(8)$</td>
<td>8.19 (0.41)</td>
</tr>
<tr>
<td>Vector of residuals is serially independent</td>
<td>$F(16, 22)$</td>
<td>1.77 (0.11)</td>
</tr>
<tr>
<td>Overidentifying restrictions hold</td>
<td>$\chi^2(29)$</td>
<td>41.14 (0.07)</td>
</tr>
</tbody>
</table>

\[ |D_i^t - \hat{E}(D_i^t)| > \hat{\delta}_i = 2 \cdot \sqrt{\text{vär}(D_i^t)}, \]

then the observed discrepancy differs significantly from its expected value and we interpret this result as a large discrepancy.

Figure 4 reports the results, which reveal statistically significant underpredictions in the trade discrepancy and overprediction for the overall discrepancy, just as one expects. Overall, the results show that the model's confidence intervals are narrow enough to detect the European switch in data recording procedures as a major development. By itself, this finding does not constitute an endorsement of the approach. However, not being able to identify a known significant change would question its usefulness.
Figure 4. 95 percent Confidence Bands for Forecast Discrepancies, 1993-98
(percent of World Imports)

We now test whether the IMF's figures for the overall discrepancy over 1999-2001 are large. To this end we use the IMF's extrapolations for the exogenous variables (Table 7). The results suggest that the IMF's current account predictions embody a global discrepancy $D_i$ that is significantly different from our model's expectation of that discrepancy, $E(D_i)$ (Figure 5). This finding calls, according to our approach, for a rethinking of the current account forecasts for the individual countries.

17We apply the growth rates for oil prices reported on IMF (2000, p. 277) to a 1998 oil price of $12.30 a barrel. For the interest rate, we combine the projections for the real world long-term interest rate (IMF, 2000, p. 277) with the projections for U.S. CPI inflation (IMF, 2000, p. 215). For the world growth rate, we use the projections reported on IMF (2000, p. 277).
IV. Conclusions

This paper offers a framework for judging when the discrepancy embodied in current account forecasts is large. The first step in implementing this notion is to develop an econometric model that generates the discrepancy’s conditional expectation. The second step is to compare this model-based forecast with the discrepancy embodied in countries’ current account forecasts. If the gap in discrepancies is below a critical value, then the discrepancy embodied in the countries’ current account forecasts is not large. Otherwise, the discrepancy is large and calls for a careful reexamination of the associated current account forecasts.

Econometric modeling of these discrepancies is not the obvious first step in addressing global discrepancies. The first obvious step is to design fixed rules to allocate the discrepancies across countries. Though appealing, reliance on rules is
MODELING THE IMF'S STATISTICAL DISCREPANCY

at odds with the large and often mutually offsetting movements in the discrepancies of the components of the current account and thus might further distort countries’ current account forecasts. Econometric modeling of discrepancies, unusual as it is, offers a well-defined framework for determining when discrepancies in the global current account are unusual.

APPENDIX

Data

Sources

The published data on discrepancies come from various issues of the IMF’s Balance of Payments Statistics. One needs to use several issues because the data are subject to large revisions. For example, the value for the 1994 discrepancy ranges from -$75 billion in the 1995 data release to less than -$50 billion in the 1998 data release. Thus each observation in Figure 1 comes from the most recent release containing data for that year. The exact dates of the releases can be found in the working paper version of this paper located at http://www.federalreserve.gov/pubs/ifdp/2000/678/ifdp678.pdf

The data for the other series also come from the IMF and the corresponding locators are given below:

Spot world U.S.$ per barrel for oil: IMF 00176AAZ
Federal funds rate: IMF 11160B
World imports: IMF 00171D
U.S. imports: IMF 11171D

Properties

The regressions presented earlier assume that the variables have the same degree of stationarity. To test whether that property holds, we use an Augmented Dickey-Fuller (ADF) test with a constant and three lags. The evidence suggests that one cannot reject the hypothesis that all of the variables used here are integrated of order one.

Fixed Rules Allocation

We now examine the issues involved in using fixed rules to allocate the “excess” or unexpected discrepancy among the countries and regions of interest. We start by considering hypothetical forecasts of global current accounts (Table A1), which embody a discrepancy of -$140 billion. Suppose now that the model predicts a global discrepancy of -$40 billion and the question is how to allocate the excess discrepancy (-$100 billion) across the countries and regions in Table A1.

To this end, we assume that the global current account discrepancy stems from discrepancies in trade, shipping, and portfolio income; we ignore discrepancies in transfers. Then we allocate the discrepancy across countries using a list of stylized rules:

1. Take the forecast of the U.S. current account as given.
Augmented Dickey-Fuller Tests, 1972–98

<table>
<thead>
<tr>
<th>Exogenous Variables</th>
<th>ADF</th>
<th>Discrepancy</th>
<th>ADF</th>
</tr>
</thead>
<tbody>
<tr>
<td>Price of oil</td>
<td>-1.91</td>
<td>Trade</td>
<td>-1.09</td>
</tr>
<tr>
<td>Federal funds rate</td>
<td>-1.10</td>
<td>Income</td>
<td>-1.88</td>
</tr>
<tr>
<td>U.S. share of world imports</td>
<td>-2.05</td>
<td>Service</td>
<td>-0.63</td>
</tr>
<tr>
<td>World growth</td>
<td>-2.88</td>
<td>Transfers</td>
<td>-2.02</td>
</tr>
</tbody>
</table>

Note: 5 percent value is -2.997; 1 percent value is -3.75.

Table A 1. Hypothetical Allocation of the Current Account Discrepancy for 2001

<table>
<thead>
<tr>
<th>Adjustments</th>
<th>Original (1)</th>
<th>Trade (2)</th>
<th>Shipping (3)</th>
<th>Inc. (4)</th>
<th>Total Adjustment (5)</th>
<th>Final (6)</th>
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<td>0</td>
<td>0</td>
<td>0</td>
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<tr>
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<td>0</td>
<td>+17</td>
<td>+17</td>
<td>159</td>
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<tr>
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<td>122</td>
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<td>0</td>
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<td>-40</td>
<td>+71</td>
<td>+68</td>
<td>+99</td>
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</tbody>
</table>

2. Assume that the trade discrepancy is +$40 billion (credits over debits). This estimate stems from the underreporting of imports in Euroland’s business surveys (IMF, 1999, p. 4). Thus we lower the balance for Euroland by $40 billion (Table A 1, column 2).

3. Assume that the shipping discrepancy is −$71 billion (debits over credits, 1997 value). Allocate this discrepancy to Eastern Europe (7 percent), Asia—as housing offshore banking centers—(25 percent), Middle East and Africa (8 percent), and other developing countries (60 percent). Assuming that “Other” includes Greece, we assign three-fourths of the 60 percent to Greece (IMF, 1987, Table 72).

4. Assume that the portfolio income discrepancy is −$68 billion (debits over credits), which is half of the reported 1997 discrepancy; we use half of the income discrepancy to keep the overall discrepancy in the target range of $40 billion. Allocate the −$68 billion discrepancy as follows: 50 percent to industrial countries (25 percent for Japan and 25 percent for Euroland), 6 percent to Middle East and Africa, 18 percent to Asia, and 26 percent distributed evenly between “Other” and Latin America (IMF, 1987, Table 56).

Column 6 reports the final current account forecasts after the adjustments. The main change is the increase in the surplus of “Other” industrial countries due to the allocation of shipping exports to Greece. Otherwise, the allocation leaves intact the surplus/deficit status of each current account forecast.
MODELING THE IMF'S STATISTICAL DISCREPANCY

Table A1 shows that allocating the excess discrepancy across countries with fixed rules is feasible but their use suffers from three limitations: ambiguities, disincentives, and lack of generality.

1. Ambiguities: There are many strategies to allocate the excess discrepancy, but there is no generally accepted method to discriminate among them. This ambiguity is not the result of using the model to estimate the expected discrepancy but rather stems from the abundance of degrees of freedom available in an exercise that allocates a whole to many parts.

2. Disincentives: If economists' forecasts of current accounts are bypassed by a fixed rule, then these economists could lose interest in forecast accuracy as their predictions can be overruled by the rule. Even if these economists do not lose their incentives, they will face a harder than needed task in trying to account for (and reduce) that forecast error. Indeed, fixed rules do not revise the paths for the forcing variables but rather the forecasts themselves, making detection of the source of the forecast error all that much harder.

3. Lack of generality: Organizational differences across institutions give rise to different allocation rules to accommodate institutional idiosyncrasies. What is suited to the IMF’s purposes need not be suitable for the Organization for Economic Cooperation and Development or the World Bank.

These limitations are not meant to suggest that reallocation is not needed but, rather, to suggest that using a fixed rule is fraught with pitfalls that do not necessarily enhance forecast accuracy. There is a need to reallocate an excess discrepancy, but this process needs to be tailored to the particulars of the institution such as forecast assumptions, regional aggregates, and other features that might be pertinent. Our paper is not a contribution on how to allocate a discrepancy but rather on how much needs to be reallocated and, in this task, we hope the results of the paper are of general interest.

REFERENCES


Reform and Growth in Latin America: All Pain, No Gain?

EDUARDO FERNÁNDEZ-ARIAS and PETER MONTIEL*

This paper addresses the adequacy of post-reform growth in Latin America in the 1990s on the basis of international comparisons as well as historical and other relevant standards. The paper analytically explores and empirically tests a number of hypotheses to explain the perceived dissatisfaction with growth performance in the region. We find that there is no “growth puzzle” in Latin America. Growth has not been higher in the post-reform period not because of a failure of reforms to yield the growth payoff that they should have been expected to do on the basis of international experience, but because of the combination of an unfavorable external environment with the insufficient depth and breadth of reform. We also estimate the long-run growth payoff of macroeconomic reforms, the additional gains that can be achieved by deepening this first generation of reforms, and the potential payoff from broadening the scope of reform into a second generation of reforms encompassing deeper structural and institutional areas. [JEL 011, 019, 042, 054]

The wave of market-oriented reforms that has swept developing countries in recent years has been most visible in Latin America, where such reforms have signified a particularly sharp break with the previous policy regime. The implementation of such a drastic change in policies has been politically difficult and owed in no small measure to the widespread expectation that the new policy

*This paper does not necessarily represent the views of the Inter-American Development Bank’s management or its Board of Executive Directors. The authors would like to thank Martin Lauer for excellent research assistance and Felipe Barrera for both helpful comments and assistance. The paper has benefited from comments by Michael Gavin, Ricardo Hausmann, Eduardo Lora, Ernesto Talvi, and participants at a conference at the Brookings Institution.
regime would usher in a new era of rapid and widespread economic growth. This era would reverse the experience of the “lost decade” of the 1980s, which left many countries in the region with living standards below those achieved at the beginning of the decade.

Despite the extent and depth of the reforms, however, the acceleration in economic growth recorded by countries in the region to date has been modest, and has particularly fallen short of the standards of success established by some observers. Latin American countries as a group, for example, have not achieved the rates of growth in the post-reform period that they had previously attained during the 1970s; have not managed to grow as fast as the East Asian “miracle” economies to which they are often compared; and have not achieved the absolute rates of growth considered by informed observers to be necessary for achieving progress in ameliorating a variety of social problems that were aggravated during the last decade.¹ Moreover, some observers have interpreted whatever gains have been achieved on the growth front as potentially transitory, reflecting a temporary boom generated by recovery from crisis or by the excessive exuberance of international creditors.²

Does this imply that the reforms have failed and should be reconsidered as an instrument to achieve their primary growth objective? Obviously, a simple comparison of actual to desired growth rates is not sufficient to answer this question, since the desired growth target may simply represent an excessively ambitious policy objective. But even if the desired growth rates are reasonable, an indictment of the reforms implemented to date for failing to reach them may nevertheless be unwarranted. This is partly because the size of the growth payoff depends not only on the merits of reforming but also on the magnitude of the actual reform effort, and partly because growth is also affected by variables other than those influenced by recent reform efforts in the region.

Nonetheless, it is important to assess whether the reforms are “working” in the sense of delivering an appropriate growth payoff. There are several ways to approach this question. One could measure, for example, the growth impetus of reform. This can be computed as the product of the marginal effects of reforms estimated from international experience and the actual changes in the set of variables measuring reform in Latin America. This can be evaluated by some standard of the adequacy of the growth payoff of reform. It may fall short of that standard during the post-reform period, either because the marginal effects on growth of unit changes in the set of reform measures implemented by countries in the region have not been of the expected magnitude, or because the reform variables did not register changes of sufficient magnitude.

This “growth impetus” approach essentially asks whether the policies undertaken have delivered the results—that is, the growth acceleration—that they could

¹The World Bank, for example, has estimated that the region needs to grow at an average annual rate of 6 percent to generate the resources required to cope with social and infrastructure needs. See Edwards (1995).

reasonably have been expected to do. But it does not specifically address whether
the reforms undertaken were in principle capable of attaining the desired growth
rates. Measured in this way, the reforms could have "worked" (in the sense of
having delivered an "appropriate" growth increase) while nonetheless leaving
growth rates in the region far short of their desired levels. Simply measuring the
growth impetus associated with the reforms would provide no indication as to why
this might be so.

A broader approach to the question takes the desired growth outcome as
its point of departure and seeks to account for the gap between actual and
desired outcomes. A failure of policies to deliver the growth response that they
could reasonably have been expected to do, an insufficient magnitude of
adjustment in the reform variables, and/or unfavorable values of growth deter­
minants other than those captured in the set of reform measures could all
contribute to such a gap. To the extent that the factors contributing to the exis­
tence of a gap between actual and desired post-reform growth rates can be
identified—and their individual contributions to the magnitude of this gap
measured—this broader approach, which we label the growth gap approach,
has the advantage that it can potentially tell more about the possibility of iden­
tifying and adopting measures to close the gap between actual and desired
growth rates.

A paper by Easterly, Loayza, and Montiel (1997, hereafter ELM) imple­
mented the growth impetus approach. They found that the response of economic
growth to reform in Latin America has not, in fact, been disappointing during the
reform period. Rather, given the estimated effects of the reform variables on
economic growth and the actual changes that the values of the reform indicators
have undergone in Latin America during recent years, the change in the observed
rate of growth in the region was not statistically different from what would have
been predicted on the basis of international evidence.

This paper extends ELM's work in two ways. First, we broaden the scope
of their analysis of the growth impetus approach in several important direc­
tions—that is, by extending the sample, allowing for dynamic effects of the
reform measures, and broadening the scope of the reform indicators. We then
present new empirical evidence designed to test whether their conclusion that
the reforms have "worked" in the narrow growth impetus sense is robust to
these extensions. We find that it is. We then use our more general specification
to produce two alternative measures of the contributions of reform to changes
in growth performance within the region. The first measure involves estimating
the contribution of our broadest set of reform measures to increasing the long­
run growth rates both of individual Latin American countries and of the region
as a whole. We estimate that the reforms implemented to date will have persis­
tent—albeit quite different—growth effects for almost all of the countries in
our sample, as well as a significant positive effect on sustainable growth for the
region in the aggregate, estimated at 1.8 percent a year. The second measure of
performance is based on the observed growth acceleration in the region
between 1991-95 and 1986-90. We conclude that the growth acceleration
payoff to the reforms was significant, but was partly offset by an adverse external environment during the reform period.

We then implement the growth gap approach to assessing the adequacy of the growth effects of reform. To do so, we require a measure of the "desired" growth rate. We use three alternative definitions, based on growth in the region in the 1970s and contemporaneous growth in two other regions, and decompose the corresponding growth gaps based on our results. Regarding the historical comparison for Latin American countries, we find that in most countries current macroeconomic policy is significantly more conducive to growth and, in the absence of a substantial deterioration of exogenous factors, would have led to surpassing the growth target. The central finding in the cross-region comparison is that, in the aggregate, while there is room for intensifying reforms in the directions already implemented to substantially increase growth, achieving the desired growth rates is likely to require a broadening of the scope of the reform effort. We conclude by summarizing the results, offer some tentative interpretations, and point to potentially fruitful areas of future research.

I. Reexamining the Evidence

This section consists of four parts. After discussing our statistical methodology, we take up the extensions listed above one at a time, and examine in each case whether the conclusion that the growth payoff to reform has not been disappointing in the narrow sense proves to be robust to the specific extension.

Statistical Methodology

Our empirical methodology is based on the estimation of panel growth regressions. The panel consists of a sample of 69 countries, 18 of which are in Latin America, with data spanning the period 1961–95. This is the largest panel of countries for which relevant information is available. We divided the period into seven five-year subperiods, two for each decade, and constructed five-year averages of our variables where appropriate, both contemporaneous and lagged. The use of a large panel of countries over an extended period of time allows us sufficient degrees of freedom to enrich the menu of variables used to measure reform and to control for nonreform growth determinants, as well as to engage in some exploration of growth dynamics associated with the adoption of macroeconomic reforms.

Since one of our objectives in this paper is to use our estimated relationship between macroeconomic policies and economic growth to implement the growth gap approach, we have conducted our estimation in level form to permit us to esti-
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mate the growth effects of country-specific "permanent" structural factors through
country-specific intercept terms (that is, we adopt a "fixed-effects" model).
Consequently, the resulting estimation of the growth contribution of the macroe-
comic reform variables included in the regression is not distorted by the attribu-
tion of permanent cross-country differences in growth to differences in
time-invariant aspects of the policy environment. We controlled for time-specific
growth effects emanating from changes in the external economic environment
(resulting from technological, financial, or other sources) across periods by
including time dummies. Finally, our explanatory variables also included a set of
traditional cross-country growth determinants as control variables. The basic
static estimation equation is thus:

\[ g_{it} = s_i + w_t + ar_{it} + bc_{it} + u_{it}, \]  

where \( g_{it} \), the explained variable, is the real per capita growth rate of GDP in
country \( i \) (\( i \) ranging from 1 to 69) during period \( t \) (\( t \) ranging from 1 to 7). The first
two terms are the structural country dummy and the time dummy, respectively.
Macroeconomic reform variables are denoted by \( r \) and control variables by \( c \). Our
reform variables included the rate of inflation, the share of government consump-
tion in GDP, the ratio of broad money to GDP, the black market premium, and a
conventional measure of openness (the ratio of the sum of exports and imports to
GDP). The set of control variables consisted of the GDP per capita and the level
of educational attainment inherited from the previous period (that is, at the begin-
ing of each five-year period), and the international terms of trade prevailing for
each country on average during the current period. The empirical counterparts of
all of the variables are described in the data appendix.

Our reform variables are among the most widely used policy indicators in the
cross-country growth literature. Since these variables are in effect macroeconomic
outcomes themselves, however, endogeneity is a potentially important problem
under Ordinary Least Squares (OLS) estimation, possibly leading to a magnifica-
tion of estimated growth effects through reverse causality. In principle, this

\[ \text{In contrast, purely cross-country studies cannot control for structural country-specific differences}
\text{unless untestable statistical assumptions are made to justify a "random-effects" model—that is, the}
\text{absence of correlation between the structural terms and the explanatory variables. In this context, this}
\text{practice appears particularly worrisome because, as explained later, our testing of that assumption with}
\text{this panel indicates that random-effect growth models are biased.}
\text{Note that since investment is not included among the control variables in the regressions reported}
\text{below, growth effects should be interpreted as overall effects, inclusive of effects operating through invest-
\text{ment rates.}
\text{The estimated effects of "initial" GDP per capita and education should be interpreted with care}
\text{because they refer to convergence effects within a five-year period, especially when compared with estimates}
\text{from cross-section regressions.}
\text{As is customary, many of the explanatory variables are used in logarithmic form and their corre-
\text{sponding coefficients have a semi-elasticity interpretation. However, neither the statistical significance of}
\text{these variables nor their overall growth effects is sensitive to this specification.} \]
problem could be addressed (at least in part) by using as reform indicators variables that more narrowly capture specific policy instruments. However, aside from the availability of such variables, the presumed superiority of using policy reform variables is weakened by the fact that policies that lack credibility are ineffective and would introduce biases if credibility is not controlled for, while outcome variables implicitly filter out ineffective policies. Moreover, the very presumption of positive biases under OLS is itself doubtful under close scrutiny. The traditional argument that outcome reform variables yield positive endogeneity biases because of reverse causation misses the opposite effect attributable to the so-called crisis hypothesis, according to which crises help the implementation of reform, thus inducing negative reverse causality.

Nevertheless, to test the appropriateness of OLS estimation, we conducted instrumental variable estimation using the lagged values of inflation and financial depth as instruments (as well as the rest of the explanatory variables). Under the reasonable assumption that these instruments are exogenous, IV estimation is consistent. A comparison of the point estimates of the macroeconomic reform variables, however, did not point to a systematic OLS magnification bias, since the IV estimates of the effects of openness, government consumption, and financial deepening—three out of the five policy proxies—were larger than the OLS estimates. In fact, the point estimate of the overall growth effect of reform obtained under both estimation methods is almost identical. A Hausman specification test showed very strongly that the consistency of OLS could not be rejected on the basis of this IV estimation. At the same time, the accuracy of the IV estimates was clearly lower than that of OLS; this implies that, all things considered, OLS appears to be the best choice between the two estimation methods.

Because the imprecision of the IV estimates may derive from the poor quality of the exogenous instruments available in this case, we also resorted to indirect evidence to satisfy ourselves that it was not worthwhile to complicate the statistical approach to the problem and that OLS would be reasonably unbiased, by comparing our results to those of ELM. We found that our simple methodology was able to closely reproduce the results of the more sophisticated econometric methodology employed in ELM to implement IV estimation in a dynamic panel. For all of these reasons, we used OLS as our estimation method.

A second econometric issue concerns the appropriate technique for panel estimation. Under the assumption that the country-specific effects are orthogonal to the regressors $r$ and $c$, a random-effects model, in which the country-specific effects are controlled for within the regression error term, is consistent and more efficient than the fixed-effects model posited above. However, a Hausman specification test shows that the estimations from the fixed-effects model and those from the random-effects model are significantly different at extremely high confidence levels, thus indicating that the random-effects model yields inconsistent estimates in this case. In other words, the validity of the orthogonality assumption required for consistency is rejected with virtually total confidence. Therefore, the best choice between the two methods to analyze this panel appears to be fixed-effects OLS. An important implication is that the convenient use of the orthogonality assumption in the context of cross-section growth regressions, in which context
fixed-effects estimation is not feasible and the consistency of random effects cannot be tested, is not only unwarranted but very likely invalid.

In preliminary estimates of the basic equation, the openness variable failed to be statistically significant at conventional confidence levels (p-value of 25 percent), while the other four macroeconomic reform variables had estimated coefficients with the theoretically appropriate signs that were statistically significant at least at the 97 percent confidence level. Separating the effects of openness by region, the variable entered with the appropriate sign for all regions except Africa, where it was statistically significantly negative. We concluded that the failure of the openness variable in this panel was associated with the role of the variable for the African countries. One possible explanation is the effect of compensatory external financial aid to Africa, which may induce a negative correlation between growth performance and openness. Given the ambiguity of interpretation, we chose to eliminate openness from the basic static specification.

Basic Static Equation

The results of estimating the basic equation without this variable are displayed in Table 1 (specification 1). Notice first that the Durbin-Watson statistic adjusted by the 68 cross-country residual differences of this panel (2.05) strongly supports the hypothesis of zero serial autocorrelation of residuals. Therefore OLS is efficient and the reported precision of the estimations is reliable. All the stabilization and structural reform variables are correctly signed, and are highly significant (p-value of less than 4 percent). In particular, we find a substantial and statistically significant positive marginal growth impact associated with lower public consumption, lower inflation, financial deepening, and exchange rate unification. Control variables all have the expected signs as well: positive for education and changes in the terms of trade, and negative for initial per capita GDP. All but that for the education variable (p-value of 40 percent) are also highly significant. Changes in the external environment—captured by the time dummies—appear to have had growth effects that were both substantial and statistically significant. In particular, the external growth environment in the 1990s appears to be about as negative as in the first half of the 1980s, when the debt crisis hit—down by about 1 percentage point relative to the second half of the 1980s. This finding is consistent with other studies and also with casual observation: relative to the previous five-year period, growth slowed in all regions, including East Asia, except in Latin America. Not only in Africa, but also in member countries of the Organization for Economic

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8 Though we chose to eliminate openness as a separate explanatory variable, we capture it below through the Structural Policy Index, as discussed later. Another strategy could have been to eliminate Africa from the study and keep openness as an explanatory variable. Sensitivity analyses comparing runs with and without African countries showed that the removal of Africa does not introduce statistically significant changes in the coefficients of interest. It marginally dampens most of the estimated coefficients (albeit it magnifies the effect of exchange rate unification), but their significance and the overall conclusions of the exercise remain. We chose to keep Africa to gain in statistical precision.

9 The so-called speed-of-convergence parameter associated with the latter is consistent with that found in ELM.
Cooperation and Development (OECD) growth slowed by more than 1\(\frac{1}{2}\) percentage points.

More important, the evidence suggests that the growth response to recent reform in Latin America is not inferior to what international experience would lead us to predict. In fact, growth in Latin American countries during the first half of the 1990s actually exceeds what should be expected according to these estimates (the average residual is positive for countries in the region during the last five-year period, amounting to 0.53 percentage points of growth). Moreover, if a Latin American dummy for the reform period 1991–95 is added, it comes out posi-
tive and has a p-value of 13 percent. The implication is that there is no reason for disappointment with the growth response to the reforms undertaken during the 1990s. The full extent of the underlying growth progress owing to reform is partially hidden by an adverse external environment, which accounts for a growth downturn of about 1 percentage point \((-0.82 + -0.14 = -0.96)\). These results suggest that expectations that do not take into account the adverse external environment would be erroneous in finding the post-reform growth acceleration to be disappointing. If anything, the evidence is consistent with the view that recent reform in Latin America led to surprisingly fast growth.10 These results are broadly similar to those obtained by ELM.

Two additional tests lend further support to this conclusion. First, the hypothesis that the coefficients of the four policy variables in Latin America were statistically equal to those in the other regions in the panel was directly tested and could not be rejected. Second, the results indicate that, while growth in the region in the late 1980s is well explained by the model, growth significantly exceeded the model’s prediction for the 1990s. In other words, it is only in the 1990s that the Latin American residuals are sizable (their average in the late 1980s was actually negative, at \(-0.04\)). This suggests that the excess growth identified in the recent period is due to actual acceleration (of 0.57 points as measured by average residuals) rather than model misspecification inducing a systematic Latin American misfit.

Basic Dynamic Equation

The previous specification does not address the dynamics of the growth response, implicitly assuming that a five-year period is sufficient for the long-run implications of changes in the explanatory variables to become manifest. However, stabilization and structural reform typically set in motion complex business cycle dynamics. Furthermore, growth theory suggests that macroeconomic policy has, at least to a partial extent, an income level effect that translates into a transitory growth effect. If five years are not enough to eliminate short-run fluctuations, it may very well be that long-run effects are overestimated in the static panel regression.11 On the other hand, if some of the growth effects are worked out only after a long delay, then the above estimates based on five-year averages would underestimate the effect of reform. In either case, the growth equations estimated in Table 1 would have been misspecified by omitting lagged values of the reform variables.

This possibility is testable. If valid, it has the important implication that future growth performance will differ from current performance, even if no further reforms are enacted. Depending on the direction of this effect, this may call for

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10 Note that, while the latter is consistent with Krugman’s (1995) view that foreign investor exuberance may have caused growth in the region to be temporarily high in the 1990s, the former implies that this cannot be combined with the observation that growth has fallen short of expectations to draw the conclusion that the reforms have failed.

11 This would be the case if, for example, as established in Inter-American Development Bank (1996), short-run economic booms follow stabilization and structural reform.
REFORM AND GROWTH IN LATIN AMERICA

either less or more policy response to a growth performance that is deemed inadequate, compared with a situation in which long-run growth effects materialize within a five-year period. To address these issues, we specified our basic equation in dynamic form, adding the lagged values of the four macroeconomic reform variables \( (r') \):

\[
g_{it} = s_{it} + w_{it} + ar_{it} + a'r_{it} + be_{it} + u_{it},
\]

The long-run effect of reform, that is, the effect that would prevail if reform were sustained, is given by the sum of the contemporaneous and the delayed impacts. If the latter is negative, some of the growth gains would be lost in the future. If positive, additional growth would occur effortlessly.

The results of including lagged reform variables in the basic equation are reported in Table 1 as specification 2.12 As shown in the table, the coefficients of the lagged reform variables are all opposite in sign to those of the corresponding contemporaneous variables. However, the lagged coefficients are quite small in absolute value, leaving substantial positive long-run effects for each of the four reform variables. None of the delayed effects is clearly statistically significant individually, but they are strongly significant jointly. This dynamic specification marginally improves upon the static one according to standard statistical measures, as well as with regard to the qualitative features of the results. As measured by the adjusted R-square, the fit of this dynamic specification is slightly better than that of its static counterpart. Moreover, in the dynamic specification, education is statistically significant (p-value of 14 percent). The evidence thus suggests the presence of a minor partial offset to the beneficial growth effect of stabilization and reform after five years.

The previous conclusion, however, that the growth response to reform in the 1990s has not been disappointing (in the narrow sense) continues to hold in the dynamic specification. The hypothesis that the coefficients associated with the four reform variables, both contemporaneous and lagged, are equal for Latin America and the rest of the world cannot be rejected at conventional confidence levels (p-value of 27 percent). Taking into account policy dynamics, the observed growth acceleration in Latin America still exceeded expectations by an average of 0.43 percentage points (the average residual in the late 1980s remains at -0.04). With this new dynamic specification, however, growth in the 1990s is statistically within what the model would predict. The Latin America 1991–95 dummy loses statistical significance in the dynamic specification (p-value of 22 percent).

In this case as well, the openness variable, both current and lagged, can be jointly rejected with a log-likelihood ratio test at the 15 percent level. Discriminating by region, the long-run effect of openness in Latin America was not significantly positive. We chose to eliminate the openness variable from this basic dynamic regression too.
Extending the Reform Coverage

Our basic equations, and the empirical panel literature in general, use a relatively short list of macroeconomic policy reform outcomes to capture the growth impact of a wide range of reform policies as a substitute for the direct measurement of policy instruments. While many policy stances can be expected to be proxied reasonably well by their predictable and easily measured outcomes, some potentially important reforms, such as tax reform or even trade reform, may not be adequately captured. The lack of broad and consistent information on actual policy instruments has precluded their use in the empirical panel literature. However, Lora (1997) has recently produced a Structural Policy Index for most Latin American countries over the last decade of our sample period that may contain important additional information to incorporate into the statistical methodology. To achieve a more comprehensive coverage of reform policies, therefore, we augmented our set of “reform” variables to include the Structural Policy Index.

Let \( l_{it} \) denote the index for country \( i \) in subperiod \( t \) over the entire panel. When the underlying index is available, this variable can be derived by computing the five-year averages of the underlying annual index for 1986–90 and 1991–95. Consider the modified dynamic regression equation:

\[
g_{it} = s_i + w_t + ar_{it} + a'r'_{it} + bc_{it} + fl_{it} + u_{it} \tag{3}
\]

This specification is an improvement over our previous one to the extent that the coefficient is positive and statistically significant. The growth contribution of the macroeconomic policy package in this specification is \( ar_{it} + a'r'_{it} + fl_{it} \). Unfortunately, this index is available only for Latin America and only for the period 1985–95. Thus, it only covers a small fraction of the panel. To use the index, therefore, statistical assumptions are needed to complete the missing information. The following two statistical assumptions about how to estimate \( l_{it} \) when it is not available give rise to simple estimating equations:

**Full coordination (Assumption A)**

To the extent that reforms are interlinked and coordinated, it may be that the four macroeconomic reform variables used above already capture growth effects that in reality arise from other reforms that have been omitted from our estimated regression, but that are included in the index.\(^{13}\) To capture the overall growth effects of reform, the relevant question is how much information the index contains that is not already captured by the explanatory variables previously included in the regression. Suppose that reforms are typically coordinated. The determination of the policy index can then be specified as:

\(^{13}\)If so, this is an argument against attributing specific contributions to individual policies based on the previous econometric results.
If the predictive equation $d_i + pr_i$ provides a good approximation of the value of the index (that is, if it accounts for almost all of the variation in the index), then the introduction of the index $l_i$ into the growth equation would make no difference for its overall fit, because the index is spanned by variables already used. However, when the index contains independent information, it may contribute to improving the goodness of fit. In the event, the specification in equation (4) turns out to explain about 70 percent of the total variation of the index where it is observed, according to the adjusted R-squared statistic. The estimates of the four parameters associated with the reform variables are positive and statistically significant, with at least 90 percent confidence. One interpretation of these results is that the structural reforms reflected in the index have tended to be coordinated with the macroeconomic reform variables included in $r$, but the index nevertheless contains independent information.

Our first approach to exploiting this information is to assume that when the index is not available, it can be closely approximated by the above-mentioned predictive function of the four reform variables (and arbitrary country-specific structural differences to control for cross-country structural differences). We refer to this as Assumption A. To implement it, let $e_i$ be the residual from equation (4), taking on the value $v_i$ when the index is available and 0 otherwise; $e_i$ can thus be interpreted as capturing components of reform that were uncorrelated with macroeconomic policy variables. Implicitly, we are assuming that there was no time variation in this dimension of reform where the index is not observed. Consider the following equation:

$$g_i = s_i + w_i + ar_i + a'r_i + bc_i + fe_i + u_i.$$  

Under the assumption just described, this equation is equivalent to (3). The growth contribution of the macroeconomic policy package is now what is directly obtained from all of the reform variables, that is, $ar_i + a'r_i + fe_i$.

The corresponding estimations are shown in column 1 of Table 2. The coefficient $f$ has a positive sign and is statistically significant (p-value of 8 percent). This new specification does not have much effect on the estimated coefficients of the other variables or on the overall fit of the growth regression, but it has the effect of raising expected Latin American growth during 1991–95, thus explaining more of the acceleration with respect to the previous five-year period than the equation that excludes the index. The Latin American 1991–95 dummy is

14 A dynamic version including both the error and its lag was attempted, but both coefficients were individually insignificant in the statistical sense.
Table 2. Explaining Annual Per Capita GDP Growth

<table>
<thead>
<tr>
<th>Explanatory variables</th>
<th>Including Structural Policy Index</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(3): Assumption A (Lags underneath)</td>
</tr>
<tr>
<td>Stabilization and structural reform</td>
<td></td>
</tr>
<tr>
<td>Lower public consumption</td>
<td>3.4 (0.8)</td>
</tr>
<tr>
<td></td>
<td>-1.1 (0.8)</td>
</tr>
<tr>
<td>Lower inflation</td>
<td>1.5 (0.6)</td>
</tr>
<tr>
<td></td>
<td>-0.2 (0.6)</td>
</tr>
<tr>
<td>Financial deepening</td>
<td>1.5 (0.6)</td>
</tr>
<tr>
<td></td>
<td>-0.2 (0.7)</td>
</tr>
<tr>
<td>Exchange rate unification</td>
<td>2.3 (0.5)</td>
</tr>
<tr>
<td></td>
<td>-0.4 (0.6)</td>
</tr>
<tr>
<td>Structural policy index (residual)</td>
<td>15.8 (9.0)</td>
</tr>
<tr>
<td>Structural policy index (change)</td>
<td>—</td>
</tr>
<tr>
<td>Control variables</td>
<td></td>
</tr>
<tr>
<td>Initial GDP</td>
<td>-3.0 (0.7)</td>
</tr>
<tr>
<td>Education</td>
<td>0.56 (0.39)</td>
</tr>
<tr>
<td>Terms of trade</td>
<td>6.2 (2.5)</td>
</tr>
<tr>
<td>Worldwide cycle</td>
<td></td>
</tr>
<tr>
<td>1966-70</td>
<td>0</td>
</tr>
<tr>
<td>1971-75</td>
<td>0.49 (0.4)</td>
</tr>
<tr>
<td>1976-80</td>
<td>0.19 (0.4)</td>
</tr>
<tr>
<td>1981-85</td>
<td>-1.52 (0.5)</td>
</tr>
<tr>
<td>1986-90</td>
<td>-0.83 (0.6)</td>
</tr>
<tr>
<td>1991-95</td>
<td>-1.62 (0.7)</td>
</tr>
<tr>
<td>Number of observations</td>
<td>376</td>
</tr>
<tr>
<td>Adjusted R²</td>
<td>0.60</td>
</tr>
<tr>
<td>DW statistic</td>
<td>1.99</td>
</tr>
<tr>
<td>Latin America 1991–95</td>
<td></td>
</tr>
<tr>
<td>Average residual</td>
<td>0.28</td>
</tr>
<tr>
<td>Dummy</td>
<td>0.59 (0.62)</td>
</tr>
</tbody>
</table>

Note: Standard error estimates are given in parentheses.

Statistically insignificant at conventional levels (p-value of 38 percent). Still, the growth acceleration in the region on average continues to exceed expectations, though now by only 0.19 percentage points.
No coordination (Assumption B)

The alternative statistical assumption (Assumption B) is that instead of perfect coordination with the four reform variables, no coordination is present—that is, the Structural Reform Index contains information that is independent of the included variables. In a sense, then, this assumption is the opposite of the previous one. In this case, the policy index would be specified as:

\[ I_t = d_t + v_t. \]  

This equation explains about 50 percent of the total variation of the index according to the adjusted R-squared statistic. If the corresponding predictive equation is taken as a good approximation of the value of the index, then the introduction of the index in equation (3), would make no difference for the overall fit, since the country dummies would already contain the relevant information. The actual observed index, however, may contribute to explaining growth when it is available. To use it, we again need to make assumptions about its value when it is not observed. We make the same assumption that we did previously: when the index is not observable, it remains constant over time. Thus, let \( e_t \) be the relevant residual. It is assumed to take on the value \( v_t \) when the index is available, and 0 otherwise. An important difference in this case, though, is that under the assumed predictive function, it makes sense to specify Latin America’s vector of dummies \( d_t \) as the observed, out-of-sample, pre-reform 1985 value of the index. Therefore, \( v_t = I_t - I_{1985} \). This means that for countries in Latin America, the constant value is taken to be that observed in 1985, while for other countries, the constant index level is arbitrary.

The corresponding estimations are shown in column 2 of Table 2.\(^{15}\) The coefficient \( f \) again has a positive sign and is statistically significant (\( p \)-value of 4 percent). This alternative totally eliminates the excess of observed Latin American growth. Even under this extreme case, however, the growth response to reform remains not disappointing (it is almost exactly as expected, at 0.02). The Latin American 1991–95 dummy turns out to be negative, but it is statistically insignificant (\( p \)-value of 81 percent).

We conclude that the Structural Policy Index contains useful information. Thus, the overall conclusion under the growth impetus approach is that, judging by international and historical standards, the growth response of recent reform in Latin America—that is, its marginal effect—was adequate. Since the validity of the inferences also depends on untestable statistical assumptions, however, the results must be interpreted with caution.

\(^{15}\)A dynamic version including both the error and its lag reproduced the partial dynamic offset found for the other macroeconomic reform variables. For comparability, the statistically insignificant lagged variable was dropped.
II. Long-Run Growth Effects and Growth Acceleration

Determining whether reform measures had the growth effects that would have been predictable on the basis of international evidence is only the first step, however, in assessing the adequacy of the growth effects of reform. We now turn to an examination of the magnitude of the actual growth impact of the reforms in two different ways. First, we apply the new estimates of growth determinants derived in the last section to quantify the long-run contribution of stabilization and structural reform in the 1990s to per capita growth in Latin American countries, that is, its contribution to sustainable growth acceleration. Second, we measure the growth acceleration induced by reforms relative to the previous period.

Before proceeding, it may be worth noting that, in general, the growth contribution of the overall macroeconomic policy stance is what is directly obtained from the reform variables—that is, \( ar_{it} + a'r'_{it} + fe_{it} \), plus the unknown contribution \( fd_{it} \), which, as explained previously, is absorbed by the country dummies. This last term is constant and therefore irrelevant for assessing the contribution of a macroeconomic policy reform package in a given country (as in Tables 3-5 below). However, it becomes relevant for the decomposition of cross-regional growth gaps into the portions contributed by policy and by other factors (as in Table 6 below). Thus, assumptions regarding this last term in non—Latin American countries are needed to decompose cross-regional growth gaps under this scenario. Fortunately, the evidence suggests that a particularly simple assumption—that the unobserved aggregate \( d_{it} \) tends to be equal across regions—may well be realistic. Specifically, the Latin American country dummies estimated from the growth regression turn out to be uncorrelated with the out-of-sample, pre-reform values of the Structural Policy Index for individual countries in 1985. This finding suggests that systematic differences in the unobservable policy index across countries are fully absorbed by differences in the four measured macroeconomic reform variables and, therefore, the expected value of \( d_{it} \) can reasonably be taken to be constant across countries.

The Long-Run Growth Effects of Reform

The long-run growth effects of the reforms can be derived by multiplying the sum of the current and lagged coefficients of each of the reform variables by the change in that variable from 1986–90 to 1991–95 and aggregating over all of the reform variables. Table 3 shows these estimations as additional percentage points in annual growth on a country-by-country basis, as well as their sensitivity to the aggregation method and the statistical assumptions used for incorporating the Structural Policy Index. In almost all countries it is estimated that stabilization and

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16The long-run growth effects and decompositions of growth gaps presented in this and the following section are based on a model that incorporates the Structural Policy Index under the coordination assumption (Assumption A), under which the four macroeconomic reform variables in the model already capture much of the index. This case appears to be intermediate between the exclusion of the index and its incorporation under the no-coordination assumption. The conclusions in both sections, however, are robust to the alternatives.
structural reform made a substantial contribution to long-run growth as measured by all of the estimation methods. The preferred specification is the one in which the Structural Policy Index was introduced under Assumption A, in which case the typical country, as measured by the simple average, experienced a sustainable growth increase estimated at about 1.6 percentage points a year (with a standard deviation of 0.3). To the extent that reforms do not affect population growth, the best estimate of growth effects in the region as a whole is obtained through the GDP-weighted average of country growth effects, as opposed to a population-weighted average. By this method, the contribution of stabilization and structural reform to aggregate long-run growth is estimated at about 1.8 percentage points a year (with a standard deviation of 0.4). Other methods yield roughly similar results.

Therefore, the conclusion that recent stabilization and structural reform made a significant contribution to sustainable growth appears to apply to almost all individual countries and to be robust to these alternative methodologies. In the long run, if reforms are sustained, some of the current gains achieved in the first half of
### Table 4. Latin America: Decomposition of Changes in Per Capita Growth, 1991-95 Compared with 1986-90

<table>
<thead>
<tr>
<th>Stabilization and structural reform</th>
<th>Typical Country</th>
<th>Regional Aggregate</th>
</tr>
</thead>
<tbody>
<tr>
<td>Current reform</td>
<td>1.88</td>
<td>2.21</td>
</tr>
<tr>
<td>Past reform</td>
<td>1.95</td>
<td>2.00</td>
</tr>
<tr>
<td>Current reform</td>
<td>1.95</td>
<td>2.00</td>
</tr>
<tr>
<td>Past reform</td>
<td>-0.07</td>
<td>0.21</td>
</tr>
<tr>
<td>Control variables</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Income</td>
<td></td>
<td>0.04</td>
</tr>
<tr>
<td>Education</td>
<td>0.07</td>
<td>0.18</td>
</tr>
<tr>
<td>Terms of trade</td>
<td>0.09</td>
<td>-0.03</td>
</tr>
<tr>
<td>Other factors</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Unmeasured external factors</td>
<td>-0.94</td>
<td>0.92</td>
</tr>
<tr>
<td>Transitory differences</td>
<td>0.15</td>
<td>0.32</td>
</tr>
<tr>
<td>Aggregation</td>
<td>0.00</td>
<td>-0.15</td>
</tr>
<tr>
<td>Total growth rate increase</td>
<td>1.05</td>
<td>1.48</td>
</tr>
</tbody>
</table>

### Table 5. Latin America: Decomposition of Per Capita Growth Reduction, 1991-95 Compared with 1976-80

| Argentina                          | 1.00            | -1.60            | 3.76 | -3.16 |
| Bolivia                            | 1.58            | -2.81            | 2.85 | -1.62 |
| Brazil                             | -1.93           | -1.58            | 1.93 | 1.58  |
| Chile                              | 1.67            | -1.69            | 0.06 | -0.04 |
| Colombia                           | -0.04           | -2.21            | 1.42 | 0.83  |
| Costa Rica                         | 0.91            | -1.92            | 0.96 | 0.05  |
| Ecuador                            | 2.45            | -2.44            | -2.26| 2.25  |
| El Salvador                        | 2.86            | -2.23            | 4.63 | -5.26 |
| Guatemala                          | 1.46            | -2.00            | -2.02| 2.57  |
| Honduras                           | 0.90            | -2.01            | -1.83| 2.94  |
| Haiti                              | -1.41           | -2.04            | -7.45| 10.90 |
| Jamaica                            | 3.51            | -1.62            | 3.79 | -5.68 |
| Mexico                             | 0.83            | -1.86            | -3.61| 4.64  |
| Paraguay                           | 0.24            | -1.67            | -5.61| 7.05  |
| Peru                               | 1.32            | -2.06            | 4.49 | -3.75 |
| Uruguay                            | 0.91            | -1.60            | 0.16 | 0.53  |
| Venezuela                          | 0.54            | -2.79            | 4.01 | -1.76 |
| Typical country                    | 0.99            | -2.01            | 0.31 | 0.71  |
the 1990s will be lost owing to negative policy dynamics. If the reform level during 1996–2000 equaled that of 1991–95, there would have been an estimated aggregate dynamic per capita growth loss of 0.2 percentage points.

Growth Acceleration

Another way to assess the growth effects of macroeconomic reform is to determine the growth increase they delivered relative to the pre-reform period. The difference between average country growth performance during the reform period 1991–95, denoted by $g_t$, and average growth performance in the same country in the previous period, denoted by $g_0$, can be expressed as follows:

$$g_t - g_0 = \left[ r_t^* - r_0^* \right] + \left[ c_t^* - c_0^* \right] + \left[ w_t^* - w_0^* \right] + \left[ \left( g_t - g_t^* \right) - \left( g_0 - g_0^* \right) \right], \quad (7)$$

where $g_t^* = [r_t^* + c_t^* + w_t^*]$ and $g_0^* = [r_0^* + c_0^* + w_0^*]$ are the fitted values of the “preferred” growth equations from the introduction, using average values of explanatory variables in the last five-year period and the previous period, respectively. The first two terms on the right-hand side of this equation capture the “explained” portion of any growth change in a Latin American country. In other words, they measure the extent to which growth changes in Latin American countries can be explained by the variables that have been influenced by macroeconomic reform (denoted by $r$) and systematic differences in the set of “control” variables (denoted by $c$). The next term captures differences in the effects of
external factors (denoted by \( w \), which can be interpreted as an exogenous temporary shock). This term captures the extent to which unaccounted international exogenous factors related to growth (such as the debt crisis or productivity of new inventions) account for differences in growth performance in Latin America. The last term corresponds to the difference between Latin American growth residuals in both periods—that is, the random (and transitory) portion of the difference in growth performance between the two periods. This final term captures the extent to which the recent growth response has been disappointing, in the sense of falling short of what could reasonably have been expected on the basis of international experience.

The above decomposition of the change in growth holds for each individual country and, consequently, for the typical country as measured by the simple average. The typical country experienced an average growth acceleration in the period of slightly above 1 percentage point a year, but the contribution of recent macroeconomic reform far exceeds this mark (see Table 4). In fact, the impact effect of reform (current reform) is estimated at about 2 percentage points; this is partially obscured by external exogenous factors, mainly a strong negative effect of external factors accounting for more than 1 percentage point of growth reduction, while other explanatory variables played a relatively minor role and largely offset each other.

This conclusion also holds for the region as a whole. Although the above decomposition does not hold in the aggregate, to the extent that population growth rates are not affected by the explanatory variables considered in this growth model, GDP-weighted averages exactly identify the aggregate growth contributions of these variables in all the growth gap decompositions analyzed in this study, and the remaining statistical discrepancy is attributable to demographic factors. The aggregate growth acceleration owing to stabilization and structural reform during 1991–95 is therefore estimated at 2 percentage points a year. Had this reform been deeper, its impact growth effect would have been correspondingly larger when multiplied by the estimated marginal growth effects. For example, if reforms had attained the levels observed in the OECD or the East Asian miracle region, the resulting aggregate growth acceleration impact would have been 4.24 and 5.14 percentage points a year, respectively. The conclusion is that while significant, the impact growth effect of reform fell short of half of its potential, judged by these standards.

III. Accounting for Reform Effects: Growth Shortfalls

The question to be posed in this section is the following: Considering some absolute standard of growth performance—we take it to be average Latin American per capita growth rates during the 1970s, as well as both the average East Asian and OECD growth levels during the Latin American reform period—how can the shortfall between such a standard and the actual growth experience be explained in terms of the growth determinants we have identified? We take up each of the alternative standards of comparison in turn and show the results in Tables 5 and 6.
Contribution of Reform to Growth Acceleration

Most Latin American countries grew faster in the 1970s than in the recent reform period 1991–95; in fact, the typical country, obtained as a simple average of all countries, experienced a growth rate about 0.70 percentage points a year lower. However, this growth decline is not evidence of reform failure. When this growth shortfall between the high growth period 1976–80 prior to the debt crisis and the reform period 1991–95 is explained along the lines of the formula in the previous section, as is done in Table 5, it becomes apparent that the macitromomic reforms contributed to a sizable increase in growth in almost all Latin American countries. In fact, in the typical country, relative to the situation prevailing during 1976–80, better macroeconomic policy, as measured by the first term of the decomposition formula, contributed to a growth improvement of about 1 percentage point a year. Such progress was more than offset by a severe deterioration of the external growth environment in all countries, as measured by the effects of external factors (contributing an estimated 1.81 percentage points a year of growth reduction) and the international terms of trade of each country, which resulted in a decline of about 2 percentage points a year for the typical country. These two factors, macroeconomic reform and external environment, explain growth performance well across the region. In fact, while other factors may have been important for explaining performance in each individual country, they made a modest contribution of about 0.3 percentage points in the typical country.

Reform and Interregional Growth Gaps

A key advantage of panels including extraregional countries is that they permit us to employ an alternative standard of comparison, relying on cross-regional comparative analyses, to supplement the country-by-country time-series dimension. In this case, the unit of analysis cannot be the country. Instead, we compare regional aggregates. The decomposition of the growth shortfall between aggregate growth performance in Latin America and other regions during the reform period 1991–95 can also illustrate the role of recent reform and the remaining reform agenda needed to close the growth gap.

Suppose for concreteness that the East Asia miracle region is taken as a benchmark. The difference between aggregate East Asian growth performance during this period, denoted by $g_{EA}$, and aggregate Latin American growth performance, $g_{LA}$, can be expressed as follows:

$$g_{EA} - g_{LA}$$
where $g_{EA}^*$ = $[r_{EA}^* + c_{EA}^* + s_{EA}^*]$ and $g_{LA}^*$ = $[r_{LA}^* + c_{LA}^* + s_{LA}^*]$ are the fitted values of the preferred growth equations from the introduction using GDP-weighted average values of explanatory variables for East Asian and Latin American countries, respectively, and $d$ is the demographic statistical discrepancy discussed in the previous section. Again, the first two terms in the right-hand side of this equation capture the explained portion of any Latin American aggregate per capita growth shortfall. In other words, they measure the extent to which Latin America's growth performance can be explained by the variables that have been influenced by macroeconomic reform (denoted by $r$) and systematic differences in the set of control variables (denoted by $c$). The next term captures differences in regional averages of country dummies (the structural differences). We interpret this term as capturing the extent to which structural features of economies in the two regions—that is, features that have been constant for some time and that are related to growth performance—account for differences in aggregate growth performance during Latin America's reform period. The next to the last term corresponds to the difference between East Asian and Latin American aggregate growth residuals—that is, the random (and transitory) portion of the difference in growth performance between the two regions during 1991–95. This term would also capture the extent to which Latin America's recent growth experience has been disappointing, in the sense of falling short of what could reasonably have been expected on the basis of international experience.

For the Latin American region as a whole, most of the enormous growth gap with East Asia of almost 5 percentage points is explained by incomplete reform. According to these estimates, if Latin America attained East Asian values for the reform variables and they remained constant, the per capita growth gap would shrink in the long run by $2\frac{1}{2}$ percentage points. The educational deficit in Latin America is responsible for about $\frac{1}{2}$ of 1 percentage point of the growth gap, most of which is offset by the region's relative poverty, which, everything else being equal, facilitates growth (conditional convergence). Structural differences also contribute to the growth gap, with their importance comparable in magnitude to that of the educational deficit. About one third of the growth gap remains unexplained by the factors we have identified.

In contrast to East Asia, per capita growth in the OECD countries during the 1990s was slower than in Latin America. Given the enormous difference in income per capita, reflected in the growth contribution of GDP per capita in Table 6, this is not surprising. What is perhaps surprising is that differences in reform make a smaller contribution in favor of the OECD than in the East Asian case. At the same time, there are very significant structural differences in favor of the OECD countries (accounting for almost 4 percentage points of the growth gap); these are more important than the contribution of stabilization and structural reform, estimated at...
less than 2 percentage points, which probably reflects the large differences in the stages of development of the two regions. Coupled with this observation, the significant East Asian growth residual suggests that part of the contribution assigned to transitory factors in East Asia may be permanent in nature and that East Asia may be following the steps of the OECD countries in this regard.

The overall analysis of both decompositions suggests that, for the Latin American region as a whole, there are significant growth gains to be achieved if reforms, including in the area of educational policy, are deepened. Attaining East Asian levels for the reform variables as well as for educational achievement would substantially close the Latin America growth gap with the East Asian region and possibly set the stage for other structural transformations as development is advanced and the gains from the first generation of reforms are completed.

IV. Conclusions

To summarize our findings, it is useful to consider alternative hypotheses that could be offered to explain Latin America's recent growth performance. The simplest would be, of course, that the fundamental thrust of the reforms has been misguided if the objective was to improve growth performance. We reject this hypothesis. Not only does the weight of the evidence in the professional literature, as well as our own results, support the view that the market-friendly reforms implemented in the region to date should have been growth enhancing, but we found no empirical evidence for the view that Latin America is "different" in this regard—that is, we have found no evidence that the growth response to the reform variables has been systematically different in Latin America than elsewhere. Moreover, the growth impetus associated with the reforms has been substantial: the estimated long-run growth effect of the 1990s reform is large for most countries in the region and amounts to almost 2 percentage points of additional annual sustainable growth in the aggregate, enough to double the real income expected in 40 years (see Table 3).

A second possibility that we were able to discard is that there is a Latin American growth "puzzle," in the sense that unidentified region-specific factors depressed growth in Latin America during the 1990s, offsetting the large positive growth impetus of the reforms just described. In fact, a time- and region-specific dummy for the reform period in Latin America was statistically insignificant when added to the panel growth equations.

In short, even after extending the sample, broadening the set of reform indicators, and taking into account possible dynamic effects, our findings are consistent with those of ELM, in the sense that we found no evidence of disappointing growth performance when disappointment is measured either in terms of the marginal effects of the reforms or in terms of the overall growth impetus that they imparted to Latin American countries during the reform period.

Why, then, did Latin America not experience a more pronounced acceleration of growth during 1991-95 leading to more satisfactory levels of growth? The answer appears to lie in a combination of factors:
- The reforms were implemented in a relatively unfavorable external environment. The effect of implementing the reforms during 1991–95, instead of in the previous five-year period, was to associate them with an international context that by itself reduced the average growth rates of the reforming countries by about 1 percentage point.

- For growth to have accelerated more than it did would have required more intensive reforms along the lines already implemented. We found evidence that there is indeed room to move further in this direction, in the sense that Latin America has not yet reached the levels of performance achieved in faster-growing regions.

- Our results would also support a case for more extensive structural and institutional reforms—that is, for broadening the scope of reform—because pushing macroeconomic reforms to the levels of performance achieved in the faster-growing regions would be insufficient for Latin America to close the growth gap. Our results suggest that only about half of the annual growth gap of about 5 percent between Latin America and East Asia during the reform period can be closed by doing more of the same—that is, intensifying the reform effort along the lines already undertaken. This conclusion emerges with even greater force in comparison to the OECD, where structural differences account for an even larger share of the current difference in growth performance relative to Latin America. This remaining gap suggests that the scope of reform in Latin America will need to be broadened. Improvements in macroeconomic management are simply not sufficient for Latin America to achieve long-run growth rates comparable to those achieved in East Asia.

The final result of our study is, therefore, that while much has been painfully achieved in Latin America, and while the reforms that have been implemented have indeed delivered the boost in growth that they could have been expected to provide on the basis of international evidence, reaching much higher long-term growth rates in the region—beyond historical growth rates and approaching the rates of high growth regions—will require both an intensification of reform along the dimensions already implemented and a broadening of reform to incorporate changes in structural characteristics of Latin American economies that are still inhibiting growth in the region. Our results in this paper do not permit us to go further in identifying such characteristics, but we have been able to document their importance indirectly. A key item on the research docket for the region, therefore, should be to identify desirable directions in which to extend the reform agenda, as well as ways to make further progress in consolidating and intensifying the reform efforts that are currently under way.

APPENDIX

The panel consists of the following 69 countries during 1961–95:

Latin America (18): Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Ecuador, El Salvador, Guatemala, Haiti, Honduras, Jamaica, Mexico, Paraguay, Peru, Trinidad and Tobago, Uruguay, and Venezuela.
REFORM AND GROWTH IN LATIN AMERICA

OECD (17): Australia, Austria, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, Netherlands, Norway, Spain, Sweden, Switzerland, United Kingdom, and United States.


East Asian miracle countries (5): Indonesia, Korea, Malaysia, Singapore, and Thailand.

Others (11): Bangladesh, Cyprus, Greece, India, Israel, Jordan, Pakistan, Philippines, Portugal, Sri Lanka, and Turkey.

The period was divided into seven five-year subperiods: 1961–65, 1966–70, 1971–75, 1976–80, 1981–85, 1986–90, and 1991–95. Five-year simple averages of the available underlying yearly information were used. The resulting information panel was unbalanced because of data limitations for some countries. Of a total of 482 possible observations, 37 were not available.

Except when noted, the data sources used are Inter-American Development Bank, World Bank, and IMF official information. The basic data were the real growth rate of per capita GDP; real consumption as a proportion of real GDP; openness measured as real imports plus exports as a proportion of real GDP; inflation rate based on monthly CPI; financial deepening measured as the ratio of real M2 (deflated by year-end CPI) as a proportion of real GDP; real per capita GDP at the beginning of each period; average years of secondary schooling in the total population of 15+ years at the beginning of each period (Barro-Lee data set); terms of trade growth rate; black market premium (for 1961–84 from Wood, 1988, and for 1985–95 from World Currency Yearbook, 1996).

The following variables were entered with a logarithmic transformation: openness ratio, government consumption ratio, inflation (as 100+ inflation rate in percent), financial depth ratio, initial GDP per capita, and black market premium (as 1+ premium).

Data on the Structural Reform Index (Latin America, 1985–95) are from Lora (1997). The volatility of inflation was measured as the standard deviation of annual inflation rates. The volatility of terms of trade was measured as the standard deviation of annual terms of trade growth rates. The inequality of income distribution is the income of the richest quintile divided by the income of the poorest two quintiles.

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Risk, Resources, and Education: Public Versus Private Financing of Higher Education

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The paper develops a public education scheme that takes aspects of uncertainty in private educational investments explicitly into account. The social merits of public education schemes are related to the lack of markets in which students can insure against educational risks. A case is made for tuition fees that depend on expected returns of investments in education. The consideration of uncertainty provides a neglected link between educational choice, resource endowment, and productivity growth, which may serve to redefine the public role of education financing. [JEL H52, I22, D81]

In virtually all developed countries the government is engaged in higher education. A commonly used argument is that private markets are unable to provide higher education up to an efficient amount. Positive externalities and capital market imperfections figure most prominently in the list of arguments claiming that investments in higher education regulated by private markets are too low.

Positive externalities result if the returns on higher education are not fully priced in private markets. Undoubtedly, university graduates provide the society with valuable services. On the other hand, they earn comparatively high incomes, which may have internalized the economic surplus of university graduates. But, even if some work by university graduates causes external effects beyond the

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respective salary differential, a Pigou-internalization by subsidizing all university students, which is common practice, will hardly be successful.¹

The imperfect capital markets argument relies on the notion that human capital acquired by education is nontradable. Human capital cannot be lent ahead and thus cannot be used as a security for student loans. This can lead to an inefficiently small demand for higher education by those who are not able to finance a cost-intensive education due to liquidity constraints.² Capital market imperfections, however, do not necessarily suggest educational subsidies. Publicly provided student loans should suffice to redress a too-small demand for higher education due to capital market imperfections.³

One should bear in mind, though, that individual returns to education are generally uncertain. The individual can neither be sure about finishing his education successfully nor about his future returns after a successful examination. In fact, educational returns display a very high variance as many students do not graduate, income differences between graduates are large, and even the risk to become unemployed exists.⁴ Publicly provided student loans, which have to be paid back irrespective of educational success, generally do not change the nature of individual educational risk. Yet, the risk an individual faces with an investment in education can be expected to constitute a significant disincentive to invest in education, as individuals are unable to insure against these risks on private markets. Therefore, it can be assumed that risk-averse individuals do not adjust expected marginal returns on educational investments to marginal costs. Instead, they underinvest in education.

What matters for society as a whole are average returns of all university graduates. Society should invest in higher education until average marginal returns of educational investments equal their marginal costs. A publicly provided educational program with a success-dependent cost participation of university students may contribute to an individual realization of this rule.

In this paper we develop a public education program that explicitly takes uncertainty of private education into account. The policy objective is charac-

¹See, for example, Lüdeke (1996) and Heckman and Klenow (1998) for a discussion on positive externalities of higher education. It should be noted that modern growth theory à la Romer (1986) and Lucas (1988) is based on positive externalities from education to endogenize the economy’s rate of growth. See Wigger (2001) for internalization policies of such externalities.

²Most of the literature concerned with the relationship between income distribution and growth assumes some kind of capital market imperfection. See, for example, Galor and Zeira (1993), Perotti (1993), and Barham and others (1995), as well as the surveys by Perotti (1994) and Benabou (1996).

³See Lou (1987) for a more in-depth survey of the normative justifications for subsidizing higher education. Trostel (1993, 1996) provides a further efficiency argument for educational subsidies relying on the notion that income taxation distorts the investment decision on education which, in turn, can be corrected by an education subsidy. In contrast to the arguments above, however, this constitutes a second-best argument for subsidizing education. It should be noted that there is an alternative strand of the economics of education literature which emphasizes a public choice perspective of public education financing rather than stressing efficiency arguments. This literature includes Peltzman (1973), Johnson (1984), and Fernandez and Rogerson (1995).

⁴Already Becker (1964, p. 104) has pointed out that educational returns are characterized by a very high coefficient of variation. This observation is supported by more recent studies such as, for instance, Miller and Volker (1993). For a theoretical approach to educational risk see Levhari and Weiss (1974).
terized by a maximization of a tax dividend of public education financing. We first derive a first-best rule of public education financing which implies full insurance against educational risk. We then consider how adverse selection with respect to educational abilities and moral hazard with respect to educational and work effort affect the shape of the educational program. Our model provides a case for *success-dependent* tuition fees as they already exist in some countries.5

The consideration of educational risk also gives an interesting insight into the interrelation between individual educational choice, initial resource endowment, and productivity growth. We show that the extent of private educational financing as well as the social surplus of a public funding of education depend on both individual resource endowments and productivity growth. The necessity for a public scheme becomes less compelling when better endowed individuals are increasingly willing to accept educational risks. Productivity growth may also lead to an increase in private educational investments. This effect, however, tends to be much weaker than the impact of an increase in initial endowments.

**I. The Basic Framework**

This paper considers individuals facing some amount of risk when deciding on the level of educational investment. Such an investment has two possible outcomes: the education undertaken may either be successful or fail, where success and failure are measured in terms of disposable income.6 Let yS and yF be disposable income in case of success and in case of failure, and let \( \pi \) denote the probability that the educational investment fails. This probability can be understood as a measure of the lack of individual talent and purposefulness. Expected utility may then be written as:

\[
Eu = \pi u(y^F) + (1 - \pi)u(y^S).
\]

The von Neumann-Morgenstern function \( u : \mathbb{R}_+ \to \mathbb{R} \) is assumed to be twice continuously differentiable. It is assumed that the individuals are risk-averse and anxious to realize positive consumption in any state. This implies \( u' > 0, u'' < 0, \) and \( u'(0) = \infty. \)

If the educational investment is not successful, individual disposable income is given by:

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5An example is the Higher Education Contribution Scheme, which was introduced in Australia in 1989. See Chapman (1997) for an economic analysis of this scheme.

6The model could easily be extended by considering more than two outcomes of educational investments. This would not, however, affect the main results.
where $e$ marks the educational investment and $x$ is the labor income of an unskilled individual. Hence, an individual failing in education has less income at its disposal than an individual who has not undertaken an education at all. If education is successful, disposable income is given by:

$$y^s = z(e) - e,$$

where the function $z : \mathbb{R}_+ \to \mathbb{R}_+$ indicates the return on an educational investment $e$. It is assumed that $z(0) = x$, $z' > 0$, $z'(0) = \infty$, and $z'' < 0$, implying that marginal returns on education are positive but decreasing.

II. Education Financing

Private Education Financing

If education is financed privately an individual will choose an educational investment $e$ such that expected utility takes on a maximum:

$$\tilde{e} = \arg\max \left\{ \pi u(x - e) + (1 - \pi) u(z(e) - e) : e \geq 0 \right\}.$$

The first-order condition for the ex ante optimal investment $\tilde{e}$ may be expressed as:

$$(1 - \pi) z'(\tilde{e}) = 1 + \pi \frac{u'(\tilde{y}^F) - u'(\tilde{y}^S)}{u'(\tilde{y}^S)} > 1,$$

where $\tilde{y}^F$ and $\tilde{y}^S$ denote disposable income in case of failure and success, respectively, if an educational investment $\tilde{e}$ was undertaken. Hence, the individuals do not equate expected marginal returns on education ($= (1 - \pi) z'(e)$) and marginal costs ($=1$), but due to risk considerations, invest a lower amount in education.

Before we can proceed to public education funding, we need a measure for the value individuals assign to privately financed education. It is convenient to employ the concept of the certainty equivalent. The certainty equivalent is that amount of income, $\tilde{y}$, which, if received for certain, the individual regards as just as good as
the expected income when it undertakes an educational investment $\bar{e}$. Thus, $\hat{y}$ is implicitly defined by:

$$u(\hat{y}) = \pi u(y^F) + (1 - \pi) u(y^S).$$

Obviously, a public education scheme has to offer a utility level of at least $\hat{u} = u(\hat{y})$. Otherwise, individuals will finance the education privately.

**Public Education Financing**

Suppose that the government encourages individuals by means of a public program to invest more in their education than they would do otherwise. This increases the average return on educational investments if there is a sufficiently large number of individuals who aim at a higher education and whose educational risks are distributed independently. It is sometimes argued that a public program of higher education financing will also be advantageous for that part of the population not participating in higher education, if the additional tax revenues exceed the costs of the scheme. The social (tax) return on higher education is called the tax dividend of a public education scheme.7

Following the notion of the tax dividend we assume that the policymaker's objective is to maximize the net social return on educational investments, which is defined as the sum of individual returns on educational investments minus the cost of education and minus the disposable incomes of the individuals having undertaken an education. The public program is an educational package that is a triplet consisting of an educational investment $e$, a disposable income in case of failure $y^F$, and a disposable income in case of success $y^S$.

Let the number of individuals being able to participate in higher education be given and normalized to 1. The net social return on educational investments is then given by:

$$T = \pi (x - y^F) + (1 - \pi) [z(e) - y^S] - e.$$

In what follows, we first determine the first-best public educational program. Afterward, we consider extensions of the model to analyze how adverse selection with respect to educational abilities and moral hazard with respect to educational and work effort affect the public program.

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7See, for example, Barr (1993, chap. 13).
First-best program of public education financing

Even in the basic framework the policymaker does not have complete freedom in designing the educational program. Since the individuals can always undertake a privately financed education, the public program must offer individuals at least as much utility as the private alternative. An individual will participate in the public education scheme only if:

\[ \pi u(y^F) + (1 - \pi) u(y^S) \geq \hat{u}; \]

that is, if the program offers at least the utility obtained by privately financed education. This constraint ensures voluntary participation in the public program. The optimal public program, denoted by PP1, can then be formulated as:

\[
\max_{\{e, y^F, y^S\}} T = \pi(x - y^F) + (1 - \pi)[z(e) - y^S] - e,
\]

subject to the voluntary participation constraint. The solution to PP1 is implicitly determined by:

\[ y^F = y^S \text{ and } (1 - \pi)z'(e) = 1. \]

Thus, the optimal program of education financing is characterized by a full public takeover of all educational risk and by the condition that the marginal expected return of educational investment equals its marginal costs.

Public education financing with different types

So far we have assumed that all individuals are identical. This is not a realistic assumption, however, as individuals normally differ with respect to their educational abilities. Therefore, we extend the model to allow for two types of individuals with differing educational risks.

Let the probabilities of educational failure of type 1 and type 2 individuals be given by \( \pi_1 \) and \( \pi_2 \) and assume that \( \pi_1 > \pi_2 \), that is, type 1 individuals face a lower chance of success than type 2 individuals. If the government could perfectly observe individual types, it would implement educational packages for each type as determined in the previous section. A reasonable assumption, however, is that individual types are not publicly observable. We assume that the policymaker knows that a proportion \( q_i \) of that part of the population, which

\[ \text{See Wigger and von Weizsäcker (1999) for technical details.} \]
is capable of participating in higher education, is of type $i = 1, 2$, with $q_1 + q_2 = 1$, but that he cannot observe each individual’s type separately. As a consequence, the first-best program is not feasible. Individuals of type 1 would simply masquerade as individuals of type 2 and choose the educational package intended for individuals of type 2 to achieve a higher utility level. Feasibility of a public financing program requires that no individual finds it worthwhile to choose an educational package designed for individuals of a different type. Standard arguments\(^9\) show that the optimal policy is constrained by the requirement that individuals of type 1 prefer to claim the educational package intended for themselves rather than the one intended for individuals of type 2. More precisely, an optimal program implies:

$$\pi_i u(y^F_i) + (1 - \pi_i) u(y^S_i) = \pi_i u(y^F_2) + (1 - \pi_i) u(y^S_2).$$

On the other hand, whether one or both of the voluntary participation constraints bind is not clear a priori, since individuals of different types would undertake different private educational investments. This implies that either both constraints bind or only the one in which individuals of type 2 prefer to undertake private educational investments binds. Hence, the optimal incentive compatible public scheme, denoted by PP II, may be written as:

$$\max_{\{e_S, y^F_i, e^S_i\}_{i=1,2}} T = \sum_{i=1,2} q_i \left[\pi_i \left(x_i - y^F_i\right) + (1 - \pi_i) \left(z(e_i) - y^S_i\right) - e_i\right],$$

subject to the incentive compatibility constraint and a voluntary participation constraint for each type. The optimal educational program in the presence of unobservable types, PP II, satisfies:

$$y^F_i = y^S_i, \quad y^F_2 < y^S_2, \quad \text{and} \quad (1 - \pi_i) z'(e_i) = 1, \quad i = 1, 2.$$

As in the first-best program, PP II is characterized by an equalization of marginal returns and marginal costs of educational investments for both types. Thus the information of the policymaker on educational types does not affect the optimal level of public educational investments. Furthermore, as in the first-best, the public program should take over the whole educational risk of individuals of type 1. Individuals of type 2, on the other hand, should face some of their educational risk. The intuition behind this result is as follows. Participation in educational risk is less attractive for type 1 than for type 2 individuals. If type 2 individuals face some educational risk this restrains type 1 individuals from masquerading and from choosing the educational package intended for type 2

\(^9\)See, for example, Kreps (1990, chap. 18).
individuals. The net social return on educational investments of PP II is less than that of PP I, even if both participation constraints are binding. If type 2 individuals are given a share in their educational risk, their average disposable income has to exceed their certainty equivalent. Otherwise, due to risk aversion, they would not participate in the public program.

Public education financing and moral hazard

The analysis so far has abstracted from incentive effects of a public takeover of educational risks on individual effort. However, public insurance of educational risks may affect incentives to be successful in two respects. First, it may undermine individual incentives to avoid educational failure. Second, it may give individuals an incentive not to reveal their success in order to reduce their tax liabilities after education. The first instance—ex ante moral hazard—may occur if the government cannot observe individual effort to be successful. The second instance—ex post moral hazard—may occur if the government cannot observe success directly but has to infer it from individual income.

Ex ante moral hazard

In this section we study how these types of moral hazard affect the shape of the public education program. For simplicity we return to the model with only one educational type. We start with ex ante moral hazard. Thus, we assume that the probability of educational success depends on individual effort. Individuals can choose between two effort levels $e_1$ and $e_2$, with $e_1 < e_2$. The probability that an individual is not successful is then determined by $\pi_i = \pi(e_i)$, $i = 1, 2$, with $\pi(e_1) > \pi(e_2)$. Educational effort $e_i$ is measured in currency units so that disposable income in case of failure and success become $y^F - e_i$ and $y^S - e_i$. If the government wants to enforce the effort level $e_2$ rather than $e_1$ (otherwise there would be no moral hazard problem), it has to consider the following incentive compatibility constraint:

$$\pi_2 u(y^F - e_2) + (1 - \pi_2) u(y^S - e_2) \geq \pi_1 u(y^F - e_1) + (1 - \pi_1) u(y^S - e_1).$$

The optimal public education program in the presence of ex ante moral hazard, denoted by PP III, then becomes:

$$\max_{\{e, y^F, y^S\}} \ T = \pi_2 (x - y^F) + (1 - \pi_2) [z(e) - y^S] - e,$$

subject to the incentive compatibility and a voluntary participation constraint. The solution to PP III is characterized by
Thus, consideration of ex ante moral hazard leads to individual risk participation in the public education program. This is because if the public took over all educational risk, individuals would have no incentive to choose the higher effort level to study towards success. In light of the analysis in the previous section this result implies that individuals of all educational types should bear some of the risk of educational failure. The optimal amount of public education, however, is not affected by ex ante moral hazard. It is still determined by an equalization of marginal returns and costs of educational investments, making the public program superior to its private alternative.

**Ex post moral hazard**

Consider next ex post moral hazard. Hence, assume that individual gross income does not only depend on educational success but also on work effort. Gross incomes are now assumed to be given by \( g^F = x l^F \) and \( g^S = z(e) l^S \), where \( l^F \) and \( l^S \) denote work effort in case of failure and success, respectively. Work effort \( l \) leads to a disutility measured in currency units by \( v(l) \), with \( v' > 0 \) and \( v'' > 0 \). If the government can only infer educational success from observed gross income, insurance in the form of redistribution between the successful and the unsuccessful is constrained by the fact that successful individuals can claim to be unsuccessful by choosing a lower level of work effort. The problem of the government then resembles that of designing an optimal income tax scheme. Employing the self-selection method, introduced by Stiglitz (1982) into the optimal taxation literature, a feasible public education program must satisfy:

\[
\max_{\{e, y^F, y^S, g^F, g^S\}} T = \pi(g^F - y^F) + (1 - \pi)(g^S - y^S) - e,
\]

saying that the successful do not find it worthwhile to claim to be unsuccessful. Thus, the optimal public education program in the presence of ex post moral hazard, denoted as PP IV, may be written as:

\[
y^F < y^S \text{ and } (1 - \pi) z'(e) = 1.
\]
subject to the incentive compatibility and a voluntary participation constraint. The solution to PP IV can be characterized by:\(^\text{10}\)

\[ y^F < y^S \text{ and } (1 - \pi) z'(e)/k > 1.\]

Clearly, PP IV cannot lead to equalization of income between the successful and the unsuccessful as this would not be consistent with incentive compatibility. Furthermore, PP IV does not lead to an equalization of marginal returns and costs of educational investment but implies a lower level of educational investment. To get an intuition of this result, assume, for a moment, that the government equalizes marginal costs and returns of educational investment. Then, reduce educational investment by one unit and simultaneously increase net income of the successful in order to satisfy the voluntary participation constraint. Such a policy exerts a negative effect on government revenues which is only of second order. However, it weakens the incentive compatibility constraint which, in turn, permits more insurance and, thus, higher taxes leading to a positive first-order effect on government revenues. Although educational investments are lower than under first-best, they are higher than the ones undertaken privately. This is because PP IV still provides some insurance against educational risk. Furthermore, the net social surplus of PP IV is positive as the policymaker could alternatively imitate private educational financing generating a net social surplus equal to zero.

### III. Initial Endowments and Productivity Growth

Hitherto, we have not considered the financial background of the individuals and its potential impact on private educational investments. It is conceivable that a privately undertaken educational investment depends on whether it has to be financed solely out of success-dependent income or whether the individual can fall back on additional financial resources. Furthermore, the absolute amount of disposable income may also affect private educational investments. This is important since, due to productivity growth, incomes of skilled and unskilled individuals have risen continuously during recent decades.

In this section we consider the impact of both initial resource endowments and productivity growth on private educational investment. We start by considering the effect of increasing initial resources. For simplicity we again assume that there is only one type of individual. Let \( a \) be an initial endowment. The state-dependent budget constraints are then given by \( y^F = a + x - e \) and \( y^S = a + z(e) - e \), where \( y^F \)

\(^{10}\)Further results are that the marginal tax rate imposed on the successful should be zero and the one imposed on the unsuccessful should be positive. These results are well known from the optimal taxation literature. For an analysis of insurance and optimal taxation see Varian (1980); for an analysis of public education and optimal taxation see Ulph (1977) and Hare and Ulph (1979).
and \(y^S\) now denote disposable resources rather than incomes in case of failure and success, respectively. The privately undertaken amount of educational investments is determined by:

\[
\bar{e} = \operatorname{argmax} \left\{ \pi u(a + x - e) + (1 - \pi) u(a + z(e) - e) : e \geq 0 \right\},
\]

providing a function \(\bar{e} = \bar{e}(a)\), which describes the relationship between the optimal amount of privately financed educational investments and the initial endowment. It satisfies:

\[
\frac{d\bar{e}}{da} \geq 0 \iff A(a + x - \bar{e}) \geq A(a + z(\bar{e}) - \bar{e}),
\]

where \(A(\cdot)\) is the Arrow-Pratt measure of absolute risk aversion. Since \(x < z(e)\) for all \(e > 0\), the amount of private educational investments is increasing in initial endowment if absolute risk aversion is decreasing. Considering the discussion of the previous section, decreasing absolute risk aversion implies that public education financing is the more important the less individuals can fall back on sufficient initial resources. In view of the standard argument that educational policies should facilitate access to higher education for the poor, this result gains in significance. In fact, it implies that the amount of educational investments undertaken by the poor differs most markedly from the socially optimal level. In this way the result provides a link between the distribution of initial resources and the net social return of public education financing.

When an increase in initial resource endowments \(a\) leads to higher private educational investments, one may ask whether a rise in disposable incomes due to productivity growth affects private investments in a similar way. If \(a\) is interpreted as an income component depending on productivity, one could conclude that private educational investments go up due to productivity growth, and a public program of education financing would cease to be necessary in the course of time. However, such an interpretation would, among other things, imply that relative educational costs decrease over time. If one observes that educational costs increase over time, such an analogy is less obvious. Suppose that educational costs are governed by the same trend as incomes. The state-dependent budget constraints may then be written as \(y^F = \tau \cdot (x - e)\) and \(y^S = \tau \cdot (z(e) - e)\), where the parameter \(\tau\) measures productivity at a certain point in time. The optimal amount of privately financed educational investments takes the form:

\[
\bar{e} = \operatorname{argmax} \left\{ \pi u(\tau(x - e)) + (1 - \pi) u(\tau(z(e) - e)) : e \geq 0 \right\},
\]

\[\text{This may be the case, for instance, if educational costs are linked to the employment of teachers.}\]
Proceeding in the same way as before we have:

$$\frac{d\bar{e}}{d\bar{e}} > 0 \iff R(\tau(s-\bar{e})) > R(\tau(\bar{e}-\bar{e})),$$

where $R(\cdot)$ is the Arrow-Pratt measure of relative risk aversion. An increase in productivity leads to higher privately financed educational investments if relative risk aversion is decreasing. The condition for productivity growth leading to higher private educational investments is thus much more restrictive than the condition for increasing resource endowments causing the same effect. Taking the results of the previous section into account, it has to be doubted that productivity growth may lead to privately financed educational investments at the socially optimal level.

**IV. Summary and Conclusion**

This paper has dealt with private educational investments under risk. It has derived an economic role for the state concerning education financing. In our framework the social merits of public education schemes are related to a nonexistence of markets in which students can insure against educational risks. Hence, a public program of education financing should contain elements of risk insurance. This could be realized, for instance, by success-dependent tuition fees. A complete public takeover of all educational risks cannot, however, be recommended. Individual participation in educational risk should be used to give individuals proper incentives to choose suitable courses of study and to study towards success. Cost-intensive courses should require more individual risk participation. This would ensure that only individuals with adequate educational abilities undertake costly education. Furthermore, individual risk participation induces individuals to make an effort to be successful and to exploit their human capital efficiently after education.

The role of public education financing also depends on the disposability of individual incomes and wealth. If absolute risk aversion is a decreasing function of income, the role of a public scheme becomes less important when educational risks are taken to a socially desirable extent due to a sufficient financial background. In this respect, the frequently used argument that public educational programs serve to facilitate access to higher education for members of low-income groups gains importance. As long as absolute risk aversion decreases, it has to be expected that members of low-income groups will invest less in their education than members of groups with higher incomes, even if they have the same educational abilities.

A general rise in incomes caused by productivity growth may also reduce the public role of education financing, but this would require a distinctly higher income effect on risk aversion. Indeed, the view that economic growth could lead to a stimulation of private demand for education has to be met with caution.
RISK, RESOURCES, AND EDUCATION

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Social Fractionalization, Political Instability, and the Size of Government

ANTHONY ANNETT*

This paper explores the relationship between the degree of division or fractionalization of a country's population (along ethnolinguistic and religious dimensions) and both political instability and government consumption, using a neoclassical growth model. The principal idea is that greater fractionalization, proxying for the degree of conflict in society, leads to political instability, which in turn leads to higher government consumption aimed at placating the opposition. There is also a feedback mechanism whereby the higher consumption leads to less instability as government consumption reduces the risk of losing office. Empirical evidence based on panel estimation supports this hypothesis. [JEL E62, O23]

When countries are heavily divided along ethnic, religious, communal, and regional lines, they are likely to experience bouts of political violence and are prone to the frequent breakdown of law and order. Clearly, serious conflict between competing groups, especially violent conflict, would be harmful to economic growth and the process of development. In such a divided country, can the government use expenditure to appease the competing groups and will this contain the potential conflict?

To isolate the sources of conflict, this paper follows the political economy literature in treating society as a collection of disparate groups, each concerned about its own interests at the expense of social welfare. Specifically, the degree

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of fractionalization within a country is measured along two dimensions, ethno-linguistic and religious. Although imprecise as a measure of the strength of competing groups, indices of fractionalization have the advantage of facilitating a broad cross-sectional empirical analysis with a large number of countries. The main hypothesis is that fractionalization leads inherently to greater levels of political instability and this imposes a political cost on the government as it risks being overthrown and losing any rents from being in power. Therefore the government attempts to placate the excluded groups by increasing the level of government consumption. This leads to a reduction in instability and thus buys the government a certain amount of insurance against being overthrown. Government consumption is thus used to reduce the level of "political risk."

A crucial assumption here is that the government does not use budgetary transfers to placate the other groups in society. This assumption is not unreasonable, given that many countries (especially in the developing world) lack the institutions necessary for a system of fiscal transfers and government consumption is often used for this purpose. For the period 1981–90, the ratio of government consumption to current central government expenditure (excluding interest payments) was 61 percent for all countries. This ratio varies substantially across countries depending on the level of development; it stands at 78 percent for low-income countries, 64 percent for middle-income countries, and 36 percent for high-income countries. A large component of government consumption is the wage bill of government workers. Another way to view the effect of consumption on instability is to think of the government using consumption to transfer fiscal resources to the various groups in order to reduce the level of discontent in society.

Empirically, it is found that higher fractionalization leads to higher government consumption in a reduced form equation. In a simultaneous equations estimation with government consumption and political instability treated as endogenous, it is shown that the channel by which fractionalization affects government consumption is via political instability. Higher fractionalization leads to higher instability, while higher instability leads to higher government consumption. In turn, higher government consumption leads to lower political instability. Thus government consumption can indeed appease competing groups and reduce the level of conflict in a society divided along ethnic and religious lines.

The model presented in this paper is related to a number of strands that have developed independently in the political economy literature. Most of the previous work on political instability and fractionalization focuses on the impact of these variables on economic growth, while the literature on the determinants of government consumption has not incorporated these political economy aspects in any systematic manner.

\[\text{Data are from the World Bank (1997). The classification of countries by level of development is taken from this source.}\]
One framework that could predict a positive relationship between fractionalization and government consumption is the common pool model. In this case, the government is inherently weak and each powerful group is able to extract resources from it. Given that society is composed of multiple self-interested groups who act noncooperatively, no single group has any incentive to constrain its demand for resources. Velasco (1997) suggests the common pool model as a reason for the perpetuation of fiscal deficits over time. Lane and Tornell (1996) and Tornell and Lane (1998) develop this idea in the context of a neoclassical growth model in which each group has common access to the capital stock. This leads to not only overconsumption and hence lower growth, but also a voracity effect in which groups actually increase more than proportionately their rate of appropriation after a positive shock to output.

The model presented in this paper posits an alternative link between fractionalization and government consumption, one that involves a strong central government strategically using government consumption to reduce the level of political risk. A society divided along ethnic or religious lines is potentially unstable; this channel has the advantage of making political instability endogenous.

There is a sizable body of literature on the negative effect of political instability on economic outcomes. Barro (1991, 1996) finds that political violence leads to lower growth in a cross section of countries. Alesina and others (1996) conclude that while instability leads to lower growth, there is no evidence that low growth affects the propensity for government change. Possible reasons for this link include the induced uncertainty and the disruption in market activities deriving from the instability (see Perotti, 1996). The exact mechanism, however, by which instability reduces growth remains unspecified. Another approach to measuring political instability is to devise an index capturing the key elements of social and political unrest. Alesina and Perotti (1996) take this approach, arguing that the disorder created by this form of political instability adversely affects productivity and the return to investment. This framework is adopted here and a similar index of political instability is derived.

In related work, Blomberg (1996) argues that the government can use defense spending as a partial insurance against political instability; instability inhibits growth while increased military expenditure decreases instability. As well as the standard result that political instability reduces growth, the author also reports significant results in the other direction: higher growth reduces the probability of coups. The role of military expenditure in this model is similar to that of government consumption in the present paper: the logic of Blomberg (1996) is that the “stick” of military expenditure reduces instability. In this paper, the “carrot” of nonmilitary government consumption reduces the instability that arises from the existence of diverse groups in society.

I. Theoretical Model

The model presented here is an adaptation of the standard neoclassical one-sector endogenous growth model in which a policymaker chooses an intertemporal path for consumption subject to a capital accumulation equation. It is assumed that the
economy comprises a number of groups, with the government controlled by a single group, or a coalition of different groups. All that matters is that the government does not represent every group in society. The most natural interpretation of consumption in this context is to think of it as the quantity of publicly provided goods provided by the government at any point in time.

The government derives utility from rents; this can be thought of as consumption allocated to the ruling group or as rents directly extracted from the economy’s resources. The government’s welfare at time 0 is the present discounted value of the sum of the instantaneous utility functions over time. Let private rents be denoted by $r_i$ and let the instantaneous utility function be $u(r_i)$; for tractability, it is assumed to be of the CRRA form so that:

$$u(r_i) = \frac{r_i^{1-\sigma}}{1-\sigma}$$

(1)

where $\sigma$ is the inverse of the intertemporal elasticity of substitution.

Since the government is selfish, in the sense that it is concerned purely with the welfare of its members, it faces some probability of being overthrown. This may take the form of a democratic loss of office or an extraconstitutional seizure of power. The probability of losing power is associated directly with the degree of political instability in the country. This political dimension can be added to the neoclassical growth model in an intuitive manner. Assume that the government has a constant positive subjective discount rate $\theta$. Following Blomberg (1996), a second component will be added to the discount rate, one that reflects the possibility of the government losing power.

More specifically, assume that the government faces a stochastic process of overthrow, $\pi$, which is constant per unit time. If $\pi$ is assumed to be constant at each point in time then the introduction of this sort of uncertainty has the effect of raising the discount rate by $\pi$. The instability generated by the constant probability of overthrow means that the government becomes more impatient. Rather than treat it as an exogenous variable, it is assumed that the government can influence $\pi$ through its consumption choices and that the probability of overthrow is directly related to the proportion of total consumption appropriated as rents by the ruling cadre. Hence the government must choose the optimal path of rents $r_i$ and government consumption $c_i$, knowing that this choice will directly affect the probability of remaining in power that period. Formally,

$$\pi(z_i) > 0, z_i = \frac{r_i}{r_i + c_i}.$$

(2)

We can think either of public goods or publicly provided private goods: it makes no difference in the context of this model. For tractability, private consumption is ignored.

One clear example of the latter would be funds deposited directly into the bank account of a kleptocratic leader.
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The intuition is straightforward: the excluded groups in society dislike the fact that the ruling cadre is deriving rents as opposed to providing consumption goods to benefit the populace as a whole. Satisfaction with the government in this model depends on the composition of government spending, rather than the total, or ratio relative to GDP, as agents view the government solely in terms of the proportion of total spending seized by the ruling cadre. The level of government consumption can be interpreted as insurance against political instability and the ensuing probability of overthrow in the sense that it is possible to placate the excluded groups with government consumption.

To complete the model, the equations of motion must be characterized. On the technological side, an AK model is assumed in which output \( y_t \) is linear in a single factor, capital \( k_t \), and \( A \) represents the level of technology, so that:

\[
y_t = A k_t. \tag{3}
\]

As always, the capital stock is a state variable and the government must decide how to allocate output between government consumption, investment, and rent-seeking. The accumulation equation can be written as:

\[
k_t = A k_t - c_t - x_t. \tag{4}
\]

There is a second state variable in this model, which shall be called “political capital”; it is defined as the accumulated probability of being in power at time \( t \), \( p_t \). From the assumptions made about \( \pi_t \), \( p_t \) can be written as follows:

\[
p_t = e^{-\int_0^t \pi_t(z_s) ds}. \tag{5}
\]

Furthermore,

\[
\dot{p} = -p_t \pi_t(z_t). \tag{6}
\]

---

4 This differs somewhat from Blomberg (1996), who uses the ratio to GDP. The present model is only concerned with the current flow of utility to the government and private agents, rather than investment.

5 McGuire and Olson (1996) argue along similar lines, claiming that even if a ruler is concerned only with personal rents, she has an incentive to provide public goods. In the authors' model, public goods are required to produce productive private goods, which in turn yield valuable revenues to the ruler.
Just as the government must choose between consuming or investing in physical capital, so it faces an intertemporal trade-off between extracting rents and building up political capital. The optimization problem is one involving two control variables and two state variables and can be written succinctly as

$$\max E \sum e^{-\int_{0}^{\infty} \left( x_{t-1} - c_{t} \right) dt}$$

subject to $k_{0} > 0$, (4), and (6). In solving this problem, the two control variables are redefined as $z_{t}$ and $w_{t}$, where $w_{t} = c_{t} + x_{t}$. The government chooses the total level of consumption and the allocation between rents and public goods. Note that (7) has an intuitive interpretation: the government maximizes the expected utility of rents at each future time period, where the expected value depends on the probability of being in power at each future date. The political preferences have the effect of turning the problem into one with an endogenous discount rate.6

The following first-order conditions can therefore be derived, where $\lambda_{t}$ and $\mu_{t}$ are the co-state variables for physical and political capital respectively.7

$$p_{t}z_{t}u'(x_{t}) = \lambda_{t}$$

(8)

$$w_{t}u'(x_{t}) = \mu_{t},$$

(9)

$$\gamma_{\lambda} = -(A - \theta)$$

(10)

$$\frac{u(x_{t})}{\mu_{t}} = (\theta + \pi_{t}) - \gamma_{\mu}.$$  

(11)

There are two transversality conditions that must be satisfied, namely:

---

6There is a large body of literature on endogenous discount rates, beginning with Uzawa (1968). Obstfeld (1990) works out a complete model for the case of one control variable. The model is also similar in nature to the endogenous fertility models of Becker and Barro (1988) and Barro and Becker (1989).

7A technical appendix, providing a detailed derivation of the steady-state plus an analysis of transitional dynamics, is available upon request from the author.
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\[
\lim_{t \to \infty} k_t \lambda_t e^{-\theta t} = 0 \quad (12)
\]

\[
\lim_{t \to \infty} p_t \mu_t e^{-\theta t} = 0. \quad (13)
\]

For this problem to give a meaningful solution, the following restriction must be imposed on the utility function:

\[
u(x_t) > 0. \quad (14)
\]

Given the particular utility function specified in (1), the condition that \( \sigma < 1 \) needs to be imposed. Equations (8) and (10) can be used to derive the growth rate of rents and consumption in this economy. First, the condition that \( \gamma_z = 0 \) is imposed; since \( \pi \) must be constant in steady-state, the growth rate of \( z_t \) must be zero. This has the implication that consumption and rents grow at the same rate; namely, \( \gamma_v = \gamma_x = \gamma_c \). Noting that \( \gamma_y = -\pi \) in steady-state, it follows that the common growth rate can be written as follows:

\[
\gamma_c = \gamma_x = \frac{1}{\sigma} [A - (\theta + \pi)]. \quad (15)
\]

This is the standard solution for an AK model; it implies that the growth rate of consumption (either private rents or general government consumption) is independent of the level of capital (see Barro and Sala-i-Martin, 1995). This equation can be rearranged in the following way:

\[
A = \sigma \gamma_c + \theta + \pi. \quad (16)
\]

The left-hand side represents the return to investment while the right-hand side represents the return to consumption. Given that \( \pi > 0 \) this implies that the return to consumption exceeds the return under the modified golden rule. The intu-
ition is straightforward: the possibility of overthrow biases the government's choice toward present consumption instead of investment. This leads to a Pareto-inefficient outcome.9

The levels solution shall now be derived. Manipulating equations (9) and (11) yields the following statement about the steady-state value of $z$:

$$z \pi'(z) = \frac{1-\sigma}{\sigma} [\theta + \pi(z)] - \frac{(1-\sigma)^2}{\sigma} A.$$  

(17)

Given the assumption that $\sigma < 1$, the right-hand side of (17) is always positive. Finally, the growth rate of capital can be written as follows:

$$\gamma_k = A - \phi, \phi = \frac{w^y}{k}.$$  

(18)

Given that the growth rate of consumption (rents) does not depend on the level of capital, it is possible to describe consumption (rents) at time $t$ as a function of initial consumption (rents). It is straightforward to show that this implies a proportional relationship between consumption (rents) and capital. If not, $\phi$ would not be constant in steady-state. Applying this result, and appealing to the linear production function in (3), it is clear that $\gamma_k = \gamma_c = \gamma_w = \gamma_k = \gamma_r$.

Setting $\gamma_k = \gamma_k$ the equilibrium level of $\phi$ can be solved for as follows:

$$\phi = \frac{1}{\sigma} (\theta + \pi) - \frac{1-\sigma}{\sigma} A.$$  

(19)

The solution can now be fully characterized by equations (17) and (19). Equation (17) allows for the retrieval of the equilibrium value of $z_t$, and plugging this into (19) gives $\phi_r$.10

To derive clearer results, a functional form on $\pi(z_t)$ needs to be imposed; in line with Blomberg (1996), the following linear specification is assumed:11

$$\pi(z_t) = \alpha + \beta z_t.$$  

(20)

The steady-state values of the main variables can be written as follows:

9Blomberg (1996) derives the same result in his model.

10This is the only possible solution to this dynamic model. All other paths can be ruled out either by the fact that they hit a boundary or by a violation of the transversality condition.

11None of the results depend specifically on the linear functional form.
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\[ z^* = \frac{1 - \sigma}{\beta(2\sigma - 1)}[(\theta + \alpha) - (1 - \sigma)A] \]  

\[ \phi^* = \frac{[(\theta + \alpha) - (1 - \sigma)A]}{(2\sigma - 1)} \]  

\[ \pi^* = \frac{\sigma}{(2\sigma - 1)}A + \frac{1 - \sigma}{(2\sigma - 1)}[(\theta - (1 - \sigma)A] \]  

\[ \gamma = \frac{\sigma}{(2\sigma - 1)}A - \frac{1}{(2\sigma - 1)}(\theta + \alpha). \]  

The optimal response of these variables to an exogenous increase in the probability of overthrow, defined here as an increase in \( \alpha \) or \( \beta \), will now be considered. In terms of the effects of an exogenous increase in the risk of overthrow on the government's choice of \( z_t \), there are two possibilities, depending on the preferences of the government. It either could respond by decreasing the level of \( z_t \) in an attempt to buy insurance against the increased political instability or it could decide that, given that it faces a higher probability for being ousted from power, it should seize more rents today.

The comparative statics for a change in \( p \) are particularly straightforward to analyze:

\[ \frac{\delta z^*}{\delta p} < 0; \quad \frac{\delta \phi^*}{\delta p} = \frac{\delta \pi^*}{\delta p} = \frac{\delta \gamma}{\delta p} = 0. \]  

This is because \( \beta z_t \) is a positive constant so an increase in \( \beta \) will necessarily be associated with a decrease in \( z_t \), leaving \( \pi_t \) unchanged. This has the property that \( \pi_t \) will always remain constant when \( \beta \) changes. Total consumption in the economy is unchanged: the only effect is a shift in the composition of \( w_t \) from rents to government consumption. The reduction in \( z_t \) is such that it completely counteracts the effect of the higher \( \beta \) and hence \( \pi \) is unchanged. This also ensures that there is no change in the growth rate.

The effect of an increase in the constant term, \( \alpha \), is more ambiguous. We can derive the following comparative static results:

\[ \frac{\delta z^*}{\delta \alpha} = \frac{1 - \sigma}{\beta(2\sigma - 1)}; \quad \frac{\delta \phi^*}{\delta \alpha} = \frac{1}{2\sigma - 1}; \quad \frac{\delta \pi^*}{\delta \alpha} = \frac{\sigma}{2\sigma - 1}; \quad \frac{\delta \gamma}{\delta \alpha} = \frac{-1}{2\sigma - 1}. \]  

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The signs of the above derivatives will depend on whether $\sigma$ is greater or less than one half. If $0 < \sigma < \frac{1}{2}$ then $\frac{\partial z}{\partial \alpha} < 0$, $\frac{\partial \phi}{\partial \alpha} < 0$, $\frac{\partial \pi}{\partial \alpha} < 0$, and $\frac{\partial \gamma}{\partial \alpha} > 0$. In this case, the government reacts to an exogenous increase in $\alpha$ by reducing both $z_t$ and $w_t$; this reduces the probability of overthrow and actually increases growth in the economy. If, on the other hand, $\frac{1}{2} < \sigma < 1$ then $\frac{\partial z}{\partial \alpha} > 0$, $\frac{\partial \phi}{\partial \alpha} > 0$, $\frac{\partial \pi}{\partial \alpha} > 0$, and $\frac{\partial \gamma}{\partial \alpha} < 0$. Here, the government is deciding to increase appropriation of rents. Total consumption is increasing, as is the probability of overthrow. Hence, growth is lower in this scenario. The government, responding to the higher exogenous probability of overthrow, responds by choosing present consumption at the expense of investment.

It is possible to view the effect of an increase in the number of excluded groups through the comparative static exercise of an increase in $\alpha$ or $\beta$. The idea is that a more fractionalized country is exogenously more unstable. One way to model a more fractionalized country is through a higher $\alpha$: holding everything else constant, a higher degree of fractionalization (more excluded groups in the economy) leads directly to higher political instability and hence a higher probability of overthrow. It is also possible to model an increase in fractionalization as an increase in $\beta$; the intuition here is that it is harder in terms of political capital to appropriate a fixed proportion of total consumption as rents in more fractionalized countries. In other words, the marginal cost of appropriating rents is greater in more fractionalized societies. A corollary is that the marginal benefit of “bribing” or “buying insurance” using government consumption is higher in more fractionalized countries. There is no single way to model fractionalization; this approach has the advantage of simplicity and consistency with the model.

In terms of the empirical predictions of the model, the clearest results pertain to the level of government consumption. At least for some parameter values, the model predicts that higher levels of government consumption should be observed in more fractionalized economies. In this model, higher fractionalization leads directly to greater levels of political instability and this is compensated for by higher levels of government consumption. This higher government consumption in turn indirectly reduces political instability. No clear predictions emerge for growth: if the $\beta$-mechanism is the channel of interest, then there should be no relationship between growth and fractionalization as the direct and indirect effects on political instability cancel exactly. The empirical section of this paper will focus exclusively between the interaction between fractionalization, political instability, and government consumption.

---

12 An alternative viewpoint is that, if a highly fractionalized country means each opposition group is really weak, it may be that the political cost of appropriating rents will be decreasing in fractionalized countries. Ultimately, it remains an empirical issue.
II. Measuring Fractionalization and Political Instability

The Concept of Fractionalization

This paper uses a number of indices of fractionalization, measured along ethnolinguistic and religious scales, as proxies for the number of competing groups in society. One weakness of this approach is that such a measure cannot distinguish between groups that are powerful and groups that are weak. This is the motivation employed by Lane and Tornell (1996) for eschewing measures of fractionalization as proxies for competing groups within society. In the present case, which focuses on instability, such indices may indeed be reasonable proxies for the degree of conflict within society. The intuition is that more fractionalized countries are inherently more prone to instability, and this necessitates placation through higher levels of government consumption. Of course, the indices will not measure the "intensity of conflict" between groups and this will remain a fundamental weakness.

The indices developed in this paper are defined as follows. For a given number of groups in society, the index measures the probability that two randomly selected individuals from the country in question will not belong to the same group. Formally, it can be calculated from the following formula:

\[
\text{Fractionalization} = 1 - \frac{1}{N} \sum_{i=1}^{M} \left( \frac{n_i}{N} \right)^2, \quad i = 1, ..., M.
\]

\(N\) is the total population and \(n_i\) is the number of people belonging to the \(i\)-th group. Specifically, two such indices will be defined, ethnolinguistic fractionalization and religious fractionalization. The former divides the country into ethnolinguistic groups while the latter concentrates on different religious groupings. Given data availability, these seem to be reasonable proxies for measuring the importance, though not the intensity, of group affiliation within each country. A further problem is that these indices will have no time-variance, and the assumption that they only change slowly over time will be maintained. This is important, because the empirical section described in Section III will use panel data from three decades.

Ethnolinguistic Fractionalization

A number of researchers have worked with indices of ethnolinguistic fractionalization in the past. Mauro (1995) presented such an index, derived from the World

\[\text{Mauro (1995) presented such an index, derived from the World}\]
Handbook of Political and Social Indicators by Taylor and Hudson (1972); this index in turn is based on data from 1960 and was constructed in the Soviet Union for the Atlas Narodiv Mira. Taylor and Hudson (1972) also consider two other indices of ethnolinguistic fractionalization. The second index was derived by Muller (1964), in a comprehensive study of the world's living languages. The third of the indices has the fewest number of countries and derives from Roberts (1962), from a study of second languages in Asia, Africa, and Latin America. Researchers in economics have tended to use the Soviet index, given the larger sample size, and the argument by Taylor and Hudson (1972) that the data are free from ideological bias. Furthermore, Easterly and Levine (1997) argue that the other two measures are flawed as they omit certain key groups within countries.

A number of papers in recent years have used the Soviet index to explore the effects of ethnolinguistic fractionalization on the overall macroeconomic environment. Canning and Fay (1993) argue that a more fractionalized population leads to lower productivity growth. Easterly and Levine (1997) focus on low growth rates in sub-Saharan Africa. They find that fractionalization of society along ethnic lines is associated with low schooling, political instability, underdeveloped financial institutions, distorted foreign exchange markets, high government deficits, and poor institutional quality. They conclude that such results provide evidence of strong rent-seeking behavior and inability to find consensus on the provision of public goods in highly fractionalized economies. Controlling for other variables, Sala-i-Martin (1997) argues that ethnolinguistic fractionalization is not robustly correlated with growth. In related work, Alesina, Baqir, and Easterly (1999) address the issue of ethnic fractionalization and the provision of local public goods in U.S. cities, counties, and metropolitan areas. They conclude that more ethnically diverse regions are associated with higher spending and deficits per capita, but with lower spending shares on basic public goods like education. Along similar lines, Kuijs (2000) presents international evidence that more divided societies spend less on public goods.

This paper starts afresh and builds a new index of ethnolinguistic fractionalization. The source of the data is the World Christian Encyclopedia (Barrett, 1982), which provides an extremely detailed breakdown of the ethnolinguistic groups within each country. It also has the advantage of coming from the same source as the data on religious fractionalization. The methodology is simply to include as many groups as possible at the most detailed level of breakdown. The larger the number of groups in a country, the higher the value of the resulting index. If 100 percent coverage can be achieved with the most detailed breakdown of ethnolinguistic groupings, this is the methodology adopted. Otherwise, a more aggregated group is taken.

This new index of ethnolinguistic fractionalization (Table 1) has a number of advantages over the previous indices, including the care taken to include as many groups as possible, the more recent data, and the larger sample size. In contrast,
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## Table 1. An Index of Ethnolinguistic Fractionalization

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Source: Author's calculations based on Barrett (1982).
the Soviet measure has data for only 119 countries in this sample. The other indices are even worse: the Muller index contains 102 countries in the sample, while the Roberts index reports only 47. From this point, the Soviet, Muller, and Roberts indices shall be denoted Index 1, Index 2, and Index 3, respectively. The correlation between the present index and the Taylor-Hudson indices are as follows: 0.84 for Index 1, 0.66 for Index 2, and 0.85 for Index 3. For the sake of comparison, the correlation between Index 1 and Index 2 is 0.83, and the correlation between Index 2 and Index 3 is 0.93.

**Religious Fractionalization**

Apparently, there has been no attempt in the literature to develop an index of religious fractionalization with a view to analyzing its impact on macroeconomic variables. The index of religious fractionalization used in this paper (Table 2), simply measures the probability that two randomly drawn people in a specific country will not belong to the same religious group, and hence measures the degree of fractionalization in society along a different dimension. Once again, the data used in compiling this index comes from the *World Christian Encyclopedia* (Barrett, 1982) and all the information pertains to the early 1980s. Any religion listed by Barrett (1982) as a distinct religion in a given country is included in the index. The only religion that is disaggregated is Christianity: the subdivisions include Catholicism, Protestantism, Eastern Orthodoxy, Indigenous Christianity, and Crypto-Christians. Other groupings include Islam, Hinduism, Buddhism, Judaism, Tribal religions, Shintoism, Chinese folk religion, as well as a plethora of smaller religions. Furthermore two secular categories are included: Nonreligious and Atheist. The data for the latter normally derives from membership of communist parties in the country in question.

The mean value of the religious fractionalization index is 0.38. The simple correlation of the indices of ethnic and religious fractionalization yields a coefficient of 0.39, suggesting a positive relationship between the indices but also confirming that each index captures a different dimension to the fractionalization within each country.

For the sake of comparison, and to assess robustness, a supplementary index of religious fractionalization called Index 4 has been compiled. The data used is from a different source and at a much less disaggregated level. Only ten cate-

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17 Some work has been done on the impact of religious affiliation and the size and efficiency of government. La Porta and others (1998) analyze the effect of religious affiliation and size and quality of government. The present study emphasizes religious division and does not consider the potential effects of religious affiliation.

18 Barro (1999) argues that the data do not significantly change over time.

19 Includes non-Roman Catholics.

20 Includes all Protestant denominations.

21 Usually associated with African and Caribbean countries.

22 This group is a residual, and estimate of the number of clandestine Christians in countries that do not grant religious freedom. It is included in the index as a separate group, as there is no way (from the data) of apportioning this group between the other denominations. It is only an issue in a small number of countries, and may serve as a proxy for the enhanced religious tensions in these countries.

23 The *Encyclopedia Britannica Online*. 

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## Table 2. An Index of Religious Fractionalization

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Source: Author's calculations based on Barrett (1982).
categories are included: Catholicism, Protestantism, Eastern Orthodoxy, Islam, Hinduism, Buddhism, Judaism, African indigenous, Chinese folk, and Official Atheism. The raw correlation between the two indices of religious fractionalization is 0.89.

Since religious and ethnic fractionalization represent two dimensions to the divisions within society, it makes sense to combine them (Table 3). Hence the index of fractionalization is defined as one half times the value of ethnolinguistic fractionalization plus one half times the value of religious fractionalization: this index will be the principle proxy for conflict between competing groups in this paper. \(^{24}\) From now on, the term fractionalization shall be taken to mean this combined index.

### Political Instability

There are a number of different approaches to the quantification of political instability. Unlike Londregan and Poole (1990), Blomberg (1996), and Alesina and others (1996), instability in this paper is not equated with the propensity of observing changes in government. Rather, the second approach is adopted and an index of political unrest is devised. Following other researchers in the area, a single index of political instability encompassing a number of different factors is created (Table 4). This is precisely the methodology employed by researchers such as Hibbs (1973), Vanieris and Gupta (1986), Gupta (1990), Alesina and Perotti (1996), and Perotti (1996).

The idea is that these factors capture different dimensions of political instability in the country. The following nine dimensions are employed: (1) genocidal incidents involving communal victims or mixed communal and political victims \((COMPOL)\), measured as a dummy variable; (2) the occurrence of a civil war, measured as a dummy variable \((WARCIV)\); (3) the number of assassinations per thousand population \((ASSASS)\); (4) the number of extraconstitutional or forced changes in the top government elite and/or its effective control of the nation’s power structure \((COUPS)\); (5) the number of illegal or forced changes in the top government elite, any attempt at such change, or any successful or unsuccessful armed rebellion whose aim is independence from the central government \((REVOLS)\); (6) violent demonstrations or clashes involving more than a hundred citizens involving the use of physical force \((RIOTS)\); (7) the number of major government crises, where a crisis is defined as any rapidly developing situation threatening to bring the downfall of the present regime, excluding instances of revolt aimed at overthrow \((CRIS)\); (8) the number of times in a year that a new premier is named and/or 50 percent of the cabinet posts are occupied by new ministers \((CABCHG)\); and (9) the number of basic alterations in a state’s constitutional structure, the extreme case being the adoption of a new constitution that significantly alters the prerogatives of the various branches of government \((CONSTCHG)\).\(^{25}\)

\(^{24}\)Values of the religious and combined (ethnolinguistic and religious) fractionalization indices for 148 countries are available upon request from the author.

\(^{25}\)These variables, including the definitions, are taken directly from the dataset of Easterly and Levine (1997).
### Table 3. An Index of Ethnolinguistic and Religious Fractionalization

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<td>Singapore</td>
<td>0.54</td>
<td>Venezuela</td>
<td>0.32</td>
</tr>
</tbody>
</table>

Source: Author’s calculations based on Barrett (1982).
These factors measure political instability along a number of different dimensions, all threatening the survival of the present government in some way. From the point of view of the model, it is necessary to develop an index of political instability that directly captures those aspects of social disruption that will lead the government to trade off private rents in order to devote resources toward the alleviation of this instability. The statistical methodology employed in deriving such an index is that of factor analysis. The index, which is called \( \text{INS} \), is defined as follows:
The factor analysis technique is designed to reduce the dimensionality of a variable by describing linear combinations of those variables that contain most of the information. In essence, the technique recovers the latent original variable by identifying a small number of common factors that linearly reconstruct the original variables. Vanieris and Gupta (1986) claim that there is "general agreement among students of the relevant literature" that there are two dimensions to political instability that can be disentangled by factor analysis. These dimensions represent the less and more violent events, the former placing weight on such variables as riots and demonstrations, the latter emphasizing those elements that result in deaths. Hibbs (1973) dubbed these factors "collective protest" and "internal war," respectively. This is indeed the pattern found with the data in question: the instability variable here reflects the more violent "internal war" factor. In this sense, the index should be interpreted primarily as one representing political violence.

One advantage of INS is that, unlike previous indices, it is time-varying with average decade values for the 1960s, 1970s, and 1980s. Data availability constrains the series to include only 108 countries.26

III. Empirical Results

Econometric Methodology

The goal of this section is to explore the data for evidence in favor of the hypothesis forwarded in Section II using data on fractionalization, instability, and government consumption. Pooled time-series cross-section data will be used, with three time observations corresponding to the decade averages for the 1960s, 1970s, and 1980s. The ten-year average is used to remove any short-term cyclical patterns in government consumption. To allow for decade-specific effects, decade dummies are included in each equation. The following sequence is adopted: first, single-equation estimation for government consumption is presented to ascertain the importance of fractionalization. Following from this, systems estimation techniques are utilized, treating both government consumption and political instability as endogenous.

The model of Section I argues that the government appropriates rents directly to benefit its own group and provides government consumption to the benefit of excluded groups. It is implicitly assumed that rents take the form of a direct transfer and do not show up in government consumption. Clearly, it may be the

26Values for this index are available upon request from the author.
case that some rents are included within government consumption. Unfortunately, this is impossible to disentangle empirically. Nonetheless, it is not unreasonable to posit that the bulk of government consumption benefits a larger group of citizens than the government faction alone. In an empirical study, Rodrik (1997b) finds no effect of resource rents on government consumption, and concludes that government consumption (in particular public employment) is driven more by social insurance considerations.

**Government Consumption Equations**

For most of the equations, the dependent variable employed will be the real share of nonmilitary government consumption to GDP. This is measured as the share of government consumption in GDP, measured in 1985 international prices from the Heston-Summers data set minus the ratio of government expenditure on defense (deflated by the investment price) to real GDP from the Barro-Lee data set. The reason for excluding military spending is straightforward: one may postulate that military expenditure is higher in more unstable countries, for reasons not accounted for by the model. There are a number of reasons for choosing to work with the Heston-Summers data. Most of the previous work in recent years on the determinants of government consumption has utilized this data source. A number of researchers, such as Ram (1987) and Rodrik (1996) maintain that the Heston-Summers data is superior to the more conventional data on government shares for the problem at hand given that the conventional data suffer from problems created by the differing relative price of government purchases across countries. Without making this adjustment, the actual size of government in poorer countries is biased downward.

Table 5 estimates a number of basic government consumption equations. It is necessary to produce a number of control variables: these will be the variables considered important in previous work in the field. There is a burgeoning literature in recent years on the determinants of the size of government; most of this work has focused on the importance of economic and demographic variables. The equations in Table 5, as will be the case with all equations in this paper, include decade dummies. To be able to compare these results with the results of the structural equations in the next section, the sample size is constrained to the 108 countries for which data are available on political instability.

---

27If the data were available, it would be possible to test compositional effects, along the lines of Alesina, Baqir, and Easterly (1999).
28Rodrik deals with economic, not political, risk.
29Penn World Tables 5.6.
30The investment deflator is chosen to convert military spending to international prices on the basis that investment and military spending probably involve fairly similar imported goods.
31This is exactly the way that Barro (1991, 1996) nets defense spending out of government consumption in order to measure that component of government consumption that does not enhance productivity (to this end he also excludes expenditure on education).
### Table 5. OLS Government Consumption Equations

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
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<td>0.02</td>
<td>0.02</td>
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<td>(1.00)</td>
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<td>(1.26)</td>
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<td>(0.57)</td>
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^Robust standard errors used; t-statistics in parentheses. Decade dummies included.

In terms of control variables, the age dependency ratio is one of the most common demographic variables in the literature, based on the idea that a higher proportion of young and old people in the population leads to greater demand for publicly provided services, especially health care and education. A higher rate of urbanization is postulated to lead to a higher demand for government services. Adding a political economy dimension, it is plausible to argue that the government is likely to favor the urban over the rural community. Alesina and Spolaore (1997) argue that there are economies of scale in the provision of public goods. Population is included as a proxy for country size; this variable is significant in Alesina and Wacziarg (1997).
In terms of economic variables, a lot of attention has been accorded to the role of income per capita in explaining the size of government. One of the oldest hypotheses in the literature is Wagner’s Law: this states that the demand for government services is income elastic so a higher GDP per capita is expected to lead to a higher share of government consumption in GDP. The reasoning behind this is that either government consumption is a luxury good or the administrative and regulatory costs increase with the level of economic development.

Cameron (1978) argued that more open economies were associated with larger government as measured by the share of consumption in GDP. Rodrik (1995, 1996, 1997a) argues that the government sector insulates the economy from the destabilizing effects of external shocks. According to this logic, countries with higher exposure to external risk tend to have larger governments. Therefore the positive coefficient on openness in a reduced form government consumption equation stems from the fact that more open economies are exposed to external risk, and in these economies, the government uses its own consumption as a buffer against this risk. Rodrik (1996) derives a theoretical measure of external risk, which is defined as the degree of openness multiplied by the standard deviation of terms of trade shocks: this is the definition of external risk in this paper.

Table 5 considers the significance of fractionalization, holding fixed the aforementioned control variables. Equation 1 shows that fractionalization (the linear average of the ethnolinguistic and religious fractionalization indices) is positive but completely insignificant when regional dummies are excluded. When the three regional dummies are entered into the model, the results change dramatically: in separate equations the point estimates on fractionalization, ethnolinguistic fractionalization, and religious fractionalization are all positive and significant. This is because there are significant regional patterns to both government consumption and fractionalization and therefore the exclusion of these regional effects may cause omitted variable bias in the fractionalization coefficients. For the remainder of the paper, regional dummies will be included in all equations.

In terms of magnitude, increasing fractionalization by one standard deviation leads to an increase in government consumption of 0.16 standard deviations: given that the standard deviation of government consumption is 7.87, a one standard deviation increase in fractionalization will increase government consumption by 1.29 percentage points. A one standard deviation in ethnolinguistic fractionalization will increase government consumption by 0.09 standard deviations; the equivalent number for religious fractionalization is 0.15. It can be concluded, as predicted by the theoretical model, that higher levels of fractionalization lead to higher levels of government consumption. Furthermore, this result

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33 The same is true for ethnolinguistic fractionalization and religious fractionalization.
34 In unreported results, there is no evidence that military expenditure is higher in more fractionalized countries.
35 The significance of fractionalization in affecting government consumption does not depend on the inclusion on any single regional dummy.
36 If government consumption is at its mean of 18.73 at the outset, then this implies a 7 percent increase.
holds up among both dimensions of fractionalization although the results, both qualitatively and quantitatively, seem to be weaker for ethnolinguistic than for religious fractionalization.

As for the control variables, the age-dependency ratio and openness are not significant in any of the equations. Real income per capita is negative and strongly significant in every equation: this conclusion is in line with Rodrik (1996), but in violation of Wagner's Law.\(^\text{37}\) The point estimate on the standard deviation of the terms of trade is negative and significant while the measure of external risk is positive and significant as predicted by the hypothesis of Rodrik (1996). Equation 5 adds the urbanization rate and a socialist dummy\(^\text{38}\) to the basic specification; neither is significant. In line with the results of Alesina and Wacziarg (1997), the log of population enters with a negative sign, evidence of economies of scale in the provision of public goods.\(^\text{39}\)

In unreported results, a number of robustness tests have been conducted, including sequentially dropping each country from the regression and using a larger sample (including countries for which no data are available on political instability). The basic results continue to hold. The same is true when the equations are estimated using random effects. In sum, it is clear from the reduced form government consumption equations that higher fractionalization leads to higher levels of government consumption. So far, this result is consistent with both the common-pool model of weak government dominated by strong interest groups and the model of Section II that emphasizes political instability as the key factor which links fractionalization and government consumption. This mechanism will now be explored in some detail.

**Systems Estimation**

From the model of Section II, the causality is expected to run as follows:

\[ \text{Fractionalization} \uparrow \Rightarrow \text{Instability} \uparrow \Rightarrow \text{Govt.Consumption} \uparrow \Rightarrow \text{Instability} \downarrow \]

This section treats both government consumption and political instability as endogenous variables.\(^\text{40}\) As a preliminary step in the systems estimation, OLS is used to estimate equations for government consumption and political instability, both with the other as an explanatory variable. The results can be seen in Table 6.

---

\(^{37}\)Wagner’s Law applies to the nominal, rather than the real, share of government consumption, suggesting that it is driven more by rising relative government wages in advanced economies.

\(^{38}\)This variable derives from Kornai (1992).

\(^{39}\)Alesina and Wacziarg (1997) argue that the Rodrik results are spurious, based on the fact that there is a strong negative correlation between openness and the size of a country. The external risk result, however, is robust to the inclusion of the population variable.

\(^{40}\)In related work, Alesina and others (1996) estimate a systems model in which economic growth and political instability are both endogenous. Instability is defined as the propensity of government change. Alesina and Perotti (1996) treat investment and instability as endogenous, where instability is defined as an index of political unrest.
<table>
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<th>Dependent Variables</th>
<th>Government Consumption (1)</th>
<th>Government Consumption (2)</th>
<th>Instability (3)</th>
</tr>
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<td>Instability</td>
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<td>Government consumption</td>
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</tr>
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<td>Africa dummy</td>
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<td>-0.31 (-3.03)</td>
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<td>-9.11 (-6.03)</td>
<td>-0.0003 (-0.02)</td>
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<tr>
<td>Fractionalization</td>
<td></td>
<td>6.24 (3.03)</td>
<td>0.37 (2.73)</td>
</tr>
<tr>
<td>Log GDP per capita</td>
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<td>-4.86 (-6.22)</td>
<td>-0.22 (-4.45)</td>
</tr>
<tr>
<td>Age dependency ratio</td>
<td>2.16 (0.65)</td>
<td>2.84 (0.87)</td>
<td></td>
</tr>
<tr>
<td>Openness</td>
<td>0.01 (0.65)</td>
<td>0.02 (0.91)</td>
<td></td>
</tr>
<tr>
<td>Sd. Dev. of terms of trade</td>
<td>-0.18 (-1.79)</td>
<td>-0.17 (-1.86)</td>
<td></td>
</tr>
<tr>
<td>External risk</td>
<td>0.003 (1.99)</td>
<td>0.003 (1.89)</td>
<td>0.002 (1.30)</td>
</tr>
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<td>Urbanization rate</td>
<td></td>
<td></td>
<td>0.16 (1.49)</td>
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<tr>
<td>Socialist dummy</td>
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</tr>
<tr>
<td>R²</td>
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<tr>
<td>N</td>
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</tbody>
</table>

Equations (1) and (2) of Table 6 estimate the familiar government consumption model, with instability as a regressor. The coefficient is negative and significant. All previous results hold up. The significance of instability is robust to the exclusion of fractionalization.

Equation (3) of Table 6 uses OLS to estimate an equation with instability as the dependent variable. The first question of interest is the importance of fractionalization in leading to political instability. As basic control variables, the log of GDP per capita, the urbanization rate, and the socialist dummy are all included. The literature on the determinants of political instability is not nearly as extensive as it is for government consumption. Income per capita captures the idea that
instability ultimately represents the dissatisfaction of the populace. Poorer countries, defined by those with lower real income per capita, should be more prone to instability holding other factors constant. It is especially important to include this variable to avoid picking up a spurious relationship between fractionalization and instability, caused by the fact that poorer countries are more unstable, and poorer countries are more fractionalized. Countries with higher rates of urbanization may be expected to lead to more political pressures that may find a violent outlet, given appropriate conditions. Finally, it may be the case that, holding standard of living constant, the particular policies followed in socialist economies may lead to more inherent instability. As with all equations, the regional and decade dummies are included.

With these basic control variables alone, the coefficient on fractionalization is positive and significant, suggesting that there is indeed a relationship between fractionalization and instability. The coefficient on income per capita is negative and highly significant, suggesting economic dissatisfaction is indeed a principal source of political instability. The coefficients on the urbanization rate and the socialist dummy possess the anticipated signs, but are not significant at conventional levels. The coefficient on government consumption is negative and significant, suggesting that countries that devote more resources to government consumption are associated with lower levels of political instability.

It can be concluded from the OLS results of Table 6 that there is a strong negative association between government consumption and instability. The exact mechanism, however, remains to be determined. A priori, it is possible to make causality arguments in both directions. It may be the case that political instability (especially violent instability of the sort considered here) lessens the government’s ability to collect tax revenue or borrow on international markets and this constrains actual government consumption to be lower than desired. An alternative hypothesis argues, in a Hobbesian manner, that the causality runs in the opposite direction with lower levels of government consumption breeding discontent among the populace which may erupt into violent political instability. The second explanation, of course, is the one provided by the model of Section I.

An attempt to disentangle these effects will be made using Two Stage Least Squares (TSLS), treating government consumption and instability as endogenous variables. If this hypothesis is correct, then the estimates from Table 6 will suffer from simultaneous equations bias. From the full set of exogenous variables, the following variables are selected as instruments for government consumption: the age-dependency ratio, openness, the standard deviation of terms of trade shocks, and external risk (the interaction of openness with the standard deviation of terms

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41 Ades and Glaeser (1995) postulate that the direction of causality runs from political instability to urbanization.
42 The results are not reported, but both ethnolinguistic fractionalization and religious fractionalization are just as significant when entered independently.
43 It is especially important to remove military expenditure from government consumption in this case. In unreported results, adding military spending as a percentage of GDP to this equation yields a coefficient that is positive and significant.
of trade shocks). Fractionalization, the urbanization rate, and the socialist dummy are used to instrument for instability. Income per capita and the regional and decade dummies are included in both equations. Basically, equations (1) and (3) of Table 6 are being re-estimated using the technique of instrumental variables, treating government consumption and instability as endogenous. This will allow for testing the validity of the hypothesized feedback mechanism.

It can be seen from Table 7 that the results are quite favorable. The key result is that while the effect of government consumption on instability remains negative and significant, the effect of instability on government consumption is now positive and significant. Furthermore, the point estimate on fractionalization in the instability equation remains positive and significant. There is evidence, therefore, that the hypothesized chain of events is indeed borne out by the data. Higher fractionalization leads directly to higher instability. In response, government spending rises and this has the consequent effect of leading to lower political instability.

Table 7 also reports the test statistics associated with the Chi-squared test of overidentifying restrictions for each equation; this basically tests the relationship between the instruments and the error term in the structural equation. The resulting low statistics suggest that the instruments in each case are validly excluded. Confidence in this result must be tempered somewhat by the weakness of the instruments for political instability. This system was also estimated by Three Stage Least Squares; the results are not reported here as they are virtually identical to the TSLS results.

Returning to the system estimated by equations (1) and (2), it is the case that an increase in fractionalization by one standard deviation leads to an increase in instability of 0.1144, which is 0.28 of a standard deviation. An increase in government consumption by one standard deviation leads to a fall in the instability index of 0.39 units, or 0.96 of a standard deviation. Finally, an increase in instability by one standard deviation leads to a rise in the ratio of government consumption to GDP by 7.89 percentage points, corresponding to one standard deviation. Let us now consider the total effect of a one standard deviation increase in fractionalization: first, instability increases by 0.1144. This in turn leads government consumption to rise by 2.203 percentage points. Finally, this causes instability to decline by 0.1102 units. As the system equilibrates, the overall effect on instability is an increase by 0.004, or 0.01 of a standard deviation. It should also be noted that the direct effect of fractionalization on instability counteracts the indirect effect through government consumption, leaving virtually no aggregate effect on instability; this was a conclusion of the model in Section I, for reasonable specifications of the instability function.

Table 8 repeats the TSLS specification of Table 7, but replaces fractionalization by both ethnonational and religious fractionalization. In the instability equation, both of these variables are significant in equations (2) and (4),

44The p-value is 0.07.
45Some experimentation with different assumptions regarding excludability of the instruments was undertaken. In particular, it is conceivable that external vulnerability is an instrument for instability. The Rodrik variables, however, fail the Chi-squared test in this case.
46The means of government consumption and instability are 18.73 and 0.28, respectively.
respectively, suggesting that ethnic and religious fractionalization independently lead to higher political instability. Similarly, the strong evidence that government consumption decreases instability remains. The only difference between this specification and that of Table 7 is that in the former, the point estimates on instability in the government consumption equation fall just outside the standard significance levels. In both equations (1) and (3) of Table 8, the p-values for instability fall to 0.12, from 0.07 when fractionalization is used as an instrument for instability. It can therefore be concluded that it is important to take both dimensions of fractionalization into account.

Table 9 assesses robustness by considering the sensitivity of the model to a number of alternative indices of fractionalization and instability. All four alternative fractionalization indices exhibit significant positive effects on instability.
Table 8. Ethnic and Religious Fractionalization in a TSLS Model

<table>
<thead>
<tr>
<th>Dependent Variable</th>
<th>Government Consumption</th>
<th>Instability</th>
<th>Government Consumption</th>
<th>Instability</th>
</tr>
</thead>
<tbody>
<tr>
<td>Instability</td>
<td>14.71 (1.55)</td>
<td>-0.06 (-3.79)</td>
<td>21.41 (1.57)</td>
<td>-0.05 (-3.78)</td>
</tr>
<tr>
<td>Government consumption</td>
<td></td>
<td>-0.06 (-3.79)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Africa dummy</td>
<td>-2.87 (-1.25)</td>
<td>-0.33 (-2.68)</td>
<td>-2.97 (-1.04)</td>
<td>-0.37 (-2.91)</td>
</tr>
<tr>
<td>Latin America dummy</td>
<td>-6.36 (-3.39)</td>
<td>-0.21 (-2.02)</td>
<td>-7.07 (-2.99)</td>
<td>-0.18 (-1.81)</td>
</tr>
<tr>
<td>East Asia dummy</td>
<td>-13.03 (-3.05)</td>
<td>-0.21 (-1.24)</td>
<td>-15.45 (-2.72)</td>
<td>-0.24 (-1.38)</td>
</tr>
<tr>
<td>Ethnic fractionalization</td>
<td>0.31 (2.46)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Religious fractionalization</td>
<td></td>
<td>0.33 (2.45)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Log GDP per capita</td>
<td>-4.04 (-3.33)</td>
<td>-0.39 (-4.43)</td>
<td>-3.89 (-2.53)</td>
<td>-0.44 (-5.10)</td>
</tr>
<tr>
<td>Age-dependency ratio</td>
<td>1.23 (0.26)</td>
<td>0.07 (0.01)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Openness</td>
<td>0.04 (1.30)</td>
<td>0.06 (1.29)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Sd. Dev. of terms of trade</td>
<td>-0.35 (-1.70)</td>
<td>-0.41 (-1.62)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>External risk</td>
<td>0.006 (2.06)</td>
<td>0.007 (1.96)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Urbanization rate</td>
<td></td>
<td>0.003 (1.39)</td>
<td>0.004 (1.84)</td>
<td></td>
</tr>
<tr>
<td>Socialist dummy</td>
<td>0.23 (1.31)</td>
<td>0.23 (1.36)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \lambda^2 )</td>
<td>0.41 1.48</td>
<td>0.87 2.52</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Robust standard errors used; t-statistic in parentheses. Decade dummies included.

Likewise, the coefficient of government consumption is always negative and significant in each case. As usual, the most fragile result is the effect of instability on government consumption. Another robustness test replaces the instability index with that of Alesina and Perotti (1996) and Perotti (1996); this index includes democracy among its factors under the hypothesis that instability is biased downward in nondemocratic countries. Although the hypothesized channel is much weaker with this index, there is still evidence of it in operation. As a final robustness test, the TSLS equations are reestimated, sequentially dropping each country in order to test the sensitivity to outliers.47

47These results are unreported, and are available upon request from the author.
In sum, it can be concluded that the mechanism is reasonably robust, the main deficiency being the paucity of good instruments for instability. Furthermore, it is important to take account of the two dimensions to fractionalization, ethnic and religious, in unison.

IV. Conclusion

This paper has presented a hypothesis concerning the interaction between fractionalization (as measured along ethnolinguistic and religious dimensions), government consumption, and political instability. Specifically, higher fractionalization is held to lead exogenously to greater political instability. Partial insurance against the political risks engendered by the instability can be bought by raising the level of consumption. The reduced form evidence shows a strong positive relationship between fractionalization and government consumption. This paper shows that endogenizing political instability is crucial and that once a system with two endogenous variables is estimated, the effect of fractionalization on government consumption is seen to operate through the political instability channel.

Furthermore, it is clear that governments use government consumption as a buffer against political instability. One lesson to be drawn from this paper is that it is possible for a country to reduce the level of instability, even if the country is
divided into competing groups. This involves the use of government consumption. Permitting a government to increase its consumption will reduce the level of political instability, and numerous researchers have pointed out that political instability is detrimental to growth.

In terms of future research, it would be useful to analyze whether it is government consumption, pure transfers, or government investment that is most effective in reducing political instability. It would also be interesting to endogenize military spending, to determine whether military or non-military expenditure is more effective in reducing instability. In terms of fractionalization, it would be useful to derive an index giving more weight to a country with a small number of large groups to test the theory that a few powerful groups are more destabilizing than a large number of disparate groups. Furthermore, inequitable income distribution could be considered as another social fractionalization factor alongside ethnicity and religion.

REFERENCES


SOCIAL FRACTIONALIZATION


Anthony Annett


Deposit-Refund on Labor: A Solution to Equilibrium Unemployment?

BEN J. HEIJDRA and JENNY E. LIGTHART*

The paper studies the employment effects of a deposit-refund scheme on labor in a simple search-theoretic model of the labor market. It is shown that if a firm pays a deposit when it fires a worker, to be refunded when it employs the same or another worker, the vacancy rate increases and the unemployment rate declines. The scheme introduces rigidities in the labor market, however, which may be undesirable in countries wanting to liberalize their labor markets. [JEL J3, J68]

Some people return empty bottles to the store because they find littering and nonrecycling unacceptable from an environmental point of view. Many people, however, are motivated to return the bottles only because there is money to be made in the form of a deposit that is refunded. The literature on environmental economics has shown that these deposit-refund schemes are efficient instruments to deal with the negative environmental externalities resulting from illicit dumping.1 If unemployment is considered to create negative externalities,2 one could argue for using a deposit-refund system in the labor market. Why not have

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2 Phelps (1997) claims that unemployment is associated with negative externalities such as crime, poor health, and cultural deprivation.
the firm pay a deposit when it fires a worker, to be refunded when it (re-) hires that or another worker?

A deposit-refund scheme on labor is essentially a tax on firing an employee matched by a subsidy on hiring a (new) worker. The deposit acts as a “stick” by providing an incentive to move away from socially “undesirable” behavior, whereas the refund acts as a “carrot” by providing an incentive to move toward “desirable” behavior. By seeking such a policy mix, one can contain the fiscal implications of hiring subsidies; firing taxes reduce firms’ incentives to lay off workers and generate government revenues to (partly) finance public spending on hiring subsidies. Schemes that rely solely on hiring subsidies to boost employment may encourage firms to generate high labor turnover—assuming that the subsidy exceeds the recruitment and training costs to the firm—with a view to securing a net gain. Fay (1996) argues that close monitoring in employment-subsidy programs is important to make sure that subsidies are not used to replace subsidized workers whose subsidy has ended nor to recruit workers for recently created temporary vacancies.

The present proposal is not merely academic. Recently, Tilburg University has implemented a subsidy-penalty scheme with a view to increase the proportion of female associate and full professors among its staff. When a new female professor is recruited by a department, the Board of the University disburses a subsidy of 75,000 guilders to that department, which has to pay a penalty of the same amount when a female professor quits or is laid off. This measure was introduced to replace a 50,000-guilder hiring subsidy on female professors, which was not successful because of high quit rates.

This paper analyzes the long-run unemployment effects of deposit-refund schemes on labor. To this end, a modified version of Pissarides’ (1990) search-theoretic model of the labor market is employed. This framework allows for flows in the labor market from the creation of new jobs and the (exogenous) destruction of superfluous jobs. It assumes a search process in which firms offering vacancies and unemployed job-seeking workers are brought together in a stochastic fashion. In this type of model, vacancies and unemployment occur simultaneously in equilibrium. Pissarides’ model thus differs radically from the aggregate (classical) labor market model in which only unemployed workers (vacancies) are observed when the wage is above (below) its equilibrium value. The model features a search externality that is represented by agents’ contact probabilities as a function of labor market tightness: the presence of an additional firm with a vacancy makes it easier for workers to find jobs but harder for firms to fill vacancies. Similarly, an additional unemployed worker makes it more difficult for workers to locate jobs but easier for vacancies to be filled with workers.

So far, the literature on unemployment policy has not studied deposit-refund systems in the context of labor markets. There exists, however, a substantial liter-
DEPOSIT-REFUND ON LABOR

Nature on hiring subsidies, on the one hand, and firing costs on the other. Informal discussions on the beneficial employment effects of hiring subsidies can be found in Fay (1996), Martin (1998), OECD (1994, 2000), and Phelps (1997). The earlier theoretical literature on employment subsidies, which assumes a clearing labor market and homogeneous labor, stresses their employment-enhancing effect, but has neglected the government's budget constraint. Recently, Martin (1998) and the OECD (2000) have informally discussed the fiscal and employment implications of various labor subsidies taking into account the heterogeneity of labor. Formal treatments of labor subsidy policies in a search-theoretic context are developed by Millard and Mortensen (1997) and Mortensen and Pissarides (2001). They show that hiring subsidies reduce unemployment duration, but have an ambiguous effect on the unemployment rate.6

The theoretical literature on firing costs—with central contributions including Bertola (1990, 1992), Bentolila and Bertola (1990), and Booth (1997)—generally finds that employment protection provisions reduce the variation of employment over the business cycle, but the effect on the average level of employment is ambiguous. Firing costs create disincentives for firms wanting to fire workers but also discourage hiring because of the discounted costs of possible firing in the future. Bertola (1992) shows that when discount and attrition rates are positive, the average employment effect of firing costs depends on its effect on the marginal product of labor (which is negative), as well as the relative slopes of the labor demand curve in booms and slumps. He argues that if hiring and firing cycles are long—in which case attrition and discount rates are large and thus the present discounted value of firing costs small—employment may increase. However, no consensus has been reached yet in the empirical literature on the employment effects of firing costs (see Hunt, 1994).

The model at hand shows that deposits and refunds on labor push up economy-wide wages, increase the vacancy rate, and reduce the equilibrium rate of unemployment. Intuitively, the scheme provides an implicit subsidy (that is, the interest earned on the "deposit") to firms hiring new employees. The model gives a very stylized description of the labor market, however; employees work (full time) during their entire life, they do not voluntarily quit jobs, and workers are either employed or unemployed, in which case they are looking for a job. As the discussion section argues, if firms can freely set working hours, deposit-refund schemes on labor are likely to induce firms experiencing bad business conditions to impose shorter working hours on their workforce as a substitute for firing them. Accordingly, the variance of employment over the business cycle is stabilized. Some observers may claim that this contributes to undesirable rigidities in the labor market.

5Experiments with employment subsidies in the United States and Italy—as reported in Woodbury and Spiegelman (1987) and Felli and Ichino (1988)—have generally been successful in increasing employment.

6These papers extend Pissarides' (1990) framework by endogenizing the job destruction process. Because their analytical framework yields indeterminate employment effects of hiring subsidies, this paper employs the much simpler Pissarides model, which assumes a fixed job destruction rate. As a result, the model cannot take into account firms' incentives to replace nonsubsidized employees by subsidized (the so-called displacement effect).
I. A Search Model of the Labor Market

Suppose that there is a fixed labor force, \( L \), of which a fraction \( u \) is unemployed per unit of time.\(^7\) The flow into unemployment originates from an exogenous job destruction process that causes a proportion \( s \) of occupied jobs to dissolve. Separations may be caused by structural shifts in demand or by changes in productivity and are interpreted as “firing” because in equilibrium workers strictly prefer employment to unemployment. Only unemployed laborers search for a job, thus excluding on-the-job search by workers. Likewise, the firm is not searching for workers to replace incumbent (but unsatisfactory) workers. Vacancies are created by new firms or by incumbent firms reopening previously destroyed jobs.

At each instant of time, \( uL \) unemployed workers and \( vL \) vacancies on offer engage in a stochastic matching process, where \( v \) is the vacancy rate. The number of successful matches depends on the number of unemployed workers, \( U = uL \), and the total number of vacancies, \( V = vL \), according to the following matching function:

\[
xL = G(U, V),
\]

where \( xL \) is the total number of matches, \( x \) denotes the matching rate, and \( G(.,.) \) is a linearly homogeneous function. By defining \( g = xUV = x/v \), the matching function can be rewritten to yield \( g = g(\theta) \), with \( g'(\theta) < 0 \).\(^8\) The probability that a firm finds a worker in the time interval \( dt \) is given by \( g(\theta)dt \), where \( \theta = V/U \) denotes the vacancy-unemployment rate, which is a measure of labor market tightness. Then, the expected duration of a vacancy is given by \( 1/g(\theta) \). Similarly, \( f(\theta) = \theta g(\theta) \) is the probability of an unemployed worker finding a job, so that \( 1/f(\theta) \) is the expected duration of an unemployment spell.

In the aggregate, unemployment evolves according to the difference between the average number of unemployed workers who find a job, that is \( f(\theta)uLdt \), and the average number of workers entering the unemployment pool, \( s(1 - u)Ldt \). This yields the following equilibrium rate of unemployment:

\[
u = \frac{s}{s + f(\theta)},
\]

which is strictly greater than zero.

Wages are taken as given during the search process. After an individual worker and firm meet each other, however, wages are determined through negotiations.

---

\(^7\) This section builds on Pissarides (1990), which is amended to allow for a deposit-refund scheme on labor.

\(^8\) Job-seeking workers flow out of the unemployment pool according to a Poisson process with rate \( x/v \).
Firm Behavior

Assume that there are many risk-neutral firms, each of which has one job that is either filled or vacant. If the job is filled, the representative firm rents physical capital, \( k \), at the constant rate of interest, \( r \), to produce output according to a constant returns-to-scale production function, \( F(k, 1) \), which satisfies the usual properties. If the job is vacant, the firm actively searches for a suitable worker and is incurring search cost, \( y_0 \), per unit of time.

Suppose that a firm hiring a worker receives a fixed one-off subsidy of \( b \) from the government, but when it fires that worker it must pay a tax of \( b \). This is like a variant—in fact, the reverse—of a cash-refund scheme on bottles where consumers have to pay a fixed amount on buying a filled bottle and receive a refund (set at the same rate) when the empty bottle is properly disposed of. In the model, it is assumed that workers do not voluntarily quit jobs (in which case the firm would not need to pay a firing tax).

Let \( J_F \) and \( J_V \) denote the present value of profits of a firm with a filled job and a vacancy, respectively. Assume perfect capital markets so that firms can borrow freely at the market rate of interest. Then, the following arbitrage equation can be derived for a firm having a vacancy:

\[
\begin{align*}
\frac{r}{} = -y_0 + g(\theta)[J_F + b - J_V].
\end{align*}
\]

Equation (3) says that the capital cost of the firm, \( rJ_V \), should equal the firm’s expected return on investment, which consists of two parts: (i) the fixed-search costs that are incurred per time unit; and (ii) the expected capital gain when it finds a worker. The latter is equal to the sum of the gain in present value from filling the vacancy (that is, \( J_F - J_V \)) and the subsidy payment received from the government, weighted by the probability of finding a suitable candidate. The number of firms/jobs is determined by the zero-profit condition—free entry and exit of firms will occur until all profit opportunities from new vacancies are equal to zero. This implies the following expression for the present value of an occupied job:

\[
J_V = 0 \Rightarrow J_F = y_0/g(\theta) - b.
\]

Equations (3)–(4) show that \( b \) acts like an implicit subsidy to firms with a vacancy: the expected search costs, \( y_0/g(\theta) \), are reduced by the subsidy payment received from the government.
For a firm with a filled job, the steady-state arbitrage equation reads as follows:

\[ rJ_F = F(k,1) - (r + \delta)k - w - s[J_F + b], \] (5)

where \( w \) denotes the wage rate and \( \delta \) is the rate of depreciation. The left-hand side of equation (5) represents the expected return on an occupied job, which is composed of the surplus created in production and the expected capital loss owing to job destruction (that is, \( s(J_F + b) \)). If the job is destroyed, the firm not only loses the value of the occupied job, but must also pay back the deposit on its worker to the government. Since the job destruction rate, \( s \), is exogenous, the firm can do nothing to reduce the probability of an adverse job-destroying shock.

The firm chooses the amount of capital it wants to rent such that the value of the firm is maximized:

\[
\max_k (r + s)J_F = F(k,1) - (r + \delta)k - w - sb. \tag{6}
\]

This yields the usual condition equating the marginal product of capital to the rental charge on capital:

\[ F'_k(k,1) = r + \delta, \tag{7} \]

and the marginal condition for labor:

\[ F'_N(k,1) - w = \frac{(r + s)\gamma_0}{g(\theta)} - rb, \tag{8} \]

where \( N \) is employment and subscripts denote partial derivatives. Equation (8) shows that labor receives less than its marginal product, owing to the positive search costs. However, the capital value of the deposit acts like a subsidy on the use of labor.

**Worker Behavior**

Workers are homogeneous, live forever, and are risk-neutral. Therefore, they care only about the expected discounted value of their income. It is assumed that each worker with a job supplies one unit of labor inelastically. Let \( Y_E \) and \( Y_U \) denote the present-discounted value of the expected stream of income of, respectively, an

\[ \text{Note that } (r + s)J_F = F_E(k,1) - w - sb. \text{ Equation (8) is obtained by using this result in equation (4).} \]
employed worker and an unemployed worker. An unemployed worker receives an exogenously given income, \( z \), during his or her search and expects to move into a job with probability \( f(\theta) \). Then, the following steady-state arbitrage condition can be derived for a worker without a job:

\[
    rY_U = z + f(\theta)\left[ Y_E - Y_U \right].
\]  

(9)

The arbitrage condition says that the return on human capital of an unemployed worker during search—that is, the reservation wage\(^{13}\)—should equal expected income, which consists of the imputed value of leisure and the capital gain from finding a job.

A worker with a job earns a wage, \( w \), and loses that job at an exogenous rate \( s \). Then, in steady state, permanent income of an employed worker, \( rY_E \), should equal expected income, which exceeds wage income to compensate for the risk of becoming unemployed (and suffer a capital loss of \( Y_E - Y_U \)):

\[
    rY_E = w - s\left[ Y_E - Y_U \right].
\]  

(10)

Solving equations (9) and (10) yields expressions for \( Y_E \) and \( Y_U \):

\[
    rY_E = \frac{r(w - z)}{r + s + \theta g(\theta)} + rY_U, \quad rY_U = \frac{(r+s)z + \theta g(\theta)w}{r + s + \theta g(\theta)},
\]  

(11)

where the first expression shows that for anybody to be willing to search for a job, wages need to exceed the imputed value of leisure.

**Wage Setting**

When an unemployed worker and a firm offering a vacancy meet, a pure economic rent is created by the encounter, which is equal to the sum of the expected (net) search costs of the worker and the firm. Upon separation this rent will be lost. The division of the rent of a particular firm-worker pairing \( i \) is a matter of bargaining over the wage rate, \( w_i \). Using the Nash bargaining solution, the wage rate is set such that the weighted sum of the worker's and firm's net returns is maximized taking behavior in the rest of the labor market as given:

---

\(^{12}\)This may represent the imputed value of leisure or income earned in the hidden economy.

\(^{13}\)This could also be interpreted as the household's permanent income: the amount the unemployed worker can consume without running down his or her capital.
Max $\Omega \equiv \beta \log \left( Y_u - Y_i \right) + (1 - \beta) \log \left( J_i + b - J_v \right)$, \hfill (12)

where the coefficient $\beta$ (0 $\leq$ $\beta$ $\leq$ 1) can be interpreted as a measure of the worker's bargaining strength and $Y_u$ and $J_v$ are the "threat points" of the worker and the firm, respectively. The rent-sharing rule derived from (12) is given by:

$$\frac{Y_u - Y_i}{J_i + b - J_v} = \frac{\beta}{1-\beta}.$$ \hfill (13)

This yields the following wage equation for worker $i$:\hfill (14)

$$w_i = (1-\beta) \gamma_u + \beta \left[ F_N(k, 1) + rb \right].$$ \hfill (14)

In symmetric equilibrium, each firm with an occupied job chooses the same capital stock, that is, $k = k_i$. Accordingly, all workers are equally productive so that the wage rate is the same for all worker-firm pairs, $w = w_i$. Using equations (13), (9), and (14), the reservation wage can be written as $rY_u = z + \beta \gamma_i (1 - \beta)$. Now, by combining this with equation (4), the aggregate wage equation follows:

$$w = (1-\beta)z + \beta \left[ F_N(k, 1) + rb + \gamma_i \right].$$ \hfill (15)

The economy-wide wage is a weighted average of the imputed value of leisure and the firm's surplus, which consists of the marginal productivity of labor, the expected search costs that are saved if a deal is struck, and the implicit subsidy on labor. Accordingly, the deposit rate pushes economy-wide wages up.

II. Model Solution and Allocation Effects

Market Equilibrium and Government

The full model can be summarized by equations (2), (7), (8), and (15). The endogenous variables of the model are $u$, $k$, $w$, and $\theta$. For a fixed capital stock, the model can be solved in a recursive fashion. Equation (7) shows that the optimal

\textit{14}Substituting the arbitrage equation for worker $i$, that is, $rY_u = w_i + s(Y_u - Y_E)$, and the expected capital gain to firm $i$ of filling a vacancy, that is, $(r + s)\dot{F} = F_N(k, 1) - w_i - sb$, yields equation (14).
DEPOSIT-REFUND ON LABOR

capital stock, $k^*$, is determined by the rental rate of capital, that is $k^* = H(r + \delta)$, with $H' > 0$. Given the capital stock, equations (8) and (15) yield equilibrium values for $\theta$ and $w$, and $\theta$, in turn, determines the equilibrium rate of unemployment (see equation (2)). Because the rate of interest is exogenously given, the wage rate is the only price variable in the model.

The analysis so far has implicitly assumed that the government finances the deposit-refund system through lump-sum taxes on households. The government cannot issue debt and does not provide unemployment benefits to workers. At introduction, all firms having a filled job receive $b$ from the government, so that a lump-sum tax of $T(0) = (1 - u)Lb > 0$ is needed to balance the government budget at time zero. Once the deposit-refund system is operational, job destruction leads to net government receipts, $s(1 - u)Lb$, and new job matches cause net government outlays of $f(\theta)uLb$. Using equation (2), it can be easily shown that the deposit-refund system is budgetary neutral in the steady state.

The model can be summarized with the aid of Figure 1. Panel (a) shows equilibrium for labor market tightness and wages. The curve $ZP$ represents the zero-profit condition—see equation (8)—and is downward sloping in the $(w, \theta)$ space. An increase in wages makes job creation less profitable at a given level of search cost. To restore the zero-profit equilibrium, the search cost to firms must decrease, that is, $g(\theta)$ must increase through a fall in $\theta$. $WS$ is the wage-setting curve—see equation (15)—and is upward sloping: at higher labor market tightness the worker receives a larger share of the search costs that are foregone when a deal is struck with a firm. The intersection of the $ZP_0$ and $WS_0$ curves yields the equilibrium point $E_0$. In panel (b) of Figure 1, the line $LMT$ through the origin represents the equilibrium vacancy-unemployment ratio (the indicator for labor market tightness). The curve $BC$ is the Beveridge curve—represented by equation (2)—which is downward sloping and convex to the origin. Intuitively, when there are less vacancies, unemployment is higher because unemployed find it more difficult to find a job. Equilibrium vacancies and unemployment are at the intersection of the $LMT_0$ and $BC$ curves.

Allocation Effects

The allocation effects of an increase in the deposit, $b$, are shown in Figure 1. The appendix derives these results analytically. In panel (a), the zero profit curve shifts up from $ZP_0$ to $ZP_1$ because the interest payments the firm earns on an employed worker—a so-called implicit subsidy—increase the value of an occupied job. These interest payments also increase wages via the wage setting equation, as represented by an upward shift of $WS_0$ to $WS_1$. Both wages and the vacancy-unemployment ratio rise; that is, the new equilibrium, $E_1$, lies to the northeast of the initial equilibrium $E_0$. In panel (b), the $LMT$ curve rotates counterclockwise from $LMT_0$ to $LMT_1$, whereas the position of the Beveridge curve is unaffected by the increase in the deposit. Given the increased equilibrium vacancy rate, unem-

15See Blanchard and Diamond (1989) for further details on the Beveridge curve.
16Diminishing returns to inputs in the matching function cause the convex shape.
employed workers find it easier to locate a job, and hence the expected duration of an unemployment spell falls, as does the equilibrium unemployment rate.

**Efficiency and the Optimal Deposit Rate**

The search model features trading externalities, implying that the decentralized market outcome without policy intervention may be below the socially optimal outcome. This section derives the optimal deposit-refund rate in the steady state.

Social welfare is defined as the present discounted value of gross output minus the sum of the social costs of employment (that is, forgone leisure), $zN$, search costs of hiring firms, $\gamma_0 V$, and (gross) investments, $I$:

$$A(t) = \int_0^{\infty} \left[ F(K(t), N(t)) - zN(t) - \gamma_0 \theta(t)(L(t) - N(t)) - I(t) \right] e^{-rt} dt,$$

where $N = (1 - \mu)L$ denotes aggregate employment, $K = \sum k_i$ is the aggregate capital stock, and $V = \theta(L - N)$ denotes aggregate vacancies. The social planner maximizes equation (16), subject to the employment constraint $N = g(\theta)\theta(L - N) - sN$, and the capital accumulation constraint, $\dot{K} = I - \delta K$. The government's problem can be solved using Pontryagin's maximum principle, which yields the following optimality conditions:

$$F_k(\hat{k}) - (\delta + r) = 0,$$

where $\hat{k}$ is the optimal capital stock.
\[ F_N(\hat{k}) - z + \hat{\theta} \gamma_0 - \left[ s + \hat{\theta} g(\hat{\theta}) \right] \lambda = r\lambda - \lambda, \quad (18) \]

\[ \lambda g(\hat{\theta}) [1 - \eta(\hat{\theta})] - \gamma_0 = 0. \quad (19) \]

where \( \lambda \) is the co-state variable of the employment constraint, \( \hat{k} \) is the firm's capital-labor ratio, the hats indicate socially optimal values, and \( \eta \) is the (absolute) elasticity of the function \( g(\theta) \). The condition for the capital stock, equation (17), is similar to equation (7), reflecting the optimality of private investment decisions. Equations (18) and (19) can be written to obtain the steady-state social optimum (with \( \lambda = 0 \)):

\[ F_N(\hat{k}) - z = \left[ \eta(\hat{\theta}) g(\hat{\theta}) \hat{\theta} + r + s \right] \gamma_0 = \Gamma^S(\hat{\theta}, \eta) \gamma_0. \quad (20) \]

The steady-state symmetric market condition is as follows:

\[ F_N(\hat{k}) - z + rb = \left[ \beta g(\theta) \theta + r + s \right] \gamma_0 = \Gamma^M(\theta, \beta) \gamma_0. \quad (21) \]

Matching the equations (20) and (21) yields:

\[ rb = \gamma_0 \left[ \Gamma^M(\theta, \beta) - \Gamma^S(\hat{\theta}, \eta) \right]. \quad (22) \]

If the Hosios (1990) condition—saying that \( \eta(\theta) \) should equal \( \beta \)—is satisfied, the social optimum is obtained in which all search externalities are fully internalized. Hence, the government should choose a zero deposit rate, \( \hat{r} = 0 \), so that \( \hat{\theta} = \theta \).

However, if \( \eta \) is smaller than \( \beta \), implying that firms at the margin cause fewer spillovers to other firms than workers cause to other employees, equilibrium

17For a Cobb-Douglas matching function it is easy to show that \( \hat{\eta}/\hat{\theta} > 0 \), \( \hat{\eta}/\theta > 0 \), and \( \hat{rb}/\theta = \hat{\eta}/\theta > 0 \). In this case, \( \eta \) is a constant so that optimality of the market equilibrium is a knife-edge property.
unemployment is above the efficient level. In that case, a flow subsidy at the rate $rb$ is needed to yield a socially optimal outcome. Conversely, if $\eta$ exceeds $\beta$, employment should be taxed. In practice, however, it is difficult to determine whether employment should be subsidized or taxed; it all depends on the properties of the matching function—which is hard to estimate on the basis of poorly measured data—and stylized assumptions on the behavior of workers.

III. Discussion\footnote{The discussion is kept informal so as to study the features of deposit-refund schemes in a wider context than the rather stylized theoretical model.} \footnote{A generalized application of the deposit-refund scheme would require a one-off lump-sum tax at introduction to finance ex post disbursement of employment subsidies on the current stock of workers. Applying the scheme at the margin to targeted groups significantly saves on the budgetary outlays needed to set up the system compared with an across-the-board application.}

As was pointed out earlier, deposit-refund systems may provide the same economic incentives as pure taxes and subsidies on labor. This is evident from the fact that the “deposit” on a hiring becomes a simple subsidy if an employer decides not to fire an employee. Conversely, an employer that fires an employee—who was hired before the implementation of the scheme—without rehiring a new one has to pay a pure tax. In the steady state, when both hiring and firing takes place, the analysis above has shown that deposit-refund schemes on labor provide for a flow subsidy—from the government to the firm—equal to the interest income earned on the deposit. Of course, if the real rate of interest were zero, the net subsidy to the firm would be zero. At this stage, it should be recognized that the qualitative allocation effects of other fiscal intervention rules such as cuts in unemployment benefits and job-finding bonuses are similar to the proposed deposit-refund scheme; they all reduce the equilibrium rate of unemployment.

How do deposit-refund schemes differ from the aforementioned instruments? Bohm (1981) shows that environmentally motivated deposit-refund systems may provide for stronger or more focused incentives than taxes/subsidies. For example, if beverage containers have been littered, and thus the refund has not been collected on them, someone else may take care of it and claim a refund, which would not happen with a simple fee. Bottles do not necessarily need to be returned to the same store, which increases the flexibility of the scheme. A similar argument applies to deposit-refund schemes on labor; firm B may hire an employee who was fired by firm A so as to cash in the “deposit.” Intuitively, the key incentive—increasing employment—is transferred to any firm wanting to hire new people.

Besides providing similar incentives as under taxes and subsidies, deposit-refund schemes in the labor and environmental field avoid some of the disadvantages of these alternatives. The budgetary effects of a deposit-refund may be more attractive to policymakers than pure subsidies. Subsidies create a need for additional government funds, which often need to be raised by distortionary taxes, whereas deposit-refund systems leave the budget intact. Once the funds are raised to establish the deposit-refund system,\footnote{A generalized application of the deposit-refund scheme would require a one-off lump-sum tax at introduction to finance ex post disbursement of employment subsidies on the current stock of workers. Applying the scheme at the margin to targeted groups significantly saves on the budgetary outlays needed to set up the system compared with an across-the-board application.} it will operate on a budgetary neutral
basis in the steady state—given the assumptions set out in the theoretical section—because the inflow of workers matches the outflow. Such schemes fail, however, to be budgetary neutral in steady state when firms go bankrupt: there is no way to recoup the revenues from firing taxes on the workforce laid off. Therefore, in transition economies featuring a large share of restructuring firms, these schemes have limited practicability as an instrument to stimulate employment. Typically, the scheme could be applied to sectors of advanced economies in which the government considers the rate of labor turnover to be "socially excessive" to motivate employers to retain and train their workers. Some industries facing very uncertain final demand patterns—mirrored in their labor demand—may need to be exempted from the scheme.

Bohm (1981) also argues that deposit-refund systems in the environmental field may provide for lower information and enforcement costs. Indeed, in a deposit-refund system on beverage containers the buyer has an incentive to prove that the commodity has been properly disposed of so as to claim the refund. Under a system of taxes, however, the buyer may try to circumvent the tax by disposing of the containers in an improper fashion. A major concern with a deposit-refund scheme on labor relates to its administrative feasibility. The scheme is easily subject to fraud: it can only become operational under the strong assumption that the government can distinguish quits from layoffs. Otherwise, a group of workers and employers can get together at the expense of the government. For example, an employee working at firm A quits and is subsequently rehired by firm B, where an employee quits to be rehired by firm A. Even if quits could be distinguished from layoffs, the government needs to set up a monitoring system to ensure that firms laying off people make their deposit repayments. Monitoring costs are likely to remain limited when the number of participating firms in the scheme is not too large. In addition, an effective penalty system must be implemented to enforce compliance of firms with firing tax payments.

Environmentally motivated deposit-refund schemes assume that the commodity eventually needs to be disposed of in some form. In the case of labor this is less clear, because labor may leave due factors beyond the control of the employer. First, if mandatory retirement applies, firms employing workers up to the retirement age could keep the subsidies received in the past on currently retiring employees. In this way, hiring old (but less productive) workers may be promoted. Under a flexible retirement arrangement, however, firms may have an incentive to push old, unproductive workers into early retirement schemes to escape the burden of the firing tax. Second, a lot of workers exit jobs through voluntary quits, from which was abstracted in the model of Section I. If the worker terminates the employment contract, firms enjoy a larger net benefit—they receive a hiring subsidy without having to pay the firing tax—at the cost of higher budgetary outlays to the government. Allowing for voluntary separations in the

---

20In practice, in evaluating the budgetary implications, the government should also take into account the saved unemployment benefits that it otherwise would have provided to unemployed workers and the lost resources on workers who voluntarily quit.

21Unless it is also the government's objective to penalize quits. This seems to be the aim of the scheme applied by Tilburg University.
model creates a moral hazard problem: firms may abuse workers to induce them to quit (or even perhaps ask them to quit and share the gain with them) so that they do not have to pay the firing tax. However, the very existence of firing taxes may also give rise to moral hazard for employers: workers are more likely to shirk because firing costs make it more costly for firms to fire employees.

Besides these considerations, there are a number of other aspects, related to the practical relevance of the deposit-refund scheme in general, that merit further discussion. Opponents of active labor market policies argue that subsidy-tax schemes on labor lead to rigidities in the labor market. An argument against job security provisions (that is, a firing tax) is that firms are likely to freeze the average size of their workforce over the business cycle—through varying working hours—to avoid firing costs during economic downturns. This may delay or even cancel the sometimes needed efficiency-enhancing restructuring in enterprises. If firms cannot vary working hours, the deposit-refund scheme may exacerbate the variation of employment over the business cycle rather than dampening it (as is the case with a system of only firing costs). New and growing firms are expected to expand employment more than without the scheme because they receive the current subsidy payments on new workers whereas the expected firing costs in the future are discounted at a positive rate. Firing costs will not discourage hiring as long as the subsidy payments exceed the expected present discounted value of future firing costs. Restructuring firms, however, face immediate (undiscounted) firing costs if they cannot vary average working hours.

The theoretical model assumed equal hiring subsidies and firing taxes. This makes perfect sense for commodities with a high rate of turnover such as bottles and car batteries. However, if turnover is low—in the case of labor, a worker may be employed at a firm for more than 35 years—and deposit rates are not indexed for inflation, it is not a priori clear whether this is an optimal strategy. In an inflationary environment, firms receive a larger net real benefit when a match dissolves (the real value of the firing penalty is below the real value of the hiring subsidy) at increased net real costs to the government. Observers may claim that this provides a rationale—in addition to voluntary quits and retirement—for a firing tax greater than the hiring subsidy. In general, the following rule should apply: the difference between the firing tax and the job creation subsidy must be greater than or equal to zero; otherwise, the firm could earn unbounded returns by choosing a very high rate of labor turnover.

\footnote{Bertola (1990, 1992) uses a simple dynamic (partial) equilibrium model to show the stabilizing effect of firing costs on the average level of employment. Also, Abraham and Houseman (1993) find that German companies—which typically face stronger employment protection policies than U.S. enterprises—rely much more on the adjustment of average work hours, including the use of part-time work, to reduce total labor input during recessions.}

\footnote{Even though the optimal deposit rate could be determined in theory, ascertaining the optimal rate in practice may not be an easy task because estimates of the net marginal social costs of unemployment are not readily available.}
IV. Conclusion

Firms firing workers impose costs upon society in terms of unemployment benefits, retraining, and other labor market measures. The formal literature on active labor market policies has suggested using (marginal) employment subsidies or hiring subsidies as one way to deal with the problem of unemployment. This paper explores a new policy strategy—a subsidy-tax scheme on labor—as an alternative to budgetary costly employment subsidy schemes. It is shown—using a simple search-theoretic model of the labor market—that if a firm pays a tax when it fires a worker to be reimbursed when it (re)hires that or another worker, the economy-wide wage goes up, the natural rate of unemployment declines, and the vacancy rate increases. Intuitively, the interest earned on the deposit during the time the firm employs the worker acts as an implicit subsidy on hiring a new worker.

Theoretically, a deposit-refund scheme seems to be a useful policy instrument to reduce the equilibrium rate of unemployment, but there are some limitations to its appeal. First, it could be argued that it creates additional rigidities in the labor market because employers may want to stabilize their average labor force—possibly through varying the average hours worked per person—out of fear for firing costs during downturns. Consequently, efficiency-enhancing restructuring in firms through massive layoffs may be delayed. However, the scheme could potentially be implemented at the margin—applying only to new recruits above a certain threshold—in sectors or employment categories where governments want to motivate employers to retain and train their employees. Second, it could create a moral hazard: firms experiencing bad business conditions may try to push employees into voluntarily quitting rather than firing them. However, employees with permanent contracts may be encouraged to shirk more as firing them becomes more costly to the firm. Third, there is no way of recouping firing taxes from firms going through bankruptcy procedures. Finally, the deposit-refund scheme is susceptible to fraud; because the government is unable to distinguish quits from layoffs, firms and employees may get together at the expense of the government. Even if the government could distinguish quits from layoffs, the scheme is administratively demanding because a monitoring plus penalty system needs to be in place to enforce compliance with firing tax payments.

APPENDIX

This appendix derives the analytical results of an increase in the deposit rate. Loglinearizing equations (8) and (15), holding the rate of interest constant, yields the following system of equations:

\[
\left[ \eta(\bar{\theta})(w - rb - F_N) - \frac{1}{\beta \gamma_0} \right] \left[ \frac{\bar{\theta}}{d\theta} \right] = \left[ \begin{array}{c} -1 \\ 1 \end{array} \right] rdb, \tag{A.1}
\]

where \( \bar{\theta} = d\theta/\theta \) and \( \eta(\bar{\theta}) = G_U / f(\bar{\theta}) \) is the absolute value of the elasticity of the \( g(\bar{\theta}) \) function (with \( 0 < \eta(\bar{\theta}) < 1 \)). Solving for \( \bar{\theta} \) and \( d\bar{\theta} \) yields the following expressions:
The vacancy-unemployment ratio and wages both rise as a result of the increase in the deposit rate. By loglinearizing the Beveridge curve (equation (2)), using $\theta = V/U$, $u = s/(s + f)$, and $f(\theta) = \theta g(\theta)$, the following equation is obtained:

$$\ddot{v} = \left(\frac{1}{1 - \eta(\theta)}\right) \dot{s} - \left(\frac{s + f(\theta)\eta(\theta)}{f(\theta)(1 - \eta(\theta))}\right) \ddot{u},$$

(A.4)

where $\dot{v} \equiv dv/v$, $\ddot{u} \equiv d^{2}u$, and $\dot{s} \equiv ds/s$. The Beveridge curve is downward sloping because $0 < \eta(\theta) < 1$. Substituting equation (A.2) in (A.4) and setting $\dot{s} = 0$ yields:

$$\ddot{v} = \frac{s + f(\theta)\eta(\theta)(1 - \beta)rb}{(f(\theta) + s)[\eta(\theta)(F_{N} + rb - w) + \beta \gamma_{0}]} > 0,$$

(A.5)

so that a rise in the deposit rate increases the vacancy rate.

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dots (...) indicate that the data are not available:

a dash (—) indicates that the figure is zero or less than half the final digit shown, or that the item does not exist:

a single dot (.) indicates decimals:

a comma (,) separates thousands and millions:

"billion" means a thousand million; and "trillion" means a thousand billion:

a short dash (—) is used between years or months (for example, 1998–99 or January–June) to indicate a total of the years or months inclusive of the beginning and ending years or months:

a slash (/) is used between years (for example, 1998/99) to indicate a fiscal year or a crop year; and

components of tables may not add to totals shown because of rounding.

The term "country," as used in this publication, may not refer to a territorial entity that is a state as understood by international law and practice; the term may also cover some territorial entities that are not states but for which statistical data are maintained and provided internationally on a separate and independent basis.

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