Sudden Stops and Currency Crises:
The Periphery in the Classical Gold Standard

Luis Catão*
Research Department
International Monetary Fund

First Draft: November 2004

Abstract

While the pre-1914 gold standard is typically viewed as a successful system of fixed exchange rates, several countries in the system's periphery experienced dramatic exchange rate adjustments. This paper relates the phenomenon to a combination of sudden stops in international capital flows with domestic financial imperfections that heightened the pro-cyclicality of the monetary transmission mechanism. It is shown that while all net capital importers during the period occasionally faced such "sudden stops", the higher elasticity of monetary expansion to capital inflows and disincentives to reserve accumulation in a subset of these countries made them more prone to currency crashes.

JEL Classification: E44, F31, F34

Keywords: Gold Standard, Exchange Rates, International Capital Flows, Currency Crises.

---

* Author's address: Research Department, International Monetary Fund, 700 19th street, Washington DC 20431. Email: lcatao@imf.org. I thank George Kostelenos, Pedro Lains, Agustín Llona, Leandro Prados, Irving Stone, Gail Triner, and Jeffrey Williamson for kindly sharing with me their unpublished data. The views expressed here are the author's alone and do not necessarily represent those of the IMF.

I. Introduction

The pre-1914 gold standard is often depicted as a singularly successful international system in that it managed to reconcile parity stability among main international currencies with both rapid and differential growth across nations and unprecedented capital market integration. Nevertheless, several countries in the periphery of the system experienced dramatic bouts of currency instability. Figure 1 illustrates the phenomenon by plotting the various national exchange rates defined as the price of the domestic currency against the British pound, then the main international currency which kept a fixed parity to gold throughout 1870-1913.[1] Among Latin American economies, not only did the exchange rate displayed a long-term depreciating trend but also witnessed large discrete downward adjustments in the mid- to late 1870s and between the late 1880s and early 1890s, with relative stability spells in between.[2] Figure 1 also shows that large discrete currency adjustments were not a Latin American monopoly. Besides countries on a silver standard which were directly exposed to the large fluctuations in the gold price of silver from the early 1870s onwards,[3] several European countries under paper-money regimes also witnessed non-trivial fluctuations in the gold parity of their currencies. Although relatively mild in Austria-Hungary and Italy, large depreciations are observed in Russia in 1875-78, Greece in both 1884-86 and 1890-95, Spain in 1890-93 and 1895-98, and Portugal in 1891-94, and in the latter case after having successively stuck to gold for over three decades.[4]

---

[1] To facilitate comparison, all series are rebased to 1900=100. Defining the exchange rate as the foreign price of domestic currency implies that a rise in the index corresponds to an appreciation of the respective national currency relative to the pound.

[2] Argentina stabilized its exchange rate and held on to a gold peg in 1870-75, 1883-84, and 1899-1913. Brazil was on gold for a few months between 1888 and 1889 and from 1906 to 1913. Chile was on a bimetalic standard through 1879 and on gold between 1895 and 1898.

[3] On the experience of silver standard countries and the correlation between world silver prices and the nominal exchange rates of countries on silver, see Catão and Solomou (2004). Reasons behind the variations in pre-1914 gold price of silver are extensively studied in Friedman (1992), Flandreau (1996), and Redish (1997).

[4] Portugal adopted the gold standard in 1856 and abandoned it in 1891. Spain was on a bimetalic standard through 1882 and never formally joined gold before world war I. Russia joined gold in 1897, while Greece was very briefly on gold in 1885 and then again in 1910-1913.

That some countries did not consistently stick to gold in the 1870-1914 era is especially puzzling in light of several pieces of historical evidence. One is the significant reduction in country spreads (40 to 60 basis points) typically associated with gold standard membership (Bordo and Rockoff, 1996; Obstfeld and Taylor, 2003) – clearly a non-neglible gain in an era of extensive international borrowing and in which several national governments endured high debt-to-GDP ratios. Inability to stick to gold was also at odds with the conventional wisdow in financial matters at the time and with the repeated statements by policy makers in those countries and elsewhere that keeping the exchange rate at par remained a central policy goal.[5] Further, even if words do not necessarily translate into deeds, the economic history of the period plainly indicates that such currency movements had deleterious side-effects. Since most countries borrowed extensively abroad but only a handful of them was able to issue domestic currency-denominated foreign debt (see Bordo et al., 2003), the resulting currency mismatch often entailed repayment difficulties and outright defaults in the wake of depreciations;[6] this in turn led to monetary crunches that triggered banking crises and macroeconomic downturns, clearly offsetting the benefits of a weaker currency on aggregate demand via the trade balance. Last but not least, as nominal rigidities were non-trivial even in this early period (Triffin, 1964; Lewis, 1978; Gordon, 1983), the rising inflation resulting from depreciations and the attendant relative price shifts away from non-tradable sectors likely had regressive effects on income distribution, as insightfully argued in Furtado (1963).

Thus, a question of central interest to both the economic historiography of the period as well as the more contemporary debate on the genesis of currency crises is what drove such large deviations from parity. Striking as it is given the voluminous

[5] On the political economy debate surrounding the desirability of maintaining gold convertibility in the various peripheral countries, see for instance della Paolera and Taylor (2001) on Argentina, Fritsch (1988) on Brazil, and Llona-Rodriguez (2001) on Chile.

[6] While more systematic empirical work on the relationship between currency swings and debt repayment problems during the 1870-1913 is needed, indications of a positive relationship can be gleaned, for instance, from the experiences of Argentina and Brazil (Fishlow, 1989) as well as for countries such as Colombia (where default followed the currency slide of the late 1870s), Greece (at several points), and Portugal (where default followed its exit from gold in 1891). See Lindert and Morton (1989), Kelly (1998), and Beim and Calomiris (2001) for a chronology of default events in the pre-1914 era.

literature on the gold standard, there is little systematic cross-country evidence on what determined the exchange rate in countries that were not strictly pegged to either gold or silver. The existing literature on the topic comprises two strands. First, there are a handful of individual country econometric studies using reduced-form regressions (Cardoso, 1980; Mata, 1987; Fishlow, 1989; Fratianni and Spinelli, 1984; Lazaretou, 1995; Tattara, 2001). One difficulty with deriving consistent inferences from these studies is that they are not based on explicitly laid-out theoretical specifications that nest alternative hypotheses;[7] as such, the robustness of the estimates and a comparison with the results obtained for other countries are hard to gauge.

The second strand of this literature comprises non-econometric work on the possible explanations for the respective countries' incapacity to stick to gold. Two polar views then stand out. At one extreme is the view that unorthodox policy behavior was key – notably, those associated with fiscal profligacy. With important nuances, this argument is central to della Paolera and Taylor's (2001) account of Argentina's pre-1930 monetary experience, and also underlies a number of other country specific studies (see, e.g, Lazaretou, 1993, and Aceña and Reis, 2000). While this literature acknowledge a more subtle mapping between group-specific interests and macroeconomic policies than that assumed in early contributions (e.g. de Cecco, 1974), the emphasis clearly lies on the side of country specificity: discretionary domestic policies are often seeing as the central deterrant to maintaining gold convertibility at the pre-established par rate.

At the other extreme lies the view that factors pertaining to asymmetries in international adjustment to the business cycle were crucial (Ford, 1962; Furtado, 1963; Triffin, 1964). This literature points to the fact that, as booming conditions in capital importing peripheral economies drained foreign exchange reserves in the

---

[7] The only partial exception is Cardoso's study on 19th century Brazil where a general equilibrium of model of exchange rate determination is fully laid out and its reduced-form version is estimated by the instrumental variable method. A discussion of her findings and estimation issues is provided later in the paper.

"core" gold standard countries,[8] interest rates tended to rise to reverse this drainage; while part of such monetary tightening was market induced, it was also reinforced by the behavior of main central banks playing by the "rules of the game".[9] On the periphery side, the absence of a sufficiently strong domestic financial system or a domestic monetary authority with sufficient international reserves and institutional clout lowered the elasticity of capital flows to domestic interest rates, implying that capital flows asymmetrically responded to monetary tightening between the two sides of the system. As a result of this asymmetry, monetary tightening in the periphery tended to be an ineffective instrument of balance of payments adjustment, leaving to the exchange rate the burden of balance of payments adjustment. Such a reliance on parity changes would be the more inevitable the greater the correlation between capital flow and terms of trade in the periphery (Cairncross, 1953). Three testable implications thus follow from this view: first, that exchange rates in much of the periphery would tend to be correlated with swings in international capital flows and net barter terms of trade; second, to the extent that the latter two variables were correlated across countries, exchange rates would also tend to co-move across these countries; third, the role of country specific policies in accounting for exchange rate fluctuations in the periphery is not as important as that of international factors.

This paper takes a new look at the issue. Relative to the literature reviewed above, the analysis has three novelties. One is to use a more structured macro model that combines the monetary and asset view approaches to nominal exchange rate determination with allowance for fundamentals-driven changes in the equilibrium real exchange rate. A key link highlighted in the model is that between nominal exchange rate variance and "sudden stops" in international capital flows, and the role of domestic financial market imperfections in amplifying this transmission mechanism. It will be argued that in countries where the financial system appears to have been

[8] While the definition of a gold standard "core" is not entirely consensual among researchers (Flandreau and Jobst, 2004), it is widely touted that main capital exporters such as Britain, France, and Germany in late 19th century were the key players.

[9] For a concise description of central bank behavior under the "rules of the game", see Eichengreen and Flandreau (1998). Both volumes also contain a detailed discussion of the extent to which this stylized behavior was inconsistent with actual pre-1914 central bank practice, as contended by Bloomfield (1959).

shallower and less well-regulated, and where financial imperfections entailed closer ties between the external provision of liquidity, collateral values and bank lending, the effects of external shocks on exchange rates tended to be dramatic.

A second contribution is to apply the same model to the experience of several countries. To this end, the estimation is based on a cross-country dataset spanning similar macroeconomic and financial indicators over a similar time period for eight countries – Argentina, Brazil, Chile, Greece, Italy, Portugal, Spain, and Russia. While this subset of existing paper money regimes is dictated by data availability, it turns out to be quite representative in GDP weighted-terms since it includes some of the largest peripheral economies at the time. By using an analytical framework and dataset that are consistent across countries, the individual country results are therefore directly comparable.

The third contribution of the paper lies on the application of distinct econometric techniques to this historical data. On the one hand, I use single equation and panel methods based on the estimation of autoregressive distributed lag (ARDL) models to extract the short- and long-run coefficients underpinning the exchange rate response to the various variables. One main advantage of this approach is that of not imposing the homogeneity restrictions of standard panel data methods, which may not hold across heterogeneous country groups with arguably important structural and policy specificities. On the other hand, because the ARDL framework models currency movements in general without distinguishing those associated with "currency crises", I complement the ARDL regressions with qualitative choice estimation in probit regressions. As will be seen below, using the two approaches in a complementary fashion will help discern the various country specific and common international forces driving large currency adjustments in the periphery of the gold standard.

The main findings are as follows. Nominal deviations from fixed gold parities can be overwhelmingly explained by three variables – external terms of trade variations, relative money growth, and fluctuations in capital outflows from the core capital exporting economies. A corolary of this result is that both short-run PPP and the random walk model of exchange rate changes are rejected for all countries studied.

The finding that gross international capital flows was an important driver of exchange rate swings in the various countries also helps explain the common timing of depreciations and gold resumptions across the periphery, since international capital flows themselves displayed a conspicuous common component across countries and regions. Moreover, as capital inflows generally boomed when country spreads were the lowest and flowed out when spreads were the highest suggests that exogenous shifts in lenders' supply schedule dominated the individual country demand "pull". The flip side of this is that uncovered interest parity (UIP) generally failed to hold, implying that interest rate differentials in favor of the periphery were of little help in preventing sharp capital flow reversals. Combined with the finding that pre-1914 capital flows were highly responsive to changes in central bank rates in the gold standard core (Einchengreen, 1992), this finding about risk-unadjusted UIP deviations are consistent with the international asymmetry hypothesis discussed above; that is, domestic monetary tightening in response to sudden stops had limited effectiveness in defending peripheral pegs, entailing that nominal parity changes had to shoulder the burden of balance of payments adjustment under those circumstances.

Finally, and as an important qualification to the view that international asymmetries were paramount in explaining the vulnerability of peripheral currencies to international shocks, there were important national differences in the monetary responses to the external stimuli and to domestic output expansion. These differences are shown to account for much of the discrepancy in the magnitude of exchange rate fluctuations across the countries in the sample, as well as between the peripheral countries that resorted to floating and those which remained in gold throughout despite facing similarly large capital account shocks. In this sense, the evidence presented in this paper brings back to fore Whale's (1937) and Ford's (1962) earlier views on the importance of the domestic financial imperfections in propagating exogenous shocks and thus accounting for differences in how the gold standard adjustment worked in different countries.

The discussion proceeds as follows. Section II lays out the exchange rate model that underpins the subsequent analysis. Section III highlights the stylized facts about capital flow and terms of trade shocks during 1870-1914, and then identify the commonalities in exchange rate and capital flow fluctuations in the different countries. Section IV delves into the relationship between fluctuations in external liquidity and domestic monetary expansion, in light of the financial imperfections framework derived from the modern literature on the monetary transmission mechanism a la Bernanke and Blinder (1988) and Bernanke, Gertler and Gilchrist (1996), as well as of what we know from the historical literature on the various countries. Section V then presents the econometric estimates of the exchange rate model and discusses the roles of external shocks and domestic financial imperfections in explaining the observed bouts of currency instability across countries and over time. Section VI concludes.

II. Exchange Rate Determination

The proposed model of exchange rate determination is a simple extension of Frankel's (1979) synthesis of the traditional flex-price monetary approach to balance of payments (MABP) and the fix-price with overshooting model á la Dornbush (1975). The model is extended to allow for a time-varying country risk premium which is a function of the international supply of liquidity, as in Jeanne and Rose (2002). It also relaxes the assumption that long-run purchasing power parity (PPP) must hold, allowing instead the real equilibrium exchange to be driven by terms of trade trends and long-run productivity differentials between home and abroad, consistent with a wide class of open economy macro models (see Obstfeld and Rogoff, 1996, for a survey and further references).

A salient feature of the model is the distinction between short- and long-run. In the short-run, prices are sticky and risk-adjusted uncovered interest parity (UIP) holds so that (the log of) the expected change in the spot exchange rate (Δe^{e}_{t+1}) responds to the nominal short-term interest rate differential between onshore (i_t) and offshore (i^*_t) asset markets adjusted by a time-varying country risk premium (δ_t) which is an inverse function of the supply of foreign finance:

$$\Delta e^{e}_{t+1} = i^*_t - i_t + \gamma\delta_t \tag{1}$$

where positive values for Δe_{t+1} denote an appreciation between t and t+1, and $0 < \gamma \leq 1$.[10] Thus, when the country risk premium rises (or equivalently foreign finance supply drops), and in the absence of arbitrage failures, a fixed exchange ($\Delta e_{t+1} = 0$) can only be expected if the short-run domestic interest rate rises relative to the foreign rate. Otherwise, the exchange rate would "overshoot" at time t, depreciating beyond its long-run equilibrium; only in this case could the home country investor be compensated for the rise in country risk by an expected currency appreciation between t and t+1. By the same asset market equilibrium condition, the only way that a rise in the foreign interest rate would be compatible with a fixed exchange rate under unchanged domestic interest rates is if the country risk premium drops: only in this case would equilibrium be maintained without entailing an expected exchange rate change.

In the long-run, the exchange rate has to converge to a level which is consistent with goods market equilibrium. As is standard in the literature, the adjustment process is expected to gradually eliminate differences between actual and equilibrium exchange rate levels as well as differences of expected inflation between home and abroad:

---

[10] The risk premium is thus being broadly defined to include bank failure or any confiscation risk that affects the representative asset holding.

$$\Delta e^{e}_{t+1} = -\theta(e_t - \overline{e}_t) + (\pi^{e}_{t} - \pi^{e*}_{t}) . \tag{2}$$

where $\overline{e}$ is the log of the long-run equilibrium nominal exchange rate, and $0 < \theta < 1$.

Combining (1) and (2) yields:

$$e_t = \overline{e}_t + \frac{1}{\theta}[(i_t - i^{*}_{t}) - (\pi^{e}_{t} - \pi^{e*}_{t})] - \frac{\gamma}{\theta}\delta_t . \tag{3}$$

Equation (3) postulates that at any given point in time the nominal exchange rate moves along its long-run equilibrium level, adjusted upwards (downwards) by any positive (negative) short-term real interest rate differential between home and abroad, and by any decline (increase) in the country risk premium.

The second pillar in the model hinges on what determines $\overline{e}$. In the long-run, PPP may not necessarily hold so that the equilibrium real exchange rate is possibly time-varying:

$$\overline{rer}_t = \overline{e}_t + \overline{p}^{*}_{t} - \overline{p}_t \tag{4}$$

where p and p^* are the long-run levels of the domestic and foreign price indices, respectively.

Given long-run price flexibility, the standard log linear version of the Keynes-Hicks money market equilibrium condition allows us to map these price levels onto the behavior of money supply and output aggregates:[11]

$$\begin{aligned} \overline{p}_t &= \overline{m}_t - \phi\overline{y}_t + \lambda\overline{i}_t, \\ \overline{p}^{*}_{t} &= \overline{m}^{*}_{t} - \phi\overline{y}^{*}_{t} + \lambda\overline{i}^{*}_{t} \end{aligned} \tag{5}$$

[11] Similar formulations of the money market equilibrium condition can also be derived from optimizing representative agent models with money in utility function or with cash-in-advance constraints, as in Stockman (1980) and Lucas (1982).

where y, m are the logs of the domestic real output and money stock, respectively, and λ is the interest semi-elasticity of money demand, with ϕ>0. Identical output and interest rate elasticities for the home and the foreign countries are assumed to conserve on notation but are not an essential feature of the model and can be easily relaxed, as discussed later.

Once in long-run equilibrium $\Delta e_{t+1} = 0$ and $\delta_t = 0$, it follows from equations (1) and (2) that $\bar{i} - \bar{i}^* = \pi^e - \pi^{e*}$. Substituting into (5) and then back into (4) yields:

$$\bar{e}_t = r\bar{e}r_t - \lambda(\pi_t^e - \pi_t^{e*}) - (\bar{m}_t - \bar{m}_t^*) + \phi(\bar{y}_t - \bar{y}_t^*)\,. \tag{6}$$

Thus, the nominal equilibrium exchange rate is directly proportional to the real equilibrium rate (which is a constant if PPP holds), adjusted for the expected inflation differential and the long-run money-output gap between home and abroad.

The third pillar of the model consists of establishing the determinants of the equilibrium real exchange rates and of long-run inflation. Following much of the theoretical literature on the open economy macroeconomics (see Edwards, 1988; and Obstfeld and Rogoff, 1996 for surveys and further references), we focus on two determinants of equilibrium real exchange rates – the productivity trend differential between home and abroad, and the country's net barter terms of trade.[12] The first seeks to capture the well-known Balassa-Samuelson effect and second follows from the straightforward argument that an improvement in the country's terms of trade should appreciate the real exchange rate insofar as there is home bias in the consumption of tradables. In the case of the expected inflation differential, we also follow the recent macroeconomic literature that relates long-run inflation determination to countries' fiscal performance (Catão and Terrones, 2004). Under a

[12] One other potentially important determinant of equilibrium real exchange rates is the evolution of countries' stock of net foreign assets (Lane and Milesi-Ferreti, 2004). The main difficulty which capturing the net foreign assets effects on the real equilibrium exchange rate is the lack of relevant information for the countries considered during the period. Yet, as seen below, I try to capture this effect by introducing of the external public debt stock variable in the regressions.

commodity standard that delivers long-term price stability we have $\pi_t^{e^*} = 0$, and equation (6) becomes:

$$\overline{e}_t = \mu + \alpha \; tot_t + \eta(\overline{y}_t - \overline{y}_t^* - n_t + n_t^*) - \lambda\psi(1+\frac{1}{\theta})\overline{d}_t - (\overline{m}_t - \overline{m}_t^*) + \phi(\overline{y}_t - \overline{y}_t^*) + \varepsilon_t \qquad (7)$$

where $\overline{y}_t - \overline{y}_t^* - n_t + n_t^*$ stand for the log of the productivity differential (n standing for the respective factor input), tot for the net barter terms of trade, and $\overline{d}$ a relevant indicator of the respective country's fiscal position, such as deficits relative to GDP or the ratio of debt to exports. The parameter ψ captures the impact of the fiscal indicator on long-run price levels, and $\varepsilon_t \sim i.i.d.(0, \sigma_\varepsilon)$ captures possible errors in the way these variables are measured and accurately gauge the equilibrium levels of the real exchange rate and long-run inflation differentials.

Substituting (7) in (3), defining $\beta = \lambda\psi(1+\frac{1}{\theta})$ and following Jeanne and Rose (2002) and setting $\delta_t = -\upsilon\theta b_t / \gamma$, and re-grouping the various terms, we end up with the equation that will be estimated:

$$e_t = \mu + \alpha \; tot_t - \beta\overline{d}_t - (\overline{m}_t - \overline{m}_t^*) + \rho(\overline{y}_t - \overline{y}_t^*) - \varphi(\overline{n}_t - \overline{n}_t^*) + \frac{1}{\theta}[(i_t - i_t^*)] + \upsilon b_t + \varepsilon_t . \qquad (8)$$

Equation (8) states that the nominal exchange rate appreciates (depreciates) as: i) domestic real GDP or productivity rises (falls) relative to abroad; ii) terms of trade shift in favor (against) of the home country; iii) the domestic fiscal position (as captured by the public debt to GDP ratio) improves (deteriorates); iv) domestic money falls (rises) relative to the foreign money supply; v) domestic short-term nominal interest rates rise (fall) relative to the foreign rate; vi) the supply of external liquidity b_t rises (falls).

It is also straightforward to see that equation (8) nests alternative views of exchange rate determination. For instance, if the short-run asset market arbitrage condition does not hold, the coefficient on the interest rate differential term (i-i*) will not be significantly different from zero. This entails a short-run ineffectiveness of monetary tightening (loosening) in preventing expected exchange rate depreciation (appreciation). Also, should relative PPP hold in the long-run and rer be thus constant, we expect that $\alpha = \beta = \varphi = 0$. Finally, should the spot exchange rate follow a random walk with drift, all the slope coefficients in the above equation will not be significantly different from zero once a first-order autoregressive term is included among the regressors.

III. External Shocks

Since the model postulates a clear relationship between exchange rate and two key sources of external shocks – the net barter terms of trade and capital inflows – it is instructive to describe the time pattern of the two variables in the different countries. Moreover, since gross capital flows to any given country should respond to a combination of domestic (or "pull") as well as international (or "push") factors, it is also important to identify any common factor in driving cross-country fluctuations in capital flows; such common factor can be taken as an indicator of the variable b in equation (8), the exogenous liqudity "push" factor that "lifts (or sinks) all boats".

Figure 2 shows that in the free capital mobility world that characterizes that pre-1914 era international capital flows were very volatile.[13] Not only did capital flows from each individual core capital exporter displayed prounounced long swings, but also the aggregate series displayed two major episodes when capital outflows to the rest of the world came to "sudden stops": one in the period 1874-78 and the other in 1890-94: in both occasions, total flows more than halved, thus severely drying up

[13] The figures for the French and the German components were converted to pounds using the respective spot exchange rates. Whether one uses the nominal pound equivalent series or deflate it by the UK WPI makes no difference for the inferences, given the stability of wholesale price deflators for the core countries.

the supply of foreign savings to the periphery. Interestingly, these sudden stops in aggregate capital outflows do not appear to have been driven by the disportionate weight of a single country or region in the series. Figure 3 shows that whether we group it by regions or by type of monetary regime, the sudden stop was pervasive: even countries that abade by fiscal discipline and with more developed financial systems and very low and declining public debt such as the United States were also effect. In fact, in both occasion the peak-to-trough decline in the US inflows by its main supplier of foreign savings (the UK) witnessed a fivefold decline. Clearly, there appears to have been a common cross-country driver in the supply of foreign capital to the periphery was impervious to country specific profiles, including those that are typically deemed key to its capacity to attract foreign investment, such as adherence to gold, per capita income levels, and policy fundamentals. In this sense, the evidence provided here bodes well with the view that "push" factors driving core countries' investment abroad were of paramount importance (Edelstein, 1982; deLong, 1999). This is further corroborated by the correlations between the various country spreads and the alternative core capital outflow indicators: should "pull" factors associated with a rise in the rate of return in the periphery relative to that of the core be the main determinant of such, one would observe a positive correlation between country spreads and capital inflows. In contrast, however, Table 1 shows an inverse correlation which is more consistent with the interpretation that shifts in the supply schedule (which increase outflows concomitantly with the lowering of spreads) were the main factor at play – an inference which is also consistent the positive correlation between world financial integration index and average emerging market spreads reported in Flandreau and Zumer, 2004, p.18). Since a well-documented recurrence that immediately preceeded those sudden stops was the sharp increase in interest rates among the core economies, and particularly orchestrated by the Bank of England in 1873 and 1889, one infers that monetary policy tightening in the core was a key driving force – as indeed has also been the case with the more recent vintage of currency crises (Calvo, Leiderman, and Reinhart, 1996; Eichengreen, 2003).

While the absence of a country breakdown for the French and the German capital outflow series limits more definitive inferences, Figure 4 uses the Stone (1999) UK-based capital export series to reinforce the view that there was significant cross-country sychronicity in the timing and often also in the proximate magnitude of such sudden stops. While the Argentina experience during the Baring crisis of 1890/91 is well known, one notes how equally dramatic were the capital flow contractions in countries as diverse as Brazil, Chile, Greece, Portugal and Spain. Because the Stone UK series as a proxy for overall capital inflows is less representative in Southern European countries as well as in Russia where French and even German capital were very important, Figure 4 also plots the variations in external government borrowing, since in all these countries government borrowing accounted for much of foreign capital inflows at the time. Either way, both indicators tell a similar story.

From the point of view of overall external adjustment, it is important to document the extent to which such capital flow shocks were synchonized with terms of trade shifts. As we know from both the 1930s and the 1980s experiences, whenever large and abrupt capital outflows are exarcerbated by adverse terms of trade, dramatic currency crashes and debt crises tend to ensue (see, e.g., Diaz-Alejandro, 1984). A distinguished strand of the literature on international payments adjustment postulates that terms of trade and international capital flows tend to be systematically correlated. Taking changes in net barter terms of trade as an indicator of relative profitability shifts between any two regions, Cairncross (1953) argues that terms of trade and foreign capital inflows tend to be procyclical, i.e., the more favourable a country's terms of trade, the more capital flows in. Triffin postulates the same positive relationship between capital inflows and terms of trade but working indirectly through interest rate changes in the core economies: as monetary tightening in the core both tends to attract capital back in at the same time as it raises stockage costs, leading to commodity dumping in world markets and thereby a deterioration in terms of trade for commodity producing countries. Albeit distinct, both theories are thus observationally equivalent in postulating a positive relationship between terms of trade and capital flows into the periphery.

Support for either theory, however, is not apparent in the data. Not only are the full sample correlations generally low for all eight periphery countries (Table 2), but it was also the case that terms of trade sometimes moved in the opposite direction as that of capital inflows during the two sudden stop periods of 1874-78 and 1890-94. As shown in Figure 4, all countries but Italy (in 1874-76) experienced a terms of trade improvement between the early to late 1870s, while capital inflows were faltering. During the 1890s Baring crisis, terms of trade either improved as in Chile between 1890 and 1894, Brazil (1891-93) and most notably in Portugal between 1890-93 amidst a major currency and debt crisis, or remained about flat as in Spain and Russia. Only in Argentina was there a juxtaposition of TOT deterioration and capital outflows between 1891 and 1894 but the peso depreciation had already reached a bottom by late 1891 where terms of trade turned down (see Figure 1). Even in the long upswing of 1900-1913 when TOT and capital inflows into the periphery did generally trend up, the relationship was not monotonic as the two variables sometimes displayed short term bumps that were inverse to each other.

In short, while there may well be other factors at play which offset the Caircross and Triffin mechanisms, the fact that terms of trade are far from perfectly correlated with international capital flows in the various countries considered indicates that both variables are not colinear and thus have have an independent role to play in accounting for the degree of currency stability in the different countries. This point is borne out by the econometric evidence to be presented in Section V.

IV. Domestic Financial Imperfections

Two generalizations seem to command concensus in the historical literature regarding financial markets in much of the 19th century periphery. One is that domestic financial markets were relatively illiquid and much shallower relative to those in the core, notably in Britain and the United States. In addition, banks were poorly regulated and information about borrowers' credit history usually hard to get by. While these imperfections were in some cases mitigated by stringent regulations restricting entry into banks, such regulations, on the other hand, curbed competition and hindered efficiency. To these structural features, one added ingredient was the existence of multiple issuing banks and the lack of a central bank holding the monopoly of fiduciary money issues.[14]

Some key implications follow. First, inexistent or illiquid domestic bond markets combined with limited access to bank credit by firms imply that borrowers were typically credit constrained; hence, any outward shift in the external supply of funds would translate into immediate credit growth, *ceteris paribus*. Second, once information about the creditworthiness is hard to obtain, lending becomes more responsive to current collateral values pledged against loans. As formalized by Kyotaki and Moore (1997), this boosts the pro-cyclically of bank lending. Third, inexistent or illiquid secondary bond markets for public bonds implied that governments were highly dependent on domestic banks for non-inflationary finance of deficits, especially when external country spreads are prohibitively high and external capital markets shut off. The likely flip side of this is that expectations of bail-outs during "bad times" are heightened, especially since illiquid domestic markets would make it harder for banks to borrow from the local private sector or sell to others the illiquid (but otherwise solvent) items in their portfolios. Last but least, the decentralization of note issuing rights in a context of limited competition and deficient regulatory practices would arguably tend to boost the pro-cyclicality of domestic money supply to aggregate demand. To the extent that only a small fraction of such expansion of bank liabilities was backed by specie holdings and foreign market intervention by governments was then very limited, any excess of domestic currency relative to gold would tend to depreciate the exchange rate.

[14] On the structural features of the banking system in the various periphery countries, see della Paolera and Taylor (2001) on Argentina; Goldmisth (1986), Haber (1997), and Trinner (2000) on Brazil; Llona Rodriguez (1990) on Chile; Fratianni and Spinelli (1997, ch. 3) on Italy; and Drummond (1976) on Russia.

Figure 5 indicates that countries which experienced currency crashes during the 1870-1914 period were generally the ones that embarked upon rapid monetary expansion which far exceeded that of other macroeconomic aggregates in the run-up to the crisis. Argentina is the most notable case, with the ratio of broad money (M3) to GDP rising from 28 to 70 percent of GDP between 1880-1888, before collapsing back to around 30 percent of GDP in the aftermath of the Barings crash. Brazil, where monetary expansion had been previously curbed by stringent regulations on bank entry, also experienced a unprecedent monetary expansion as a result of new loose regulations on bank issuance (the "Encilhamento") leading to a twofold increase in broad money to GDP between 1889 and 1891. In both countries, the ensuing currency collapse was likewise dramatic. In the case of Brazil also, a comparison with the early sudden stop episode of 1874-1878 is suggestive of the role of monetary expansion in exacerbating currency movements. As Figure 5 shows, while the exchange rate did depreciate when capital inflows faltered in 1874-78, the more moderate rate of money growth appears to help explain why the far smaller magnitude of the currency depreciation relative to that observed during 1889-92.

Data for other countries is likewise suggestive of the significant role of monetary expansion in brewing currency crashes. While the absence of a broad money series for Greece make it difficult to derive inferences for that country, in Chile, Portugal, and (to a less degree) Spain, the elasticity of money supply to income far exceeded unity in the two to three years prior to the 1890-92 depreciations. Conversely, the smoother rate of money growth in Italy between 1882 and 1889 was associated with a milder depreciation in 1890-94.

A key question is what accounts for such rapid monetary expansion. An obvious candidate is fiscal behavior. The latitude of the latter in fuelling the monetary expansion and contribute to currency crises does vary widely, however, across countries. On the one hand, fiscal policy – whether measuring by the ratio of central government spending to revenues or by the ratio of public debt to GDP – was clearly expansionary in Argentina, Chile and Greece,[15] and this expansion has been shown to have fed directly through to faster money growth via attendant changes in base money as the counterpart of domestic credit to governments (della Paolera and Taylor, 2001; Llona Rodriguez, 2000; Lazaretou, 1993). On the other hand, fiscal behavior appears to have been reasonably under control in Brazil, where not only public debt to GDP was far lower than elsewhere, but also the average ratio of public expenditure to revenues barely exceed five percent in the four years before the 1890-92 crisis. Likewise, there is no evidence of a significant fiscal deterioration in Portugal in the run-up to the 1891 depreciation; and while the ratio of public debt to GDP was at a highish 80 percent, it had been high for over two decades while parity was successfully maintained; in fact it only deteriorated in 1892-93 largely as a result of the currency collapse due to "original sin" effects documented in Bordo et al. (2003) and the attendant contraction in domestic activity. Likewise, neither in Italy nor in Russia does one observe significant fiscal deterioration. This mixed evidence about the importance of fiscal deterioration in accounting for rapid monetary expansion and the subsequent currency crises is broadly corroborated by the econometric estimates to be presented in Section V.

A second factor pertains to unsound banking practices and regulatory deficiencies. In several of these countries, not only the specie coverage of issued paper was low – ranging from around 1 percent in Chile in the 1880s to between 15 and 20 percent in Italy during 1880-1900 – but there is also evidence that banks tend to reduce their cash-in-vault to deposit ratio during cyclical upswings. The likely rationale is that the opportunity cost of holding cash greatly increased in those circumstances, as borrowers' balance sheet positions improved and so did the value of collaterals such as land and public bond prices, making new loans a specially attractive business. In Argentina, for instance, this ratio dropped from 24% in 1885

---

[15] Using the ratio of expenditure (G) to tax revenues (T) was chosen instead of the traditional (G-T)/GDP measure because of the well-know deficiencies of GDP data for much of the 19th century periphery. Another reason is that rapid GDP growth can mask the deterioration of a country's fiscal position especially when measured as a ratio to GDP. Yet, this choice does not change the basic inferences of the paper since both indicators tended to move in tandem.

to 15% in the peak of the boom in 1889.[16] At the same time, the specie coverage of broad money also fell, from 12% to 7.5%. A similarly sharp decline in coverage is also observed elsewhere. In Chile, for instance, the drop was marginal from an already low basis (less than one percent), but in Greece the drop was dramatic: the ratio of metalic reserves to currency in circulation was down to 13 percent in 1889 from 40 percent in 1885.[17] Thus, when the respective domestic financial systems had a golden opportunity to increase their international reserve holdings emanating from the boom in foreign capital inflows, the opposite happened. This underinsurance against prospective bad times ahead is all the more striking against the background of a long history of sharp depreciations and bank runs in all three countries. This suggests that bank owners and managers either expected a government bail-out or to be able to shift the bulk of losses of a possible currency depreciation to depositors. Either way, this points to the existence of financial frictions in these economies that made them vulnerable to sudden stops.

To further corroborate this point about the high elasticity of bank money to various (capital flows, fiscal and other demand-related) stimuli, the first panel of Figure 6 plots the M2 multiplier (defined as the ratio of broad to high powered money) in six of the eight non-gold periphery countries comprising our sample.[18] These can then be compared it with the counterpart multipliers for a set of peripheral countries that stuck to gold throughout. As is clear from these plots, not only was the level of the M2 multiplier generally higher in the currency crisis countries, but also the evolution of the multiplier is shown to be much less smooth, displaying for instance a conspicuous cyclical pattern rising during the capital flow/demand boom of the late 1880s and crashing during the early 1890s bust.

[16] Data on banks' specie reserves from della Paolera and Taylor, 2001, p.53.
[17] Computations for Greece, Italy, and Russia were based on data from Lazaretou (1993), Fratianni and Spinello (1984), and Flandreau (2003). Chilean specie reserve data was kindly provided by Agustin Llona-Rodriguez. International reserve data for Brazil is unavailable.
[18] Lack of broad money data for Greece and Portugal precludes us from including the two countries in these calculations.

Finally and in addition to this higher M2 multiplier in the non-gold periphery, shallower financial markets exacerbated the effects of such highly pro-cyclical money growth on the currency. Again with the caveats for the lack of a better indicator, Figure 7 illustrates the differences between two large capital importing countries in this respect. While capital inflows rose to a high 20 or so percent of broad money in the United States in 1887-89, they were twice as high in Argentina. As a result, while both countries were confronted with a massive sudden stop in 1890-92, the shallower financial markets in Argentina entailed a far greater aggregate liquidity shock. In the absence of sufficient international reserves in the domestic banking system or of a central bank capable of mouting a credible interest rate defense of the peg, the burden of external adjustment would then fall squarely on the exchange rate.

V. Econometric Evidence

While instructive, raw data descriptions and simple graphical analyses as above are bound to fall short of capturing the typically rich dynamics of exchange rate adjustment and giving us a sense of the relevant elasticities. To this end, regression analysis is needed. This section presents econometric estimates of the exchange rate model laid out in Section II. Since a salient feature of that model is the separation between short- and long-run determinants of exchange rates, an econometric specification that accommodates this feature is required. In addition, potential reverse causality running from exchange rate changes to changes in budget deficits and monetary aggregates suggests that such a source of simultaneity biases in the estimates should be mitigated. A standard approach to these problems consists of estimating an auto-regressive distributed lag model (ARDL), where dependent and independent variables enter the right-hand side with lags of order p and q, respectively:

$$e_{i,t} = \mu_i + \sum_{j=1}^{p} \lambda_{i,j} e_{i,t-j} + \sum_{l=0}^{q} \boldsymbol{\delta}_{i,l}^{'} \boldsymbol{x}_{i,t-l} + \varepsilon_{i,t} \tag{9}$$

where $e_{i,t}$ stands for the log of the nominal exchange rate of country i at time t; μ_i is an intercept (or fixed effects in a panel as discussed below); $\lambda_{i,l}$ is a scalar, $\boldsymbol{\delta}_{i,l}$ is a (kx1) coefficient vector and $\mathbf{x}_{i,t}$ is a (kx1) vector of explanatory variables which includes e, i.e., $\mathbf{x}_{i,t} = \begin{vmatrix} e_{i,t} \\ \mathbf{x}^*_{i,t} \end{vmatrix}$, and $\mathbf{x}^*_{i,t}$ is a $(k-1,1)$ vector which comprises the explanatory variables in (8). As extensively discussed in the time series econometrics literature (Banerjee, Dolado, Galbraith, and Hendry, 1993; Stock and Watson, 1993), if there exists a long-term equilibrium relationship between the right and left hand side variables in (8) implying that either all or a subset of these variables are cointegrated so that $\varepsilon_{i,t}$ is stationary, then such a relationship can be identified within an ARDL structure such as in (9), provided that p and q are suitably chosen. A further advantage of this methodology is that of dispensing unit root pre-testing, since it can be applied regardless whether the individual regressors are stationary I(0) or contain a unit root and are thus first-order integrated, I(1) series.

The respective long- and short-term coefficients to be estimated can be seen from a re-parametrization of Equation (9):

$$\Delta e_{i,t} = \mu_i + \phi_i[e_{i,t-1} - \boldsymbol{\theta}_i \mathbf{x}_{i,t}] + \sum_{j=1}^{p-1} \lambda^*_{i,j} \Delta e_{i,t-j} + \sum_{l=0}^{q-1} \boldsymbol{\delta}^{*'}_{i,l} \Delta \mathbf{x}^*_{i,t-l} + \varepsilon_{i,t} \qquad (10)$$

where the vector $\boldsymbol{\theta}_i = \begin{vmatrix} \beta_i \\ \boldsymbol{\gamma}_i \end{vmatrix}$ defines the long-run or "equilibrium" relationship between the variables involved, with β_i being the elasticity of the export to import ratio with respect to the REER and $\boldsymbol{\gamma}_i$ is a vector of elasticities with respect to other variables ($\mathbf{x}^*_{i,t}$), whereas ϕ_i measures the speed of adjustment toward equilibrium. The coefficients in $\boldsymbol{\delta}^{*'}_{i,l}$ capture the "impact" or short-run elasticities of the exchange rate with respect to the various left-hand side variables.

Pesaran and Smith (1995) and Pesaran, Shin, and Smith (1999) show that two methods can yield consistent panel estimates when groups (in our case countries) are heterogeneous. One is the so-called mean group (MG) estimator which derives the full panel estimates $\boldsymbol{\theta}$, $\boldsymbol{\delta}^*$ and ϕ as simple averages of individual country coefficients. However, because MG estimates will be inefficient if $\boldsymbol{\theta}_i$ is the same across groups (i.e., if the long-run slope homogeneity restriction holds), Pesaran, Shin, and Smith (1999) also propose a maximum likelihood-based "pooled mean group" (PMG) estimator which combines pooling and averaging of the individual regression coefficients in (10). By allowing the researcher to impose cross-sectional long-run homogeneity restrictions of the form of $\beta_i = \beta$, $\boldsymbol{\gamma}_i = \boldsymbol{\gamma}$, $\forall$ i =1,2,...N, the PMG estimator also has the attractive feature of enabling one to test this restriction via standard Hausman-type tests.

In light of the evidence presented earlier in the paper of significant country heterogeneity of exchange rate adjustments and of the role of the different variables therein, both invidual country and full panel estimates are presented. In all cases, the ARDL order p,q are optimally chosen according to the Schwartz Bayesian criterion (SBC) subject to the constraint that $p,q \leq 2$ so as to conserve on degrees of freedom.

Table 3 reports the pooled mean group estimates for the 8-country panel. Because domestic interest rate indicators are available only for four countries, the i-i* variable had to be excluded from the panel; however, as will be seen below, this omission should not bias the results since individual country estimates shows that the interest rate differential is not a significant explanatory at conventional levels of statistical significance. All the reported coefficients are for the log level of the variables and also consistent with the above discussion, all regressions comprise a common external risk-premium factor proxied by the log of the foreign investment by core countries expressed in deviations from trend ("efico").[19]

---

[19] For all countries with the exception of Russia and Portugal, the indicator "fic" stands for the (log of the) sum of total recorded foreign investment by the UK, France, and Germany denominated in the same currency (Sterling). In the case of Russia, evidence that the overwhelming share of foreign investment in that country was of French origin (Crisp, 1970), I use the log of France's total foreign investment. This is substantiated by the results of Table 1 which indicate a far higher correlation between Russia's sovereign spreads and French foreign investment than between Russia's spreads and UK or Germany foreign investment. For similar reasons, the "fic" indicator stands for the log of the

The first line of results in Table 3 indicates that all the variables included have coefficients with the expected sign. Also consistent with the theoretical model, the elasticity of the exchange rate to the money differential variable (m3-m3*) is negative and about unity, with its high t-ratio indicating that it is very precisely estimated. Other coefficients also yield reasonably high t-ratios with the exception of the fiscal deficit variable (g/t). The estimated error correction coefficient of -0.29 indicates, however, that the speed of adjustment to equilibrium following a shock is somewhat slow, with a half-life of 2 years.[20] To the right of the table I also report the respective Hausman-tests (h) on the long-run slope homogeneity restrictions, which gauge whether the coefficients for each variable are in fact the same across groups. The results show that the homogeneity restrictions cannot be rejected at 5 or 10 percent for all variables with the exception of the money differential variable, for which an h-statistic of 16.41 rejects the restriction with a negligible probability of error (p). This indicates that the model be re-estimated without imposing that restriction. However, since g/t does not show up as significant, it is sensible first to verify whether the rejection of this restriction is not being caused by the undue inclusion of that variable in the regression. As shown in specification II, excluding g/t leaves the remainder estimates essentially unaltered, with the h-test still rejecting the homogeneity restriction on m3-m3*.

Accordingly, specifications III and IV report the estimation results without imposing the cross-country homogeneity restriction on the money differential variable, with and without including the g/t variable. The main differences are the drop across the board in the estimated elasticities and the loss in statistical significance of the income differential variable (y-y*). Yet, all the variables still yield the theoretically postulated signs and the exchange rate elasticity to the money variable is not so far away from unity.

UK total foreign investment. Since Portugal was such a small country in the aggregate UK capital exports, the exogeneity assumption of this regressor with respect to expected changes in Portugal's exchange rate is clearly warranted.

[20] As standard, this is calculated as ln(0.5)/ln(1-|EC|).

Table 4 reports the results of individual country regressions. By not requiring the same set of explanatory variables for all countries as in a panel, individual country regressions allow us to include the interest rate differential variable for the countries for which we have such data. Both the long- and short-run coefficients on a country by country basis are also reported.

Starting with Argentina, the first two lines of results in Table 4 largely reinstate the findings of the panel regressions in Table 3: terms of trade, money growth differential and foreign capital flows are significant determinants of long-run exchange rate fluctuations. The availability domestic interest rate data which allows us to estimate the exchange rate elasticity to the i-i* differential and this happens to be statistically insignificant, thus corroborating the exclusion of this variable in the panel regressions. One difference, however, between the panel and the individual regressions for Argentina has to do with the role of fiscal variable. The log ratio of fiscal expenditures to revenue g/t, albeity not statistically significant at 10 percent, has the expected negative sign and its t-ratio approaches statistical significance in the short-run. In addition, the third line of results indicates that once we replace the g/t variable by a widely used alternative metric of fiscal solvency – the ratio of external public debt to exports – the fiscal variable shows up as highly significant both in the short- and long-run.[21] In all cases, the fit of the regressions is excellent: about 98 percent of exchange rate movements on an annual frequency are explained by terms of trade variations, changing fiscal positions, differential money growth, and the common cross-country foreign investment cycle.

An important question that arises in countries like Argentina, which switched between on- and off-gold regimes in the pre-1914 era (Argentina was gold between 1870 and 1875, 1883-85, and 1899-1913), is whether the model can explain exchange

[21] I also experimented with the ratio of total public debt to GDP with results that were no better than that yielded by the other two indicators. Note that, because observations for the public debt variable are not available before 1883, the estimation period differs between regressions II and III. The shorter estimation period for regression III may well explain also the limiting case of a unit error correction coefficient

rate behavior about equally as well across such regimes. In other words, is there any evidence of a structural change in the estimates due to factors such as shifts in policy credibility that are unaccounted for by the fundamental variables in the model? To gauge this hypothesis, the last column in Table 4 reports the results of F-test on a set of intercept and slope dummy variables defined as 1 for gold standard periods (1883-84 and 1899-1913) and zero otherwise.[22] The low value of the respective F-statistic indicates that, in all specifications considered, there is no evidence of structural changes in the estimates. On the one hand, this suggests that the country's capacity to stabilize the currency and thus sustain a gold peg is fully explained by the model's fundamentals without resort to extraneous credibility factors. This finding has important implications for the existing literature which are further discussed in the conclusion.

Turning to Brazil, the estimates continue to show terms of trade, money differential and (to a mixed extent) core capital exports as strongly significant explanatory factors of exchange rate behavior. In contrast with Argentina, however, the fiscal variable is not only insignificant but also yields the wrong coefficient in the first specification. As discussed in in Section IV, this reflects significant differences in the stance of fiscal policy in the two countries: while in Argentina, fiscal policy appears to have been strongly pro-cyclical and resulting in a much higher level of fiscal debt overall (regardless whether weighted by GDP or by exports) in Brazil fiscal policy was far better behaved. Another salient difference between the two countries lies in the sign of the income differential (y-y*) variable, which in Brazil yields the expected positive and statistically significant coefficient when the g/t variable is appropriately excluded (see specifications II and III in the Table). These results corroborates some of the findings of Cardoso (1983) and Fishlow (1989) on the determinants of nominal exchange rate Brazil. Cardoso finds that external money growth and terms of trade changes (as proxied by the relative external v. internal price differential due to data unavailability at the time) had strong explanatory power

[22] Argentina abandoned gold in March 1885 so by the year's end the exchange was some forty percent below par (della Paolera and Taylor, 2001, pp.48-49). Accordingly, the gold dummy in the regression was set to one in 1883-84 but zero for the fully year of 1885.

in explaining the mil-reis/sterling exchange rate during the period 1870-1906; Fishlow adds capital inflows to this list. Since gross foreign capital flows in Brazil followed a time pattern is very similar to that of overall capital exports from the core countries, it is not surprising that similar inferences obtain. In the absence of a series on domestic interest rate, the variable i-i* had to be omitted.

A new interesting result is that, as with Argentina, the model does not break down between the pre- and the post-1906 period when Brazil joined gold. As shown by the F-tests on the intercept and slope gold dummies reported on the far right of the table, there is no evidence of a structural break between the two periods at any reasonable level of statistical significance. Together with an adjusted R^2 of 0.97, this implies that the model explains rather well the behavior of Brazil's nominal exchange rate during both the gold and off-gold periods.

Estimates for Chile tell a similar story, with terms of trade changes, differential money and population growth, and fluctuations in core capital exports standing as the main drivers of the peso exchange rate. As in Brazil, with fiscal policy being reasonably well-behaved in the medium- and long-run (Llona Rodriguez, 2000), it is not surprising that fiscal indicators do not show up as having a significant effect on the exchange rate. Also in tandem with previous results, the estimated coefficient on the onshore-offshore interest rate differential i-i* is not only statistically insignificant but also has the "wrong" sign, while the income differential variable y-y* continues to flip-flop between the negative and positive terrains being statistically insignificant over the long-term. Bearing in mind all the above mentioned caveats about national income estimates during period, this result however does not appear to be unique to the period or produce by the chosen econometric especification: Frankel and Rose (1995) also note in their survey of the literature that relative real income growth is largely unsuccessful as an explanatory variables in exchange rate regressions using post-war data. Finally, as with Argentina and Brazil, F-tests on the gold dummies for the spell Chile was on gold (1895-1998) does not indicate any structural break at a 5 or 10 percent significance level relative to the off-gold sub-periods.

Relative to the results so far, the estimates for Greece come out as outliers. Terms of trade, the ratio of government expenditure to revenues, and relative income all have the opposite sign predicted by the model. Further, an asymptotic t-ratio of -0.43 for the relative money variable in the first regression for Greece indicates that the long-run effect of money on the exchange rate is very imprecisely estimated, though its short-run effect with a t-ratio of -4.46 is clearly statistically significant at 5 or even at 1 percent. A closer look at the underlying data reveals, however, that the main source of such anomalies is high multicollinearity between the g/t and tot variables. This is hardly surprising since much of the tax revenue base in the pre-1914 periphery centered on foreign trade, thus rendering the g/t ratio as partially endogenous to external terms of trade fluctuations. This can be gauged from the second line of estimates for Greece which, as was done for Argentina, Brazil, and Chile, omits the g/t variable. As expected under strong multicollinearity, the tot coefficient shrinks dramatically and does its t-ratio, while the t-ratio on the money differential variable rises well above the standard critical values and the associated long-run elasticity gets close to unit in accordance with the theoretical model. This result is robust to dropping the y-y* variable from the equation as shown in the subsequent line.

Since Greece is the only country in the dataset for which the money supply variable had to be measured by currency issued due to lack of broad money data, these regressions result provides some qualified support for Lazaretou's (1993) contention that fluctuations in the dracma exchange rate in the pre-1910 period directly stemmed from budget financing through seignorage. Yet, since the fiscal deficit variable (at least as measured by ratio of expenditures to revenues) does not prove significant even when it enters the regression instead of money (and with or without including terms of trade), this hypothesis is not fully substantiated by the statistical evidence. Further, all specifications for Greece indicate that fluctuations in the core capital export cycle ("efic") were a significant and clearly exogenous factor at play in exchange rate determination that thus deserve inclusion in the analysis. Finally, as in the other individual country regressions above, there is no evidence of a structural break in the estimated coefficients between off- and on-gold periods (Greece was on gold in 1885 and 1910-13).

Turning to Italy, all the explanatory variables yield the expected signs, as usual with the exception of interest differential term i-i*. The ratio of fiscal expenditures to revenues does not quite rise to conventional significance thresholds and the associated elasticity is low, but the estimated coefficients on the terms of trade, money differential, n-n* and the efic variables are all generally significant and of sensible magnitudes. This finding that fluctuations in the lira exchange rate did not conform to PPP and are explained by other variables is consistent with the early findings of Fratianni and Spinelli (1984) that reject both absolute and relative PPP in pre-1914 Italy. The significance of the capital flow variable is also consistent with the econometric evidence mustered in Tarttara (2000).[23] There are two striking difference between Table 4 estimates and Tattara's. One is the rather low coefficient of the money variable (which in his study is simply the log of the domestic money stock rather than the domestic-foreign log money differential) which range between 0.09 and 0.006 in his regressions (depending on the specification) and yield much lower t-ratios. The other is that the interest rate differential between the domestic and foreign currency denominated Italian bond (the "rendita") is a key explanatory variable in his regressions, where Table 4 results indicate no significant support for UIP once other fundamentals are included in the regression.[24] Since the inclusion of other variables

[23] Tattara nevertheless uses different metrics for those two effects: the terms of trade change indirectly enter his equation through the differential between domestic and foreign prices and the capital flow variable is the current account balance net of reserve variations. One problem with both indicators is their endogeneity to exchange rate changes, implying that robust inferences will otherwise have to rely on the superconsistency of OLS estimates. In contrast, this reliance is not required in the interpretation of the estimates reported in Table 4, since the external terms of trade and core capital flow indicators can be taken as exogenous for a small open economy such as pre-1914 Italy.

[24] This finding proves to be invariant to all the three measures of interest rate differential available for Italy. Besides the one reported in Table 1 – the differential between the yield of the Italian "rendita" and the yield on the British consol – I also experimented with both the differential between the lira denominated Roma rendita yield and Paris market gold-denominated yield on the same instrument, as

besides the interest rate differential in his econometric model is not formally derived from a common theoretical specification (or, indeed, from any specification derived from first principles), it becomes difficult to gauge what is driving the inference. The fact that his regression model makes no allowance for fiscal variables, terms of trade, and the domestic-foreign productivity differential in driving the equilibrium exchange rate, this may well be the reason. Finally, since gold convertibility never became fully effective during the period (see also Fratianni and Spinelli, 1984), gold dummy F-tests were not conducted.

The results for Portugal are also broadly consistent the model with two differences. One is the coefficient on the terms of trade variable which falls short of statistical significance and actually takes on an unexpected negative sign. It is possible that this reflects the serious measurement errors observed in the Portuguese trade statistics of the period and the remaining imprecisions after attempts to correct them (Lains, 1995). One other possible reason is the combination of extremely low trade openness of the pre-1914 Portuguese economy (Portugal's ratio of exports plus imports to GDP averaged 10 percent during 1870-1913 thus being far lower than any other country studied) and high trade barriers on domestic manufacturing and a price control scheme on a basic staple commodity such as wheat (Reis, 2000, p.99), which likely dampened the effect of external terms of trade changes on domestic relative prices and hence on the exchange rate. The other main difference with the Portuguese estimates is that, in contrast with the remainder countries in the sample, gold dummies are statistically significant, pointing to a structural break between the pre-1891 period when the country was on gold and the 1892-1913 off-gold period. Considering that Portugal was the first European country after Britain to adopt a full-fledged gold standard, that it stuck to this monetary standard for 37 years and witnessed gold circulation accounting for over two-thirds of the domestic money supply (Reis, 1996), the significance of the gold dummy should not come as a

---

well as with the differential between Italy's Banca Nazionale discount rate and an average of the discount rates of the Bank of England, Bank of France, and the German discount rate. The respective t-ratios in the short-run first-difference regressions are 0.11 and -0.98, thus well below any conventional level of statistical significance.

surprise. Re-estimation of the equation for the post-1891 period (not shown in Table 4 but available from the author upon request) corroborates the finding that foreign capital pushes, money, and population growth differentials explain around 95 percent of the currency fluctuations in the off-gold regime, with the income variable no longer showing up as significant, however.

Results for Spain and Russia accord well with what the theoretical model would predict, again with the notable exception of the UIP building block since the interest rate differential variable for both countries falls well short of statistical significance. The Russian estimates in particular show a strong direct effect of the fiscal variable on the exchange rate in contrast with that found for all other countries but Argentina, and the speed of adjustment to disequilibrium is a lot faster – in fact approaching the limiting case of an estimated ECM of minus one in two of the specifications. Interestingly in light of the polemic in the literature about the extent of the credibility of Russian gold standard and the role of the large reserve accumulation in ensuring its sustainability (Drummond, 1976), the econometric estimates reveal no break between the pre-1897 off-gold and the post-1897 on-gold regimes. This suggests that the stability of the ruble in the post-1897 period should be attributed to whatever changes in the fiscal, monetary and external fundamentals took place between the two sub-periods, rather than by large reserve accumulation or by an extraneous improvement in policy credibility not backed up by improving macroeconomic fundamentals.

To sum up, the above econometric estimates indicate that both in the short- and long-run three main variables stand out in accounting for exchange rate fluctuations in the cross-country dataset – namely, net barter terms of trade, differential money supplies and variations in country risk premia stemming from a capital export "push" from the core.[25] In contrast, a variable much touted in the

[25] Because measurements errors are likely higher in pre-1914 data, it seems reasonable to work with 10 percent as the cut-off level of statistical significance rather than the usual 5 percent threshold. Thus, unless otherwise stated, an explanatory variable will be deemed significant once its t- of F-statistic passes that threshold.

theoretical exchange rate literature – the onshore-offshore interest rate differential (i-i*) – is unable to account for exchange rate variations in *all* of the countries for which this measure is available. Results for the relative real income variable (y-y*) are mixed and this may partly reflect the quality of the existing GDP data for some of those countries during the period. Yet, this does not seem a feature unique to the estimation period or to the specification/regression technique chosen: in their survey of the more recent empirical literature, Frankel and Rose (1995) also note that both interest rate differentials and relative real income growth are largely unsuccessful explanatory variables in exchange rate regressions using post-war data. Likewise mixed are the results for the fiscal variable. One possible reason is multicollinearity between the fiscal balance and money. However, such a multicolinearity does not appear to be strong in the proposed specification for two reasons. First, the money aggregate enters the equation in log levels rather than in the log of first differences. Second, for all countries but Greece and Portugal (due to data limitations), broad money is used instead of currency in circulation (Greece) or M1 (Portugal); since currency issued and broad money followed quite distinct paths in several of these countries, and given that governments had growing access to international borrowing in an era of increasing capital mobility, the link between public deficits and currency issued was not one-to-one. Indeed, for all countries we studied with the exception of Argentina dropping money as a regressor does not improve the economic or statistical significance of the fiscal variables. A more plausible reason is that the effects of the fiscal variable are asymmetric: weak fiscal positions increase the likelihood of a currency crisis, whereas strong fiscal positions do not readily translate into an appreciation of the currency. This interpretation is corroborated by the probit estimation below.

As noted in the introduction, in addition to examing what drives currency movements more generally, I also look at what factors change the likelihood large downward currency adjustments. To this end, a discrete choice model is estimated comprising a similar set of explanatory variables as the ARDL regressions above. A main issue in this connection is the definition of "currency crisis" or of "large" exchange rate adjustments. In the absence of a commonly agreed criterion in the literature,[26] I define a currency crisis or crash as an exchange rate depreciation greater than at least one standard deviation of the annual percentage change of the nominal exchange rate (relative to sterling) over the entire 1870-1913 period, and which is not reversed within a three-year window. Thus defined, our sample comprises 18 such events.[27] As in Frenkel and Rose (1996) and many others, the dependent variable is set to 1 in the first year of a crisis episode and zero otherwise, while the observations pertaining to the period during which the crisis is ongoing (i.e. as the exchange continues to slide) are excluded since they are part of the same crisis already counted once.

In addition to providing complementary evidence on what determined gold parity adjustments in the periphery, a discrete choice approach along these lines also allows us to include in the panel countries that never experienced a currency crash during the period. Data availability permits the inclusion of eight such countries – namely, Australia, Canada, Denmark, Finland, New Zealand, Norway, Sweden and the United States. With the exception of the United States which sometimes are classified as part of the core "gold club",[28] these are countries typically deemed as "emerging economies" in the pre-1914 world and so constitute direct counterparts to the non-gold countries featuring in the original sample. By including these contrasting experiences and adding relevant information to what makes a country's currency more or less prone to crashes, this enlargement of the sample helps improve the robustness of the inferences.

---

[26] See Frenkel and Rose (1996), Kaminsky and Reinhart (1999), and Milesi-Ferretti and Razin (2000) for the different classification criteria employed in this literature. One feature of the pre-1914 period which helps minimize disagreements on crisis dates is the more stable price environment and the absence of policy devices such as crawling pegs, which make it harder to separate currency crashes from non-crisis related discrete exchange rate adjustments in response to high inflation bouts.
[27] These are: Argentina (1875, 1885, 1890), Brazil (1876, 1890), Chile (1876, 1885, 1891, 1907), Greece (1886, 1891), Italy (1892), Portugal (1891, 1908), Spain (1892, 1897), Russia (1876, 1891).
[28] See Catão and Solomou (2004) for evidence on some distinctive features of US effective exchange rate during the period compared with the European core. A discussion of what classifies a country as part of the "core" or "periphery" is provided in Flandreau (2004).

Table 5 reports the probit estimates.[29] Starting from an specification that includes both contemporaneous and lagged levels of the explanatory variables, column 1 shows that relative money-output ratio ((m2-m2*)/(y-y*)), the ratio of government expenditure to revenues (g/t), core capital export cycles ("efico") and the UK discount rate ("iuk") are all significant determinants of the likelihood of a currency crisis. In contrast, and somewhat in tandem with the evidence presented earlier, the terms of trade variable (which as in the case of capital exports is measured in terms of deviations from a log-linear trend) does not show up as statistically significant at conventional levels. A pseudo-R^2 of 0.50 and the fact that the model correctly classifies 97% to 98% of the annual observations between crisis and non-crisis events depending of the cutting-off point (scoring a high of 98% once we define a crisis as an estimated probability in excess of 0.5) indicate that the model fits the data extremely well.[30]

The second column of Table 5 further reinforces these results in a specification where all variables with the exception of the government expenditure to revenue ratio enter the regression in first differences which is the alternating signs of the level specification of column (1). As usual with first-difference specifications, the pseudo-R^2 is lower but the model still correctly classifies between 96 and 98% of the observations. No less importantly, with the exception again of the terms of trade and of the UK interest rate which yields an unexpected negative sign (apparently reflecting a strong negative colinearity with the capital flow variable), all variables yield economically and statistically significant coefficients. Moreover, the estimated elasticities at mean are strong: for instance, a 1 percentage point contemporaneous increase in the relative money growth to output growth ratio increase the likelihood of

[29] While the empirical literature on determinants of currency and debt crises sometimes uses logistic distribution rather than a probit one, the choice is ultimately an empirical matter and dependent on which distributional assumption better fits the data. Since for the same set of variables, logit estimation yields a marginally smaller pseudo-likelihood than the probit estimation, the latter was preferred. In many practical application, the two methods yield similar results (see Greene, 2000, chapter 19 for a discussion).

[30] As a benchmark comparison, Frankel and Rose (1996) obtain report full-sample pseudo-R^2s of 0.17 to 0.20 in their probit estimates of currency crises using post-World War II data. Similarly, empirical estimates of the likelihood of debt crises using probit-logit specifications also typically report pseudo-R^2s well below 0.5 (see Catão and Kapur, 2004).

a currency crisis by 5.4 percentage points, whereas a 1 percentage increase in last period's government expenditure to revenues raises the crisis probability by 1.5 percentage points.

As usual, the inclusion of contemporary values of the explanatory variables in the regression raises the question of simultaneity biases. For instance, a downward pressure on the currency (say due to a sudden stop or falling terms of trade) may lead the government to expand money to avoid a banking crisis or an output crunch which case money growth will be responding rather than causing a crisis. In other words, while a contemporary monetary expansion may still reinforce the currency collapse, what causes the latter lies elsewhere and as a result the estimated coefficient on the contemporary money variable will be positively biased. As an alternative, column (3) reports the coefficients of a regression where all variable except for the more truly exogenous capital outflow from the core are lagged one period. We can then see that while the coefficient on the lagged money-output ratio is smaller, it remains nevertheless statistically significant. At the same time, lagging the terms of trade and the UK interest rate indicator yields more sensible estimates for both variables. In particular, the lagged terms of trade deviation from trend is now statistically significant at 5 percent and its estimated coefficient implies that a one percentage point deviation of the terms of trade relative to trend increases the likelihood of a future currency crisis by 2.8 percentage points at mean.

The subsequent columns in Table 5 verify the robustness of the foregoing results to the inclusion of alternative variables in the regressions. Once such a variable is exchange rate overvaluation. As Frankel and Rose (1996), the latter is measured as a percentage deviation from average real exchange rate over the entire 1870-1913, using the Catão and Solomou (2004) multilateral or "effective" real exchange indices. Consistent with the prior embodied in the theoretical model of Section II according to which real exchange rate misalignments are accounted for by the behavior of the various fundamentals already included in the model, the coefficient on the added overvaluation indicator proves to be statistically insignificant at 5% or 10%, albeit taking on the expected positive sign. Next column then considers the inclusion of the ratio of foreign currency denominated debt to exports. The inclusion of this variable is motivated by the recent and growing literature on the so-called “original sin” which emphasises the role of maturity mismatches associated with sizeable stocks of foreign currency denominated sovereign debt (Eichengreen and Hausman, 1999). To the extent that this does not have an immediate counterpart in foreign asset holdings (as is usually the case among capital importing peripheral countries), or in sizeable foreign currency revenues accruing from exports, then such a mismatch is deemed to greatly raise the likelihood of financial crises, including runs on the currency when the economy is hit by adverse shocks (Calvo, Izquierdo and Mejia, 2004). Yet, when added to the regression, that variable too turns out to be unimportant at conventional levels of statistical significance. One possible interpretation is that fiscal imbalances already captured by the expenditure to revenue ratio (g/t) and thus partly correlated with debt to export ratios, are are a sufficient proxy for fiscal risks driving currency movements. One other interpretation is that debt ratios are typically a poor proxy for default risk (as discussed in Catão and Kapur, 2004), and hence also for the currency meltdowns that usually followed sovereign defaults. The latter hypothesis is corroborated by the statistical insignificance (and the wrong sign) of the inclusion of the log of the debt to GDP ratio, as reported in column (6) of Table 5.[31]

Finally, the two last columns of Table 5 show estimates for the inclusion of two other potentially important explanatory variables. One is a trade balance indicator measured as the log of the ratio of exports to imports. This is shown not to add to the explanatory power of the existing variables either, a result consistent with both the contemporary and the historical evidence discussed above that the main source of large exchange rate disturbances typically comes from sudden stops in capital flows

[31] One benefit of considering total debt rather than only external debt as an indicator of the effects of debt burden on currency risk is that the separation of domestic from external debt is especially problematic. This is not only due to the later introduction of gold clauses in domestically issued government debt as for instance in Brazil in the 1890s (see Levy, 1995), but also because countries such as Italy issued public debt (the “rendita”) with the option of being paid in domestic currency at home on in gold in abroad (in the French market more particularly). Such arrangements clearly blurred any meaningful distinction between the two concepts.

and other asset markets developments, rather than from exogenous shifts in trade flows. Last but not least, and also consistent with the results of the ARDL regressions presented earlier, the inclusion of a domestic interest rate variable does not add any significant explanatory power to the regressions. Once again, this suggest that higher domestic interest rates – associated, for instance, with governments or domestic banks' attempts to take off the pressure from the currency and mitigate or prevent reserve losses – typically had no significant impact in minimizing the likelihood of a currency run.

In sum, the econometric results presented in this Section have several noteworthy implications. First and foremost, macroeconomic fundamentals add a significant explanatory power to the various exchange rate determination regressions; as a corollary, the random walk view of exchange rate determination is clearly rejected. Second, a handful of variables can explain a sizeable share of exchange rate movements. This is so both in the short- and long-run, as well as during the less frequent and extreme distress cases typically characterized as "currency crises". Third, in the case of countries that switched between on- and off-gold regimes, the model explains well both spells with no evidence of a structural change across those sub-periods. This is so despite the absence of an explicit indicator of shifts in domestic policy credibility, a factor that some of the existing literature deems crucial in explaining adherence to gold.

VI. Conclusion

The pre-1914 gold standard is widely touted as a singularly successful international system of fixed exchanges. Yet, this view does not aptly describe the experience of much of the periphery of the system. While some peripheral countries did manage to stick to gold throughout, a no less important sub-group experienced dramatic currency crises and painful cyclical contractions associated with such crises.

Given the obvious benefits of sticking to gold in terms of interest rate reduction, government reputation, and avoidance of large output and employment contractions following large depreciations, the puzzle is why several countries did not do so. A first stylized fact documented in this paper is that a clear constraint to maintaining continuous adherence to gold was the behavior of international capital flows: then, as now, capital flows ocasionally came to “sudden stops” which, in the absence of capital account restrictions, called for sharp balance of payments adjustments in recipient countries. However, this cannot be the full explanation since many other countries that also experienced such sudden stops did manage to stick to gold despite the external shock.

The evidence presented above suggests that a key reason is the procyclicality of domestic money and bank credit. In countries for which the historical literature finds the strongest evidence of prevalence of loose regulations on banks of issues, institutional obstacles to loan recovery, and high credit elasticity to cyclically sensitive collateral values, are precisely the ones where large currency crashes immediately followed sudden stops in capital inflows. Not surprisingly, these are also countries where domestic financial systems were shallower at the onset of the boom. Such financial shallowness implies that large swings in capital inflows tend to have stronger direct effects on money supply, while the low market liquidity typically associated with financial shallowness implies that banks tend to either have to resort to fire sales of their assets during “bad times” or else rely on a government bail-out. The latter, in turn, not only further exacerbates the accumulation of bad credit during a downturn, but also further deteriorates the public finances and put extra downward pressure on the currency following the sudden stop. Overall, thus, this brings back to fore Whale’s (1937) and Ford’s (1962) earlier views on the importance of the domestic financial imperfections in propagating exogenous shocks and thus accounting for differences in how the gold standard adjustment worked in different countries.

Where does this leave us in terms of the debate on the relative importance of domestic v. international factors in generating exchange rate instability in several peripheral countries? Clearly, the empirical evidence presented in this paper supports at least two aspects of the international asymmetry view posited by Triffin (1964) and others: one is that the supply of foreign capital faced by individual peripheral countries was volatile; in addition, shallower financial markets in the periphery and an apparently lower elasticity of such flows to interest differentials with the core, implied that monetary tightening in the periphery was of little use in luring capital back in the short run to help stabilize the currency. Accordingly, the burden of adjusting to sudden stops would squarely fall on the exchange rate. This was the more so whenever a negative capital account shock was reinforced by a terms of trade deterioration, as for instance, during the famed Baring crisis of 1890-91.

This is not to say, however, that domestic fiscal policy did not matter. Loose fiscal policies was clearly a relevant factor especially in some countries, notably Argentina and Greece. But in contrast with some of literature which depicts inconvertible paper money as a direct result of fiscal progligacy, reality was more complex. Countries where the fiscal stance was not out of control also experienced sudden stops and currency crisis; and, conversely, parts of the periphery that run fiscal deficits and had higher debt to GDP ratios did not experience currency runs. Quite aside from the fact that a country's fiscal stance typically has a sizeable endogenous component, the econometric evidence presented in this paper indicates that the effects of fiscal behavior on the exchange rate were not one to one.[32] What instead appears to have been crucial was the *combination* of procyclical fiscal policies with an even higher procyclicality of monetary expansion to capital inflows and

[32] Fiscal behavior also responded to events more truly endogenous to economic policy decision making such as wars. War-related pressures on the budget and currency issuance were important, for instance, in Greece during the 1877 Russian-Turkish war (in which Greece participated) as well as in the war with Turkey in 1897 (Lazaretou, 1993); similar fiscal-related pressures on the currency were also documented for Spain during the late 1898 Cuban war (Prados, 2003), as well as Chile during the Chaco war with Peru between 1879 and 1883 (Llona Rodriguez, 2000). It is striking that despite the idiosyncratic and country specific nature of these events that had a bearing on the respective currencies, one can still clearly detect the systemic co-movements and causal regularities highlighted in the regressions. If anything, this reinforces the robustness of the above results.

aggregate demand booms, as butressed by the fact that the largest exchange rate crashes occurred precisely in those places where the monetary stance was more procyclical.

Finally, this paper's econometric evidence calls into question the hypothesis that improved policy credibility and commitment to convertibility directly translate into currency stability. This is clear from the remarkable coefficient stability in the regressions: pegging to gold did not significantly alter the elasticity of one's currency to the different factors highlighted above. It thus seems that a more sensible interpretation of the data is that what gold membership did allow was to reconcile monetary deepening with a fixed exchange rate, but only during periods of higher capital inflows and improving terms of trade. It is therefore not surprising that much of the periphery only managed to join and stick to gold during the world boom of the 1900s, when commodity terms of trade were on the rise and the world capital market embarked upon an unprecedentally long boom only brought to a halt by World War I. In other words, such gold pegs had a marked cyclical character which was endogenous to the state of the international business cycle. All in all, while the weight of the evidence does not deny that fiscal discipline and the credibility of policy makers are important, it does suggests that they cannot by themselves buy currency stability wherever domestic financial imperfections loom large and sudden stops strike.

References

Aceña, Pedro M. and Reis Jaime, 2000, *Monetary Standards in the Periphery. Paper, Silver and Gold, 1854–1933*, London.

Bernanke, B.S. and A. Blinder, 1988, "Credit, Money, and Aggregate Demand," *American Economic Review*, 73, pp. 257–76.

__________, M. Gertler and Simon Gilchrist, 1996, "The Financial Accelerator in a Quantitative Business Cycle Framework," *Review of Economics and Statistics*, Vol. 78, no. 1 pp. 1-15.

Blattman, Chris, Jason Hwang, and Jeffrey Williamson, 2003, "Terms of Trade and Economic Performance in the Periphery 1870-1940," NBER Working Paper 9940 (Cambridge, Massachusetts: National Bureau of Economic Research).

Bloomfield, Arthur, 1968, *Patterns of Fluctuation in International Investment before 1914*, Princeton.

Bordo, Michael and H. Rockoff, 1996, "The Gold Standard as a 'Good Housekeeping Seal of Approval,'" *Journal of Economic History,* 56(2), pp. 384–428.

___________, Christopher Meissner, and Angela Redish, 2003, "How Original Sin was Overcome: The Evolution of External Debt Denominated in Domestic Currency in the United States and British Dominions," NBER working paper 9841, Cambridge, Mass.

Cairncross, A.K. 1953, *Home and Foreign Investment, 1870-1913: Studies in Capital Accumulation* , Cambridge.

Calvo, Guillermo, Alejandro Izquierdo, and Luis Fernando Mejia, 2004, "On the Empirics of Sudden Stops: The Relevance of Balance Sheet Effects", NBER WP 10520.

Cardoso, Eliana, 1983, "Exchange Rates in Nineteenth Century Brazil: An Econometric Model," *Journal of Development Studies*, 19, pp.170-78.

Catão, Luis and Marco Terrones, 2004, "Fiscal Deficits and Inflation", *Journal of Monetary Economics*, forthcoming.

_________ and Sandeep Kapur, 2004, "Missing Link: Volatility and the Debt Intolerance Paradox," IMF Working paper WP/04/51.

_________ and Solomos Solomou, 2004, "Effective Exchange Rates and the Classical Gold Standard Adjustment," *American Economic Review*, Forthcoming.

Cooper, Richard, 1971, "Currency Devaluation in Developing Countries", Essays in International Finance No. 86, Princeton University.

Crisp, Olga, 1976, *Studies in the Russian Economy before 1914*, London.

De Cecco, Marcello, 1974, *Money and Empire: The International Gold Standard, 1890-1914*, Oxford.

deLong, B. 1999, "Financial Crises in the 1890s and the 1990s: Must History Repeat?", *Brookings Papers on Economic Activity*, Vol.2, pp. 253-294.

della Paolera, G.and A. M. Taylor, 2001, *Straining at the Anchor: The Argentine Currency Board and the Search for Macroeconomic Stability, 1880–1935.* NBER, Chicago.

Diaz-Alejandro, Carlos F., 1963, A Note on the Impact of Devaluation and the Redistribution Effect", *Journal of Political Economy*, 84(6).

________ 1984, *"*Latin American Debt: I Don't Think We Are in Kansas Anymore*," Brookings Papers on Economic Activity, Vol. 2, pp. 335–403*

Drummond, Ian, 1976, "The Russian Gold Standard, 1879-1914," *Journal of Economic History*, 36(3), pp. 663-88.

Edelstein, M., 1982, *Overseas Investment in the Age of High Imperialism,* New York: Columbia University Press.

Edwards, S., 1986, "Are Devaluations Contractionary?", Review of Economics and Statistics, 68, pp.501-508.

_________ 1989, *Real Exchange Rates, Devaluation, and Adjustment: Exchange Rate Policy in Developing Countries*, Cambridge, Mass.

Eichengreen, B., 1992, "The Gold Standard Since Alec Ford", in S.N. Broadberry and N.F.R. Crafts (eds.), *Britain in the International Economy 1870–1939*, Cambridge, pp 47-79.

______, 2003, *Capital Flows and Crises*, Cambridge, Mass.

______, and M. Flandreau, 1998, *The Gold Standard in Theory and History*, New York.

———, and Ricardo Haussman, 1999, "Exchange Rate Regimes and Financial Fragility," NBER Working Paper 7418 (Cambridge, Massachusetts: National Bureau of Economic Research).

Fishlow, Albert, 1989, "Conditionality and Willingness to Pay: Some Parallels from the 1890s," in Barry Eichengreen and Peter H. Lindert, *The International Debt Crisis in Historical Perspective*, Cambridge, Mass.

Flandreau, 1996, "The French Crime of 1873: An Essay on the Emergence of the International Gold Standard," *Journal of Economic History*, 56, pp. 862-97.

_______ and Clemens Jobst, 2004, "The ties that divide: A network analysis of the international monetary system, 1890-1910", mimeo.

_______ and F. Zulmer, 2004, *The Making of Global Finance* 1880-1913, OECD, Paris.

Fratianni, Michele and Franco Spinelli, 1984, "Italy in the Gold Standard Period, 1861-1914," in Michael D. Bordo and Anna J. Schwartz (eds.) *A Retrospective of the Classical Gold Standard 1821-1931*, Chicago.

Ford, A.G., 1962, *The Gold Standard 1880–1914: Britain and Argentina,* Oxford.

Frankel, Jeffrey, 1979, "On the Mark: A Theory of Floating Exchange Rates Based on Real Interest Differentials," *The American Economic Review*, 69(4), pp.610-22.

____________, and Andrew Rose, 1985, "A Survey of Empirical Research on Nominal Exchange Rates," in Kenneth Rogoff and Gene Grossman (eds.), *Handbook of International Economics*, Amsterdan.

___________, 2004, "Contractionary Currency Crashes in Developing Countries," Paper presented at The 5th Mundell-Flemming Lecture, International Monetary Fund, mimeo.

Furtado, Celso, 1963, *Economic Growth of Brazil: A Survey from Colonial Times to Present*, Berkeley.

Gordon, R. J., 1983, "A Century of Evidence on Wage and Price Stickiness in the US, the UK and Japan", in J. Tobin (ed.), *Macro-Economics, Prices and Quantities*.

Gregory, Paul R., 1979, "The Russian Balance of Payments, the Gold Standard, and Monetary Policy: A Historical Example of Foreign Capital Movements," *Journal of Economic History*, 39(2), pp. 379-99.

______________, 1982, *Russian National Income, 1885–1913,* Cambridge.

Haber, Stephen, 1997, Financial Markets and Industrial Development: A Comparative Study of Governmental Regulation, Financial Innovation, and Industrial Structure in Brazil and Mexico, 1840-1930", in S. Haber (ed.), *How Latin American Fell Behind. Essays in the Economic Histories of Brazil and Mexico,* Stanford.

Jeanne, Olivier and Andrew Rose, 2002, "Noise Trading and Exchange Rate Regimes," *Quarterly Journal of Economics*,

Kaminsky, Graciela and Carmen Reinhart, 1999, "The Twin Crises: The Cause of Banking and Balance of Payments Problems," *American Economic Review*, Vol.89, 3, pp. 473-500.

Kelly, Trish, 1998, "Ability and Willingness to Pay in the Age of the Pax Britannica," *Explorations in Economic History*, 35, pp. 31-58.

Kiyotaki, N. and J. Moore, 1997, "Credit Cycles," *Journal of Political Economy,* 105(2).

Lains, Pedro, 1995, *A Economia Portuguesa no Século XIX*, Lisbon.

Lazaretou, Sophia, 1993, "Monetary and Fiscal Policies in Greece: 1833-1914," *Journal of European Economic History*, 22(2), pp. 285-311.

______________, 1995, “Government Spending, Monetary Policies, and Exchange Rate Regime Switches: The Dracma in the Gold Standard Period,” *Explorations in Economic History*, 32, pp.28-50.

Levy, Maria Barbara, 1995, “The Brazilian Public Debt – Domestic and Foreign, 1824-1913,” in Reinhard Liehr (ed.), *The Public Debt in Latin America in Historical Perspective*, Fankfurt.

Lewis, W.A., 1978, *Growth and Fluctuations, 1870–1913***,** London.

Llona Rodriguez, Agustín, 2000, “Chile During the Gold Standard: A Successful Paper Money Expeience,” in Acena, P.M. and Reis, J. (eds.), *Monetary Standards in the Periphery: Paper, Silver and Gold, 1854–1933*, London.

Mata, Maria Eugénia, 1987,”Câmbios e Política Cambial na Economia Portuguesa, 1891-1931”, *Cadernos da Revista de História Económica e Social*, Vol. 8, Lisbon.

Meese, Richard and Kenneth Rogoff, 1983, “Empirical Exchange Rate Models of the Seventies: Do they Fit Out of Sample?,” *Journal of International Economics* 14, pp.3-23.

Millesi-Ferreti, Gian Maria and Phillip Lane, 2004, “The Transfer Problem Revisited: Net Foreign Assets and Real Exchange Rates,” *Review of Economics and Statistics*, forthcoming.

Morley, Samuel, 1982, “On the Effects of Devaluations During Stabilization Programs in LDCs,” Review of Economics and Statistics, 74(1), pp. 21-7.

Obstfeld, Maurice and Kenneth Rogoff, 1996, *Foundations of International Macroeconomics*, Cambridge, Mass.

________ and Taylor, A. M., 2003, “Sovereign risk, credibility and the gold standard: 1870-1913 versus 1925-31”, *Economic Journal*, 113, 487, pp 241-75.

Pesaran, M.H. and Smith, R.P., 1995, “Estimating Long-Run Relationships From Dynamic Heterogenous Panels,” *Journal of Econometrics*, 68, pp. 79–113.

__________, Y. Shin, and R. P. Smith, 1999, “Pooled Mean Group Estimation of Dynamic Heterogeneous Panels,” *Journal of the American Statistical Association*, 94, pp. 621–33.

Prados, Leandro, 2003, *El progreso economico de Espana, 1850-2000*, Madrid

Reis, Jaime, 2000, “The Gold Standard in Portugal 1854-91,” in Aceña, P.M. and Jaime Reis (eds.), 2000, *Monetary Standards in the Periphery. Paper, Silver and Gold, 1854–1933*, London.

Tattara, Giuseppe, 2001, “Was Italy Ever on Gold?,” in Aceña, P.M. and Jaime Reis (eds.), 2000, *Monetary Standards in the Periphery. Paper, Silver and Gold, 1854–1933*, London.

_______ and M. Volpe, 1997, "Italy, The Fiscal-Dominance Model, and The Gold-Standard Age," in Marcuzzo, M.C., Officer, L.H. and Rosselli, A. (eds.), *Monetary Standards and Exchange Rates*, New York and London.

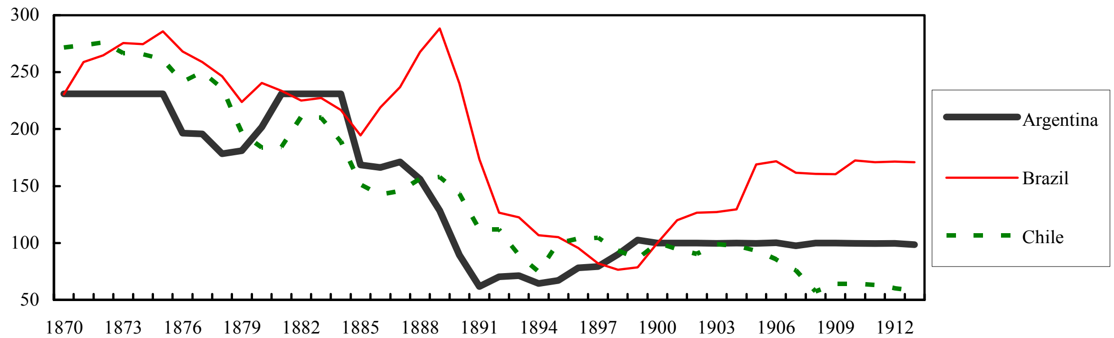
Triffin, R., 1964, "The Evolution of the International Monetary System: Historical Reappraisal and Future Perspectives," Princeton Studies in International Finance, No.12.

Triner, Gail, 2000, *Banking and Economic Development: Brazil, 1889-1930*. New York.

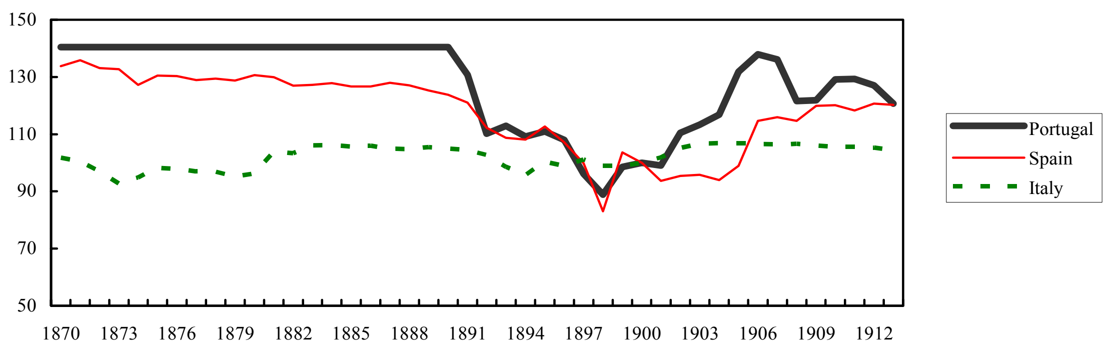
Whale, P.B., 1937, "The Working of the prewar Gold Standard," *Economica*, pp. 18-32.

Figure 1. Nominal Exchange Rates in the Periphery
(1900=100)

Paper-Currency Regimes: Latin America



Paper-Currency Regimes: Europe



Paper-Currency Regimes: Europe (cont.)

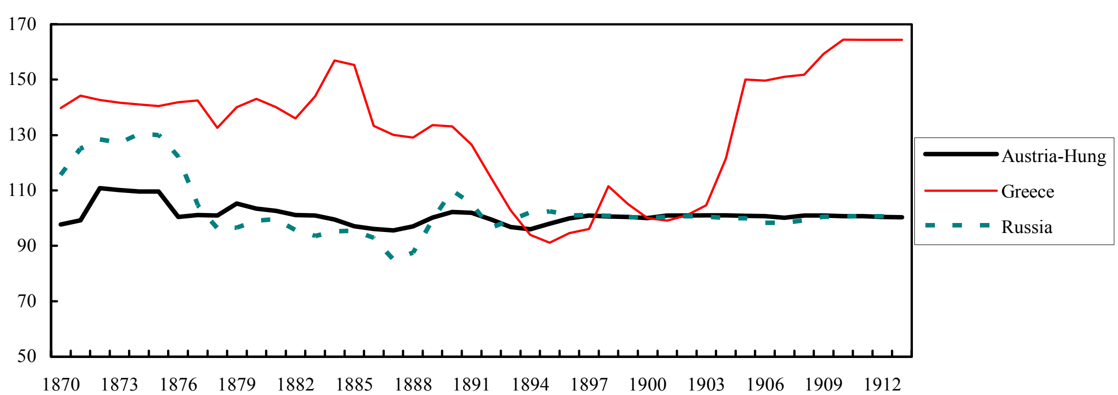


Figure 2. Net Capital Exports from Core Countries (million pounds)

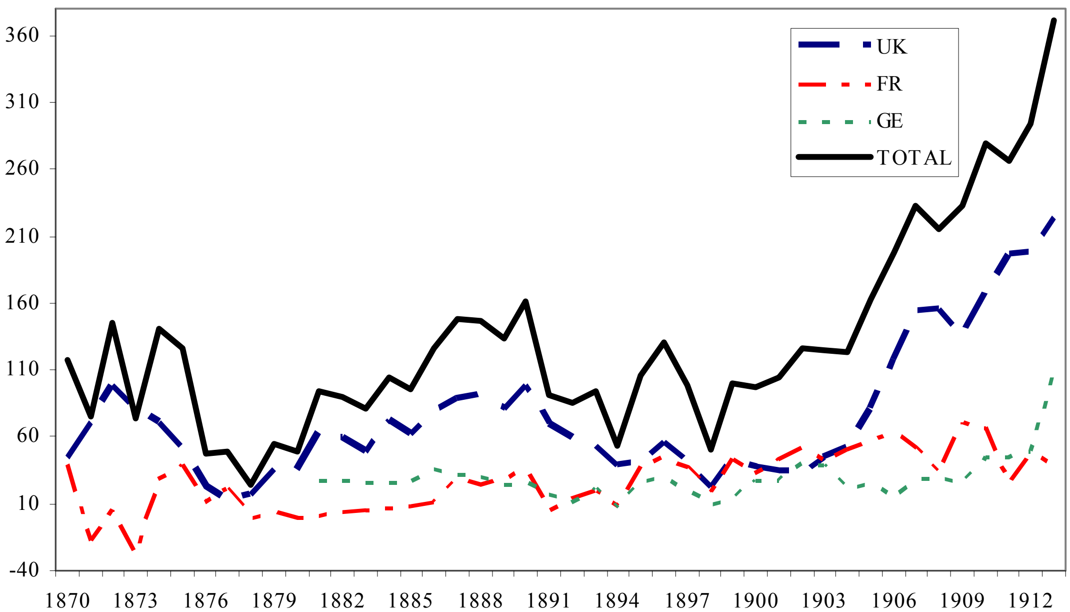
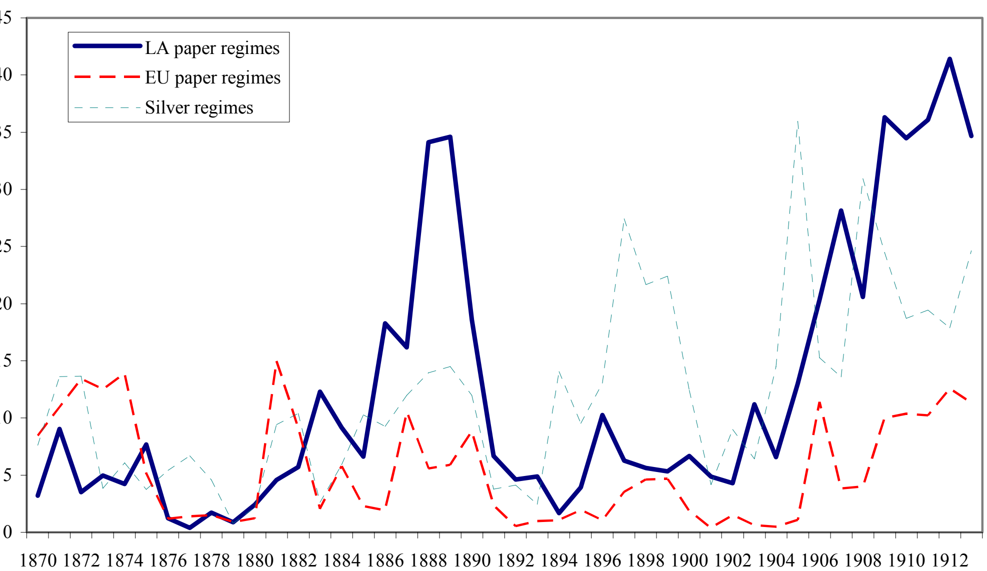
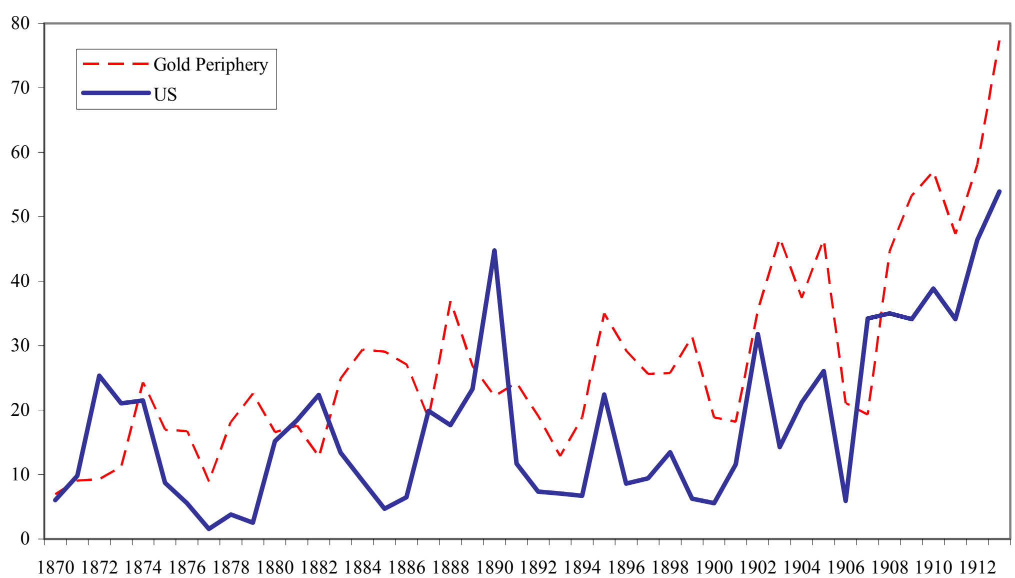


Figure 3. UK Capital Exports by Monetary Regimes and Regions (millions of sterling)

Non-Gold Periphery

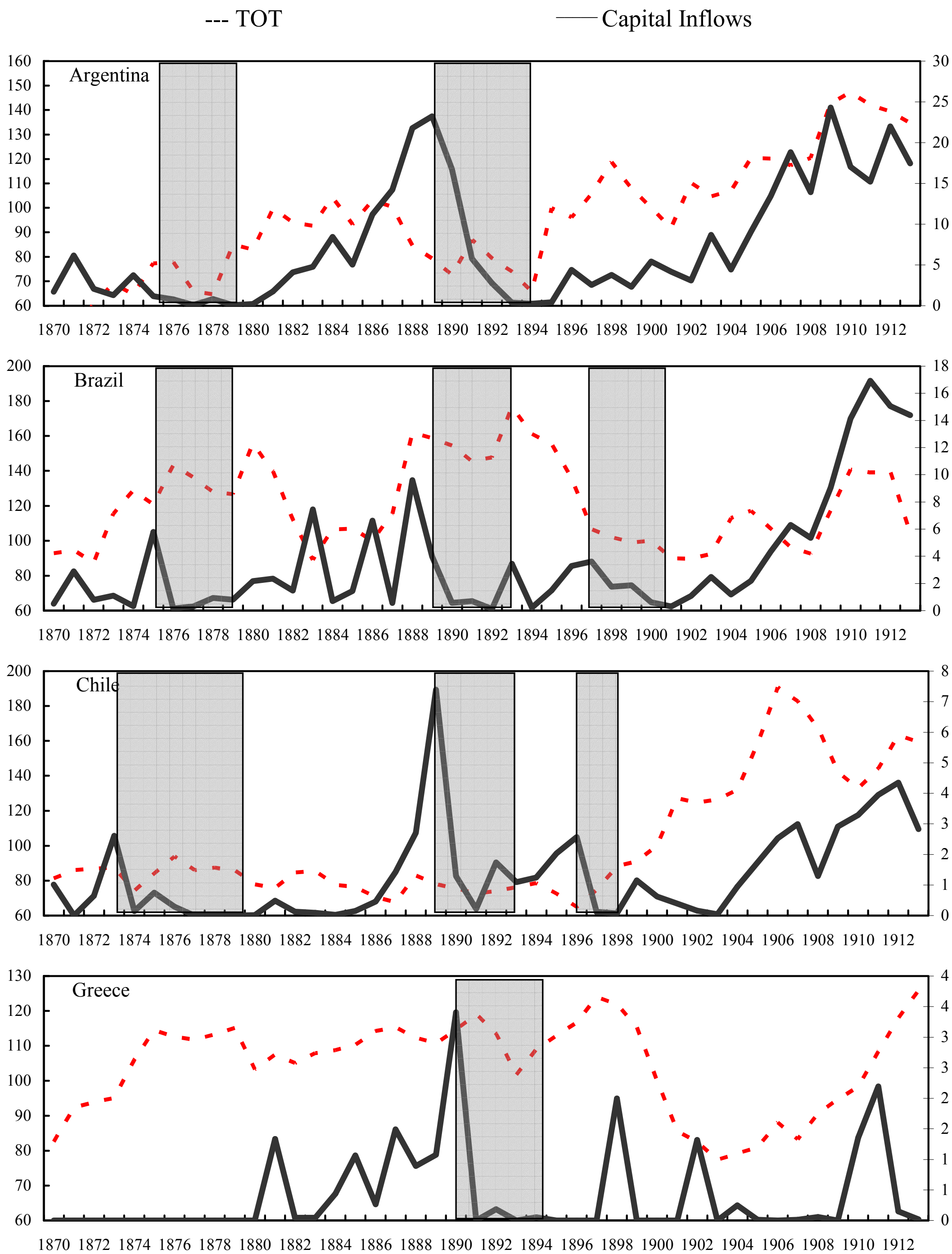


Gold Block



Source: Stone (1999).

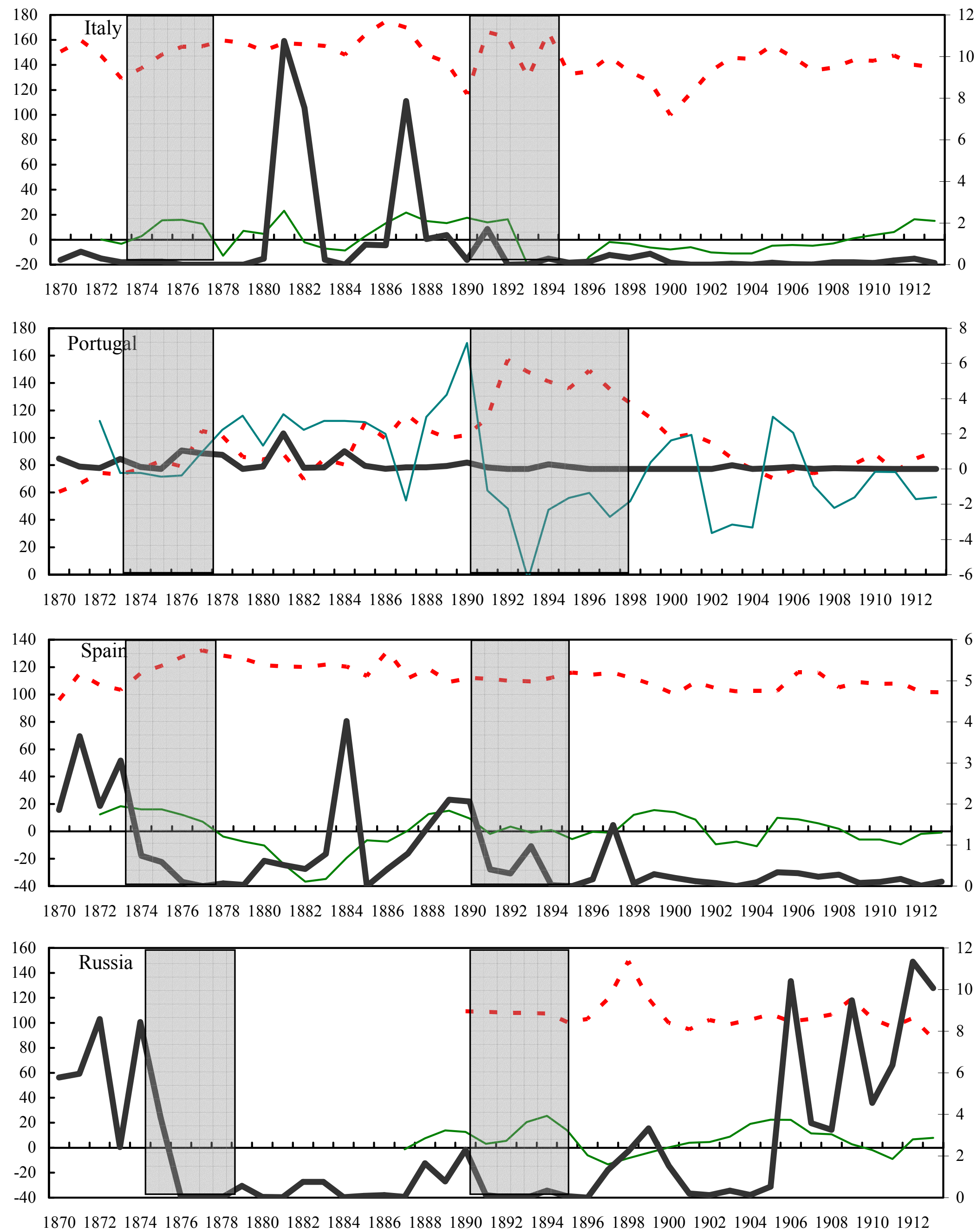
Figure 4. Net Barter Terms of Trade and Gross Foreign Capital Inflows



Sources: Stone (1999), Catão and Solomou (2004), and Clemens, Hwang and Williamson (2004).

Figure 4 (cont.). Net Barter Terms of Trade and Gross Foreign Capital Inflows

--- TOT — UK Capital Inflows — Net External Public Borrowing (3-year MA)



Sources: UK capital inflows from Stone (1999). TOT from Lains (1995), and Clemens, Hwang and Williamson (2004), and Prados (2003). Public debt data, see appendix.

Figure 5. Nominal Exchange Rate, Fiscal Indicators, and Monetization Ratio

- ι - NER — M2/GDP (right axis) - - G/T - D/Y(%)

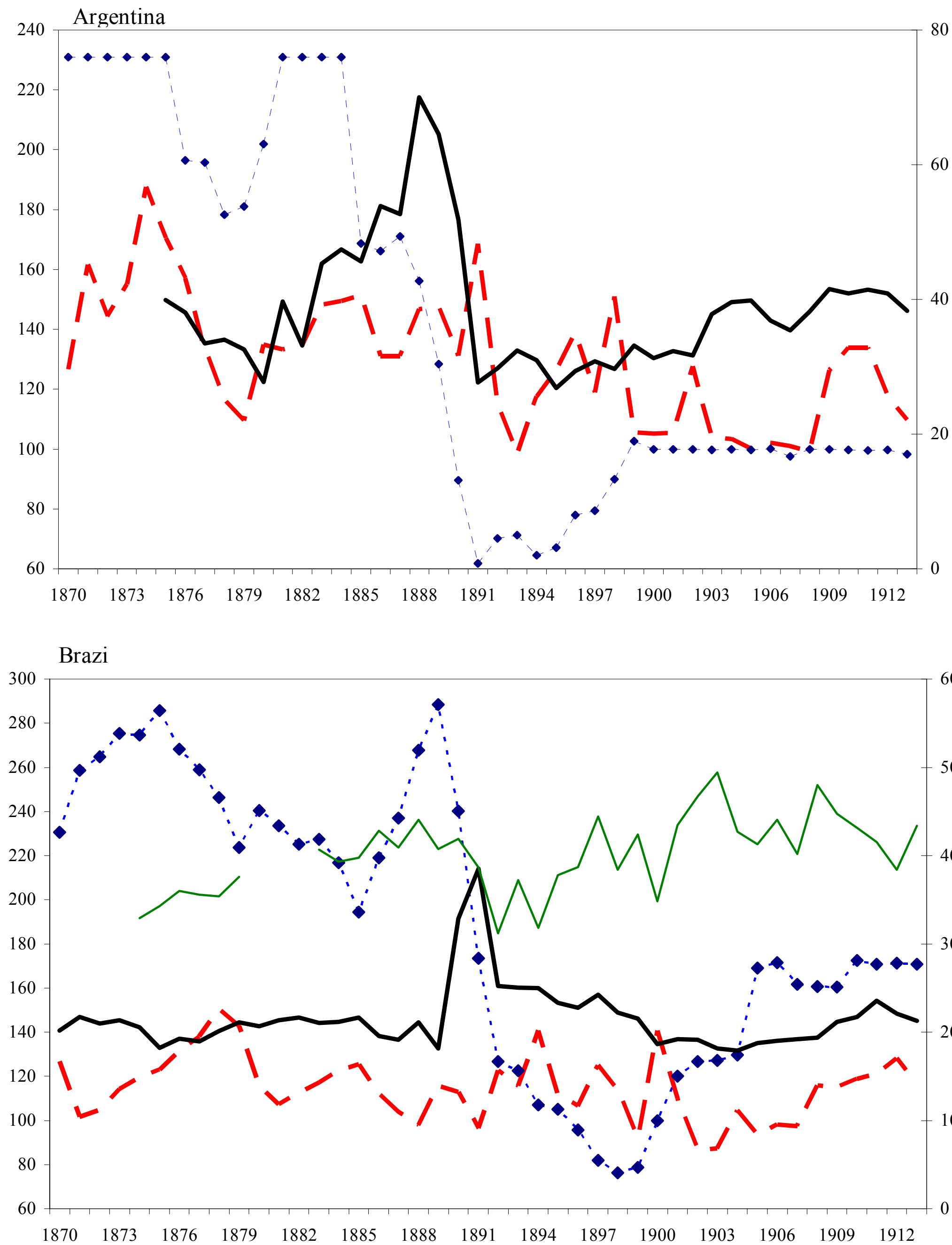
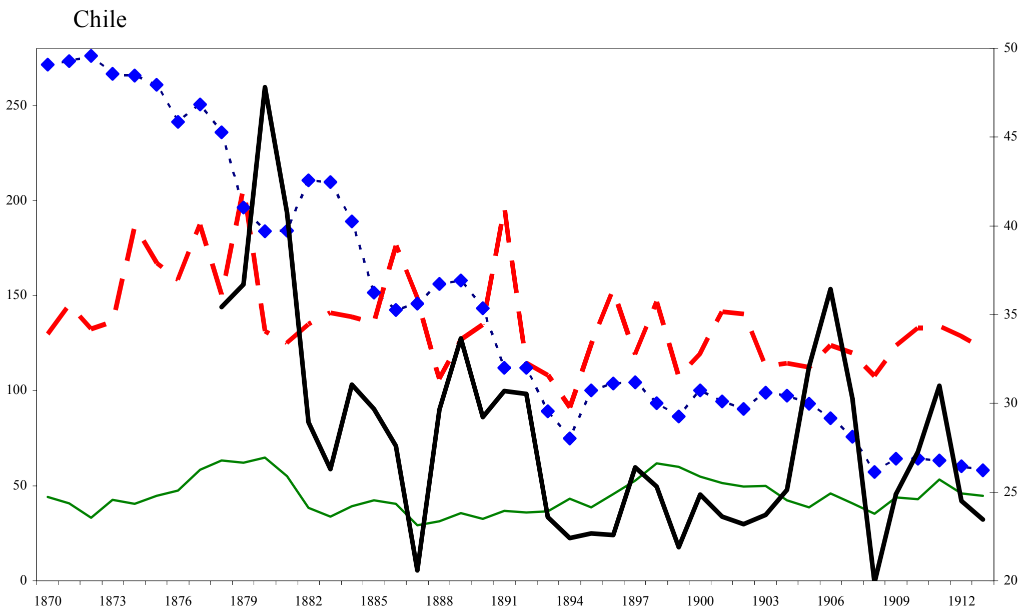
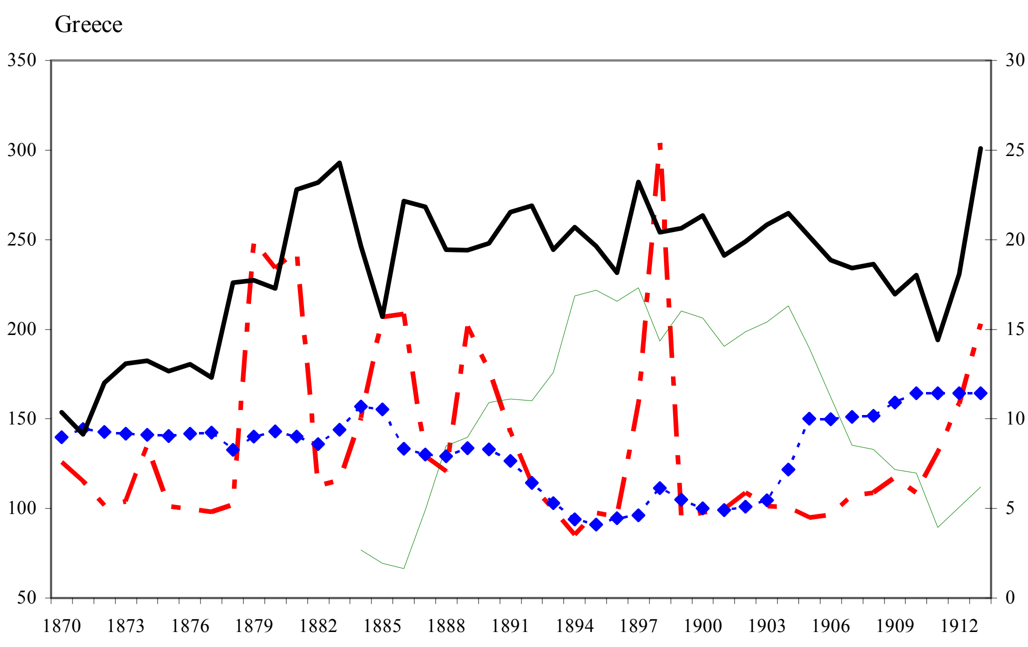


Figure 5 (cont.) Nominal Exchange Rate, Fiscal Indicators, and Monetization Ratios

->- NER(1900=100) — M2/GDP(right axis) - - G/T(%) - D/Y(%)



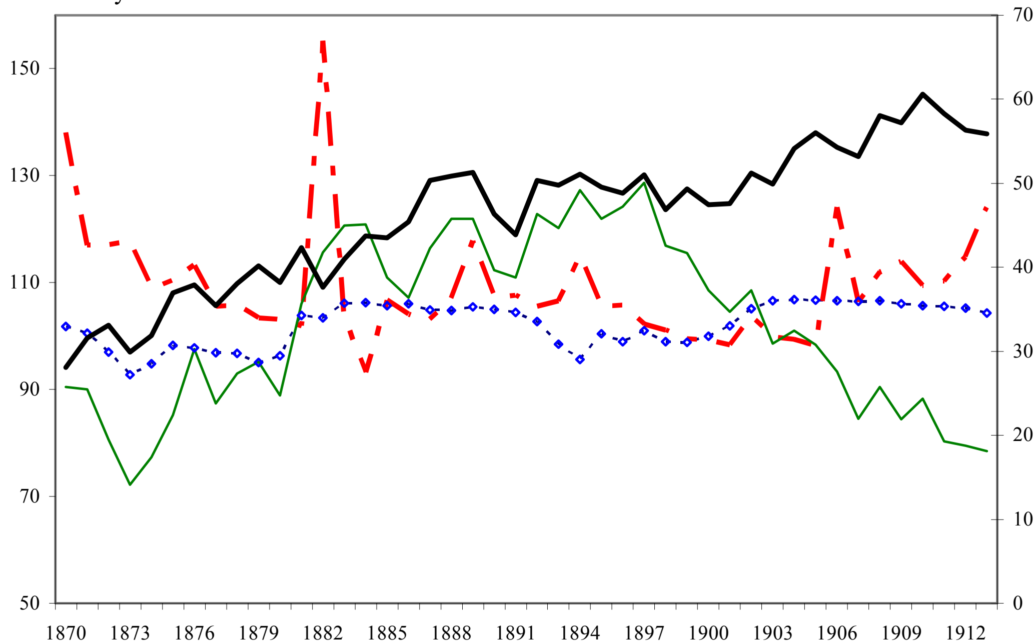


Note: Money figures for Greece correspond to currency in circulation only.

Figure 5 (cont.) Nominal Exchange Rate, Fiscal Indicators, and Monetization Rati

->- NER(1900=100) — M2/GDP(right axis) - - G/T(%) - D/Y(%)

Italy



Portugal

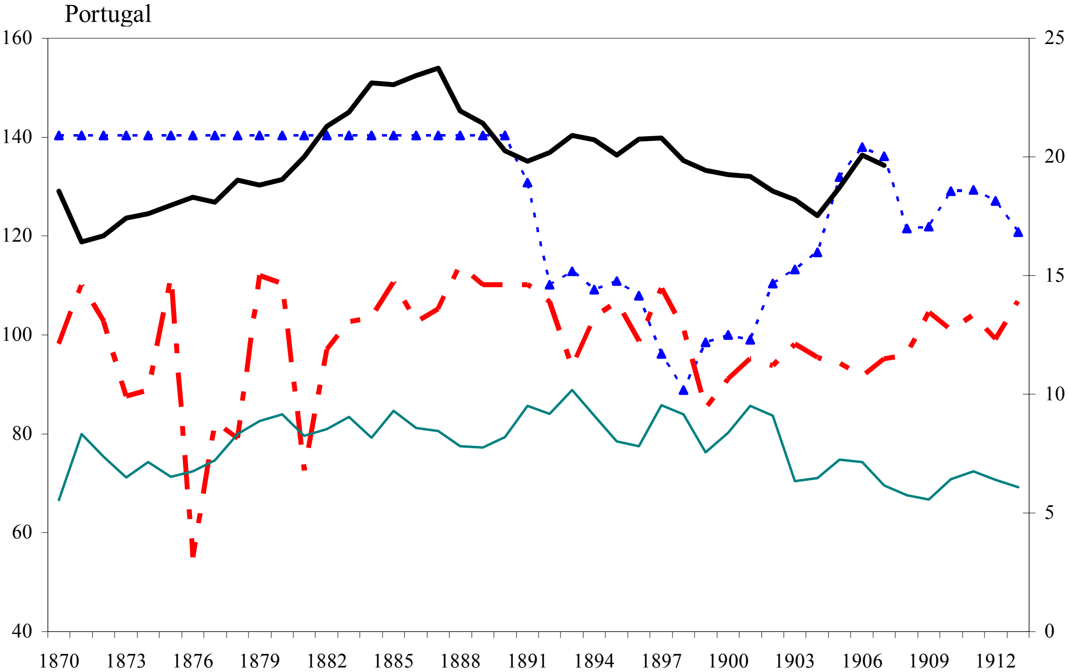
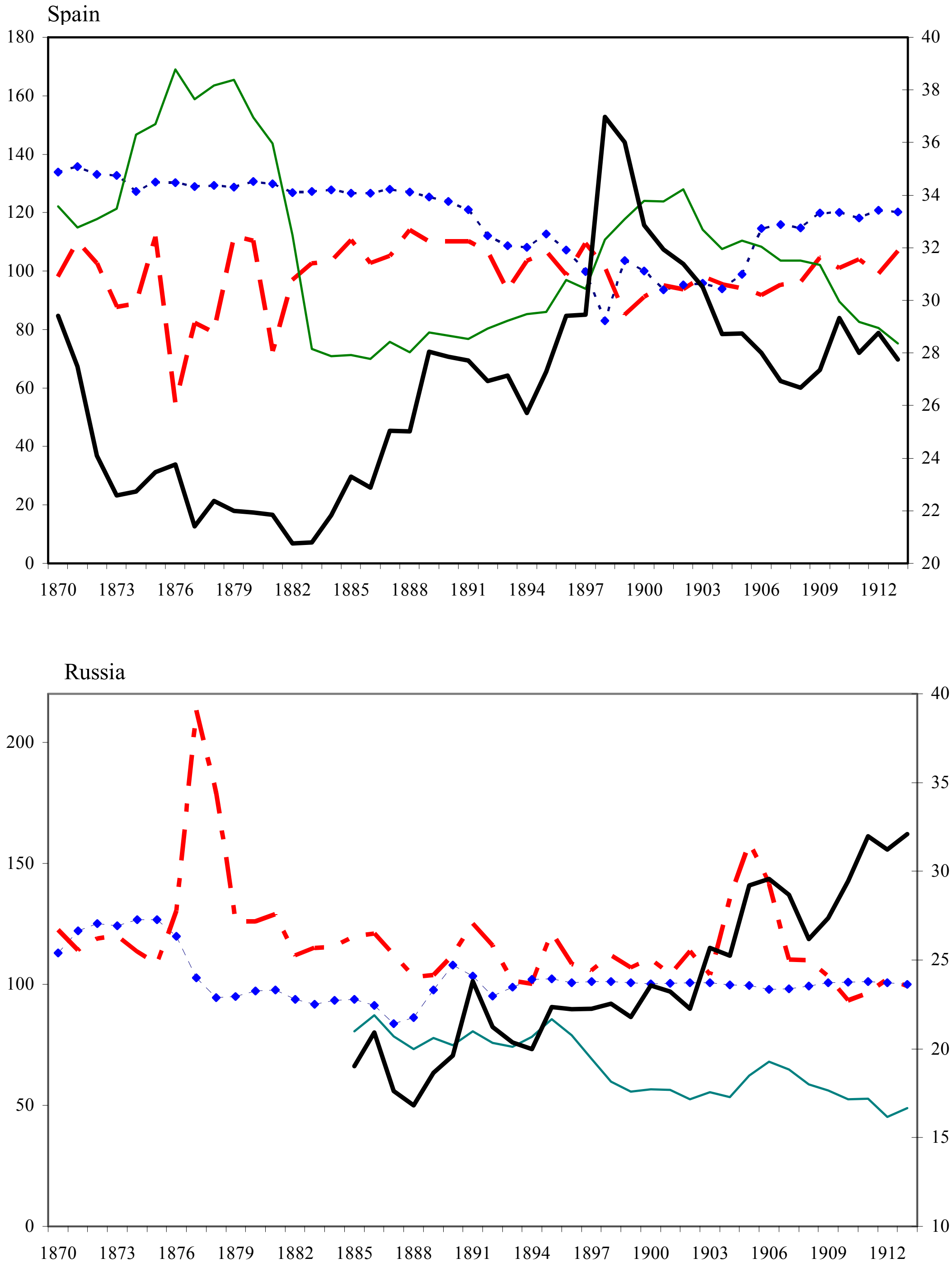


Figure 5 (cont.) Nominal Exchange Rate, Fiscal Indicators, and Monetization Rati

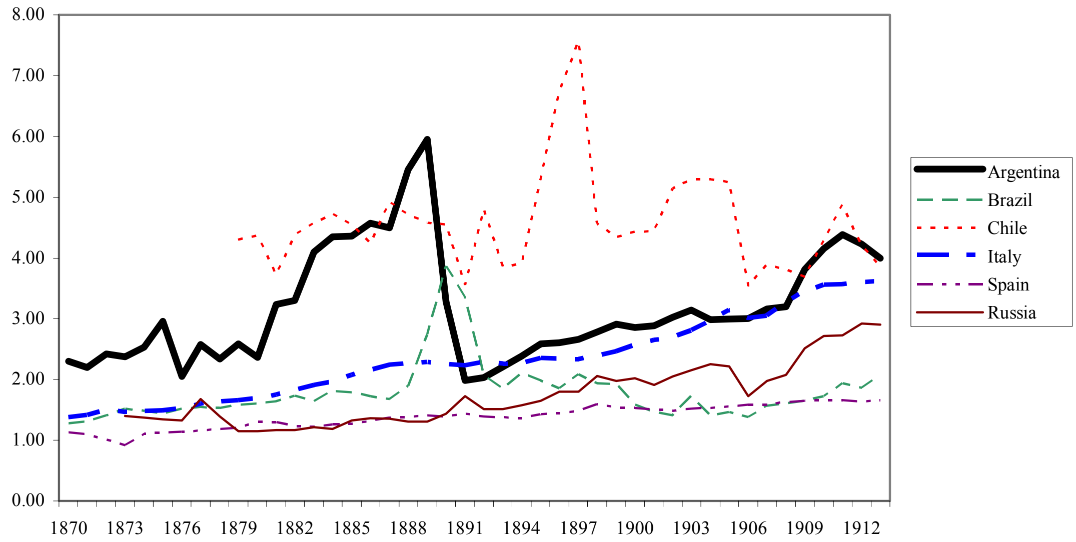
->- NER(1900=100) —M2/GDP(right axis) - - G/T(%) - D/Y(%)



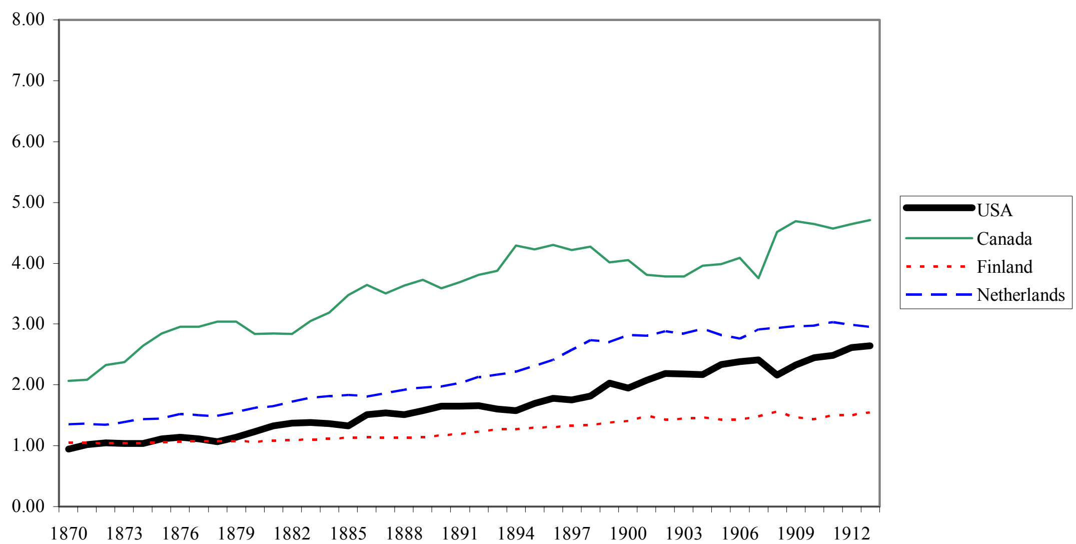
Sources: See Appendix.

Figure 6. The M2 Multiplier

The Non-Gold Periphery



Gold Countries



Source: See appendix.

Figure 7. Argentina and the United States: Ratio of Gross Capital Inflows to Broad Money (percent)

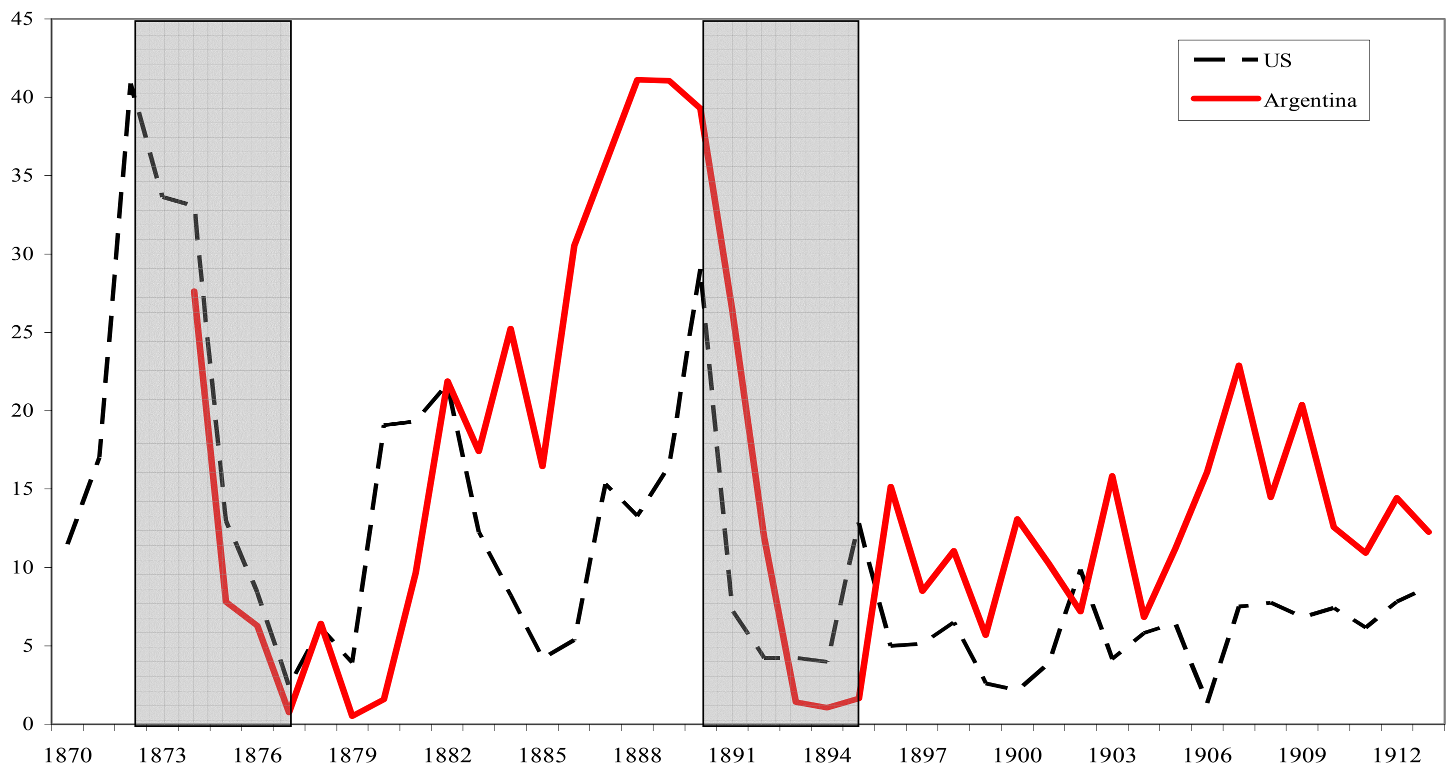


Table 1. Correlation between Sovereign Spreads and International Capital Flows, 1879-1913
(coefficients computed over annual data)

	LOFICORE	LOFIUK	LOFIFR	LOFIGE	FICORECY
Argentina	-0.84	-0.59	-0.48	-0.6	-0.68
Brazil	-0.39	-0.58	0.04	-0.53	-0.44
Chile	-0.74	-0.75	-0.56	-0.50	-0.65
Greece	-0.48	-0.46	-0.38	-0.28	-0.55
Italy	-0.71	-0.68	-0.31	-0.65	-0.56
Portugal	-0.43	-0.68	0.07	-0.57	-0.59
Spain	-0.68	-0.51	-0.56	-0.50	-0.50
Russia	-0.44	-0.02	-0.58	-0.09	-0.01
Mean	-0.59	-0.53	-0.35	-0.47	-0.50

Notation:

LOFICORE = log of the sum of capital outflows from Britain, France and Germany in million pounds.
LOFIUK = log of gross capital outflow from Britain in million pounds.
LOFIFR = log of net capital outflow from France in million francs.
LOFIGE = log of net capital outflow from Germany in million marks.
FICORECY = deviations of LOFICORE from a linear regression trend.

Sources: Spreads from Mauro et al. (1999); Capital Flow data from Stone (1999).

Table 2. Correlation between Terms of Trade and Foreign Capital Inflows, 1872-1913 1/
(coefficients computed over annual data)

	Contemporaneous Levels	Contemporaneous Growth Rates	Growth Rates with Lagged TOT
Argentina	0.69	0.19	0.37
Brazil	0.06	0.09	-0.04
Chile	0.48	-0.13	0.11
Greece	0.17	0.17	-0.26
Italy	0.34	0.51	-0.40
Portugal	0.04	-0.22	-0.10
Spain	0.32	0.03	-0.10
Russia	-0.11	-0.03	0.18
Mean	0.25	0.08	-0.03

1/ Capital Flows for each country correspond to the Stone (1999) series on UK gross portfolio flows to the respective country in million pounds deflated by the UK wholesale price index reported in Mitchell (1988). All data starts in 1872 except for Russia, for which available TOT data restricts the computations to the 1890-1913 period.

Sources: Spreads from Mauro et al. (1999); Capital Flow data from Stone (1999).

Table 3. Nominal Exchange Rate Determinants: Pooled Mean Group Estimates
(all variables in logs, t-ratios in parenthesis)

	tot	g/t	m3-m3*	y-y*	TREND	fic	EC	ML	h_{tot}	$h_{g/t}$	h_{m3-m3^*}	h_{y-y^*}
I.	0.38	-0.02	-1.08	0.36	-0.01	0.07	-0.29	470.35	0.65	0.55	16.41	2.16
	(4.93)	(-0.27)	(-13.20)	(2.61)	(-2.48)	(3.03)	(-7.67)		[p=0.42]	[p=0.46]	[p=0.00]	[p=0.14]
II.	0.35	…	-1.07	0.39	-0.01	0.07	-0.30	476.41	0.09	…	13.95	0.28
	(4.73)	…	(-13.68)	(2.95)	(-2.51)	(2.96)	(-8.86)		[p=0.76]	…	[p=0.00]	[p=0.60]
III.	0.15	-0.04	-0.79	0.03	-0.01	0.07	-0.43	483.49	0.01	0.45	…	0.87
	(3.35)	(-0.97)	(-4.49)	(0.46)	(-2.06)	(3.12)	(-5.74)		[p=0.91]	[p=0.45]	…	[p=0.87]
IV.	0.16	…	-0.78	0.068	-0.008	0.07	-0.43	487.58	0.67	…	…	0.55
	(3.52)	…	(-4.91)	(0.90)	(-2.02)	(2.98)	(-6.32)		[p=0.91]	…	…	[p=0.46]

Table 4. ARDL Estimates of Nominal Exchange Rate Determinants
(all variables in logs except for interest rates which are in percent a year)
(t-ratios in parenthesis)

Dependent Variable: Nominal Effective Exchange Rate (log of Index, 1913=100)

		Long-Run Elasticities								*Short-Run Elasticities*								*Other Statistics*			
		tot	**g/t**	**m3-m3***	**y-y***	**n-n***	**efic**	**i-i***	**edebt/x**	**tot**	**g/t**	**m3-m3***	**y-y***	**n-n***	**efic**	**i-i***	**edebt/x**	**ECM**	**Adj-R^2**	**GS F-test**	**Period**
Argentina	I	**1.32**	-0.25	**-0.75**	-0.23	-0.53	**0.26**	0.06	--	**0.2**	-0.14	0.02	-0.14	**5.86**	**0.14**	0.03	--	**-0.58**	0.97	0.45	1879-1913
		(4.45)	(-1.38)	**(-3.15)**	(-0.55)	(-1.25)	**(2.65)**	(1.25)	--	**(1.56)**	(-1.46)	(0.11)	(-0.53)	**(2.36)**	**(2.55)**	(1.35)	--	**(-5.81)**		[p=0.77]	
	II	**1.31**	-0.25	**-0.57**	-0.35	**-0.82**	**0.21**	--	--	**0.2**	-0.15	0.18	-0.14	**5.86**	**0.14**	--	--	**-0.57**	0.97	0.52	1879-1913
		(4.67)	(-1.19)	**(-3.06)**	(-0.85)	**(-1.78)**	**(2.07)**	--	--	**(1.58)**	(-1.46)	(0.12)	(-0.54)	**(2.36)**	**(2.55)**	--	--	**(-5.82)**		[p=0.72]	
	III	**1.00**	--	**-0.53**	--	**-0.79**	**0.21**	--	**-0.11**	-0.01	--	**-0.2**	--	**-0.79**	**0.21**	--	**-0.11**	**-1.00**	0.97	0.67	1885-1913
		(8.58)	--	**(-9.39)**	--	**(-6.35)**	**(5.13)**	--	**(-4.12)**	(-0.06)	--	**(2.06)**	--	**(-6.35)**	**(6.61)**	--	**(-4.12)**	**--**		[p=0.62]	
Brazil	I	**0.71**	0.91	**-1.36**	-0.19	**-1.24**	**0.26**	--	--	**0.24**	0.00	**-0.46**	-0.07	**-0.42**	**0.09**	--	--	**-0.34**	0.97	0.06	1879-1913
		(4.28)	(1.64)	**(-4.69)**	(-0.21)	**(2.35)**	**(2.14)**	--	--	**(3.25)**	(0.01)	**(-9.91)**	(-0.22)	**(-2.72)**	**(2.14)**	--	--	**(-3.87)**		[p=0.99]	
	II	**0.68**	--	**-1.06**	**1.05**	**-0.73**	0.12	--	--	**0.30**	--	**-0.47**	**0.46**	**-1.32**	0.06	--	--	**-0.44**	0.97	0.16	1879-1913
		(4.99)	--	**(-6.83)**	**(2.67)**	**(-2.02)**	(1.34)	--	--	**(3.92)**	--	**(-9.31)**	**(1.97)**	**(-1.98)**	(1.22)	--	--	**(-5.18)**		[p=0.92]	
	III	**0.64**	--	**-1.09**	**1.07**	-0.18	0.13	--	-0.19	**0.29**	--	**-0.49**	**0.48**	-0.08	0.06	--	-0.09	**-0.46**	0.97	0.48	1879-1913
		(4.65)	--	**(-6.81)**	**(2.77)**	(-0.23)	(1.56)	--	(-0.78)	**(3.78)**	--	**(-7.95)**	**(2.00)**	(-0.24)	(1.37)	--	(-0.78)	**(-5.15)**		[p=0.75]	
Chile	I	**0.74**	0.17	**-0.52**	-0.04	**-0.26**	**0.11**	-0.007	--	0.21	0.13	**-0.42**	**-0.55**	**0.02**	**0.09**	-0.01		**-0.82**	0.98	1.08	1879-1913
		(2.01)	(1.48)	**(-2.75)**	(-0.08)	**(-16.63)**	**(2.15)**	(-0.67)	--	(1.17)	(1.55)	**(-3.63)**	**(-1.72)**	**(0.95)**	**(2.51)**	(-0.64)		**(-5.56)**		[p=0.40]	
	II	**0.78**	0.17	**-0.57**	0.01	**-0.26**	**0.12**	--	--	0.19	0.13	**-0.44**	-0.51	0.02	**0.09**	--	--	**-0.52**	0.98	1.77	1879-1913
		(2.06)	(1.44)	**(-3.12)**	(0.03)	**(-16.02)**	**(2.53)**	--	--	(1.11)	(1.51)	**(-4.01)**	(-1.64)	(1.22)	**(2.86)**	--	--	**(-4.31)**		[p=0.19]	
	III	**0.68**	--	**-0.61**	--	**-0.26**	**0.16**	--	0.05	**0.20**	--	**-0.47**	--	**0.27**	**0.12**	--	0.04	**-0.77**	0.98	1.26	1879-1913
		(3.22)	--	**(-5.37)**	--	**(-16.96)**	**(2.42)**	--	(0.54)	**(2.04)**	--	**(-5.56)**	--	**(1.64)**	**(2.93)**	--	(0.55)	**(-6.49)**		[p=0.31]	
Greece	I	**-2.19**	**0.79**	-0.18	**-1.04**	1.72	**0.31**	--	--	0.15	**0.11**	**-0.21**	**-0.16**	**0.26**	**0.05**	--	--	**-0.15**	0.97	1.64	1879-1913
		(-2.67)	**(2.91)**	(-0.43)	**(-1.76)**	(0.81)	**(2.31)**	--	--	(1.27)	**(5.18)**	**(-4.46)**	**(-2.64)**	**(0.76)**	**(2.31)**	--	--	**(-3.44)**		[p=0.21]	
	II	-0.23	--	**-1.15**	0.34	2.4	**0.32**	--	--	-0.04	--	**-0.19**	0.06	0.43	**0.06**	--	--	**-0.16**	0.96	1.09	1879-1913
		(-0.60)	--	**(-3.22)**	(0.81)	(1.14)	**(1.99)**	--	--	(-0.63)	--	**(-2.80)**	(0.78)	(1.04)	**(2.07)**	--	--	**(-2.67)**		[p=0.37]	
	III	-0.47	--	**-1.03**	--	3.31	**0.31**	--	--	0.16	--	**-0.23**	--	0.54	**0.05**	--	--	**-0.18**	0.96	1.27	1879-1913
		(-0.97)	--	**(-2.69)**	--	(1.60)	**(1.94)**	--	--	(1.25)	--	**(-3.05)**	--	(1.37)	**(1.86)**	--	--	**(-2.98)**		[p=0.31]	

Table 4. ARDL Estimates of Nominal Exchange Rate Determinants (cont.)
(all variables in logs except for interest rates which are in percent a year)
(t-rations in parenthesis)

Dependent Variable: Nominal Effective Exchange Rate (log of Index, 1913=100)

		Long-Run Elasticities								*Short-Run Elasticities*								*Other Statistics*			
		tot	g/t	m3-m3*	y-y*	n-n*	fic	i-i*	edebt/x	tot	g/t	m3-m3*	y-y*	n-n*	fic	i-i*	edebt/x	ECM	Adj-R^2	F-test: Gold Dummies	Period
Italy	I	0.11	-0.08	-0.44	0.21	1.54	0.07	-0.03	--	0.05	-0.04	-0.20	0.02	0.70	0.03	-0.02	0.00	-0.45	0.82	--	1879-1913
		(1.41)	(-1.17)	(-3.00)	(1.75)	(2.79)	(2.87)	(-1.59)	--	(1.34)	(-1.22)	(-3.58)	(0.53)	(3.45)	(3.21)	(-1.56)	(-0.13)	(-4.54)		--	
	II	0.16	-0.09	-0.5	0.28	1.35	0.08	--	--	0.06	-0.04	-0.21	0.05	0.56	0.04	--	--	0.41	0.82	--	1879-1913
		(2.21)	(-1.21)	(-2.96)	(2.13)	(2.35)	(3.22)	--	--	(2.00)	(-1.27)	(-3.62)	(1.15)	(2.99)	(3.70)	--	--	(-4.18)		--	
	III	0.19	--	-0.39	0.14	0.93	0.08	--	0.01	0.09	--	-0.18	0.06	0.43	0.04	--	--	-0.39	0.79	--	1879-1913
		(2.93)	--	(-2.61)	(1.47)	(1.59)	(3.13)	--	(0.28)	(2.55)	--	(-3.03)	(1.42)	(1.98)	(3.79)	--	--	(-4.07)		--	
Portugal	I	-0.13	-0.15	-0.92	1.44	3.80	0.09	--	--	-0.08	-0.1	-1.16	0.46	2.41	0.06	--	--	-0.63	0.94	2.98	1879-1913
		(-1.29)	(-0.71)	(-3.07)	(3.09)	(3.67)	(3.11)	--	--	(-1.37)	(-0.76)	(-3.51)	(2.94)	(3.18)	(2.19)	--	--	(-4.42)		[p=0.41]	
	II	-0.1	--	-0.95	1.42	3.74	0.09	--	--	-0.06	--	-1.24	0.48	2.52	0.06	--	--	-0.67	0.94	3.42	1879-1913
		(-1.17)	--	(-3.41)	(3.29)	(3.89)	(3.20)	--	--	(-1.18)	--	(-3.98)	(3.32)	(3.42)	(2.24)	--	--	(-5.05)		[p=0.29]	
	III	-0.08	--	-1.02	1.40	3.73	0.08	--	0.06	-0.6	--	-1.19	0.45	2.64	0.06	--	--	-0.54	0.94	--	1879-1913
		(-1.02)	--	(-3.66)	(3.44)	(4.09)	(2.78)	--	(1.10)	(-1.03)	--	(-3.78)	(2.94)	(3.55)	(2.09)	--	--	(-3.97)		--	
Spain	I	0.57	0.09	-0.66	0.28	2.06	0.14	--	--	0.23	0.04	-0.69	0.12	0.85	0.06	--	--	-0.41	0.81	--	1879-1913
		(1.06)	(0.38)	(-3.07)	(0.47)	(0.96)	(1.84)	--	--	(.1.20)	(0.37)	(-4.84)	(0.46)	(0.91)	(2.02)	--	--	(-2.67)		--	
	II	0.84	--	-0.69	0.95	--	0.15	--	-0.01	0.35	--	-0.69	0.02	--	0.06	--	--	-0.42	0.81	--	1879-1913
		(1.66)	--	(-3.66)	(3.53)	--	(2.11)	--	(-0.09)	(1.86)	--	(-4.69)	(0.08)	--	(1.90)	--	--	(-2.73)		--	
	III	0.84	--	-0.69	0.95	--	0.15	--	--	0.35		-0.68	0.02	--	0.07	--	--	-0.38	0.82	--	1879-1913
		(1.69)	--	(-3.77)	(3.61)	--	(2.54)	--	--	(1.89)		(-4.86)	(0.08)	--	(2.23)	--	--	(-2.58)		--	
Russia	I	0.38	-0.18	-0.72	0.03	0.56	0.02	-0.01	--	0.15	0.014	-0.4	0.01	-6.04	0.02	-0.01	--	-1.00	0.87	0.61	1887-1913
		(4.84)	(-3.51)	(-6.68)	(0.55)	(9.87)	(2.23)	(-0.95)	--	(2.34)	(0.37)	(-3.19)	(0.35)	(-4.14)	(2.23)	(-0.94)	--	--		[p=0.67]	
	II	0.37	-0.18	-0.69	0.02	0.55	0.02	--	--	0.13	0.025	-0.36	0.03	-5.68	0.02	--	--	-1.00	0.87	0.81	1887-1913
		(4.76)	(-3.67)	(-6.67)	(0.35)	(10.14)	(2.29)			(2.15)	(0.65)	(-3.08)	(0.54)	(-4.04)	(2.29)	--				[p=0.55]	
	III	0.5	--	-0.69	0.18	0.47	0.00		-0.14	0.14	--	-0.35	0.15	-6.51	0.00	--	-0.12	-0.84	0.88	0.82	1887-1913
		(3.53)	--	(-3.14)	(2.27)	(4.83)	(-0.43)	--	(-1.97)	(2.49)	--	(-2.35)	(3.07)	(-4.84)	(-0.42)	--	(-2.77)	(-4.96)		[p=0.55]	

Table 5. Probit Estimates of the Determinants of Currency Crises
(coefficients refer to marginal effects at mean, robust z-statistics in parentheses)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
tot cycle	-0.36							
	(-1.04)							
tot cycle_1	-0.64		2.82	2.05	3.3	3.12	2.34	2.51
	(-1.40)		(2.17)**	(-1.32)	(-1.89)*	(1.98)**	(2.01)**	(1.39)
Δ tot cycle		-0.16						
		(-0.89)						
(m2-m2*)/(y-y*)	1.48							
	(5.94)***							
(m2-m2*)/(y-y*)_1	-1.42		0.20	0.13	0.24	0.23	0.16	0.18
	(5.66)***		(3.27)***	(2.91)***	(2.72)**	(2.81)***	(3.73)***	(2.79)**
Δ(m2-m2*)/(y-y*)		5.45						
		(4.81)**						
g/t	0.34	0.87						
	(3.57)***	(3.41)**						
g/t_1	0.45	1.49	3.8	2.404	4.59	4.34	3.05	2.67
	(3.00)***	(2.45)**	(5.64)***	(5.67)***	(6.18)***	(5.86)***	(9.10)***	(5.22)***
efico	-0.36		-2.06	-1.29	-2.39	-2.13	-1.49	-1.79
	(4.88)***		(6.74)***	(7.01)***	(6.82)***	(7.10)***	(7.52)***	(6.64)***
efico_1	-0.41		2.06	1.288	2.19	2.01	1.5	1.83
	(3.39)***		(3.77)***	(3.74)***	(3.82)***	(3.69)***	(3.92)***	(3.93)***
Δefico		-1.23						
		(4.93)***						
iuk	-0.22							
	(2.68)***							
iuk_1	0.24		0.51	0.61	0.58	0.52	0.45	0.41
	(2.96)***		(-1.74)*	(-1.68)*	(1.54)	(1.51)	(1.74)*	(0.9)
iuk_2			0.25	0.24	0.38	0.34	0.26	0.301
			(-1.29)	(-1.06)	(-1.43)	(1.35)	(1.44)	(2.74)**
Δ_iuk		-0.68						
		(3.01)***						
Reergap_1				3.03				
				(-1.33)				
Ext.Debt/X_1					0.19			
					(-0.84)			
Total Debt/GDP_1						-0.149		
						-0.77		
Trade Balance_1							0.81	
							(1.50)	
Domestic Int. Rate								0.12
								(0.99)
Observations	566	566	567	496	470	499	567	482
Pseudo R-squared	0.50	0.41	0.34	0.34	0.34	0.34	0.35	0.36

*significant at 10%; ** significant at 5%; *** significant at 1%