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STANDARD ERA**

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# Sovereign Risk in the Classical Gold Standard Era\*

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## Abstract

This paper explores the determinants of sovereign bond yields during the classical gold standard period (1872-1913). Using the Pooled Mean Group methodology, we find that the main benefit of the gold standard was as a short-hand device that enhanced a country's reputation in international capital markets. By conveying important information to investors and enhancing the speed of adjustment of sovereign bond spreads to long-run equilibrium levels, the gold standard allowed country risk to be priced more effectively. In contrast to other studies, our results suggest that fundamental factors were more important in determining a country's creditworthiness in the long-run than the exchange rate regime per se.

*Keywords:* Gold standard, sovereign risk, heterogeneous dynamic panels, pooled mean group estimator.

*JEL Classification:* F33, F34, F41, N10, N20.

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# 1 Introduction

Membership of the gold standard is often regarded as a significant factor driving yield spreads on sovereign bonds issued during the period leading up to World War One. A seminal study by Bordo and Rockoff (1996) argues that adherence to the gold standard served as a ‘good housekeeping seal of approval’, shaving as much as 40 basis points off sovereign bond spreads.<sup>1</sup> An important recent contribution, using a much larger country sample and a wider range of macroeconomic variables, by Obstfeld and Taylor (2003) also confirms the role played by the gold standard in enhancing sovereign credibility before 1914. Their findings also indicate that, to a large extent, country risk was priced without much reference to fundamental macroeconomic variables such as the level of public debt and the terms of trade.

While such results suggest that the exchange rate regime plays an important role in the long-run determination of spreads, they shed little light on how it helps a country acquire a reputation for financial probity. A country’s reputation for creditworthiness in previous periods is likely to have a substantial bearing on its current reputation. This may be because the ‘type’ of policymaker running the country could change over time in ways not readily transparent to foreign lenders.<sup>2</sup> Creditors, therefore, are required to constantly update their beliefs about the type of policymakers that they face. With Bayesian updating of country risk, bond spreads are likely to exhibit a high degree of persistence and deviate from steady state levels for a long time. These deviations are likely to vary from country to country, particularly if investors are wary of ‘emerging’ countries at the periphery of the international financial system.

Obstfeld and Taylor (2003) tackle the dynamics of country risk using a lagged dependent variable model with the Arellano-Bond (1991) one-step dynamic panel estimation technique

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<sup>1</sup>A number of country-specific studies find similar results. See, for example, Sussman and Yafeh (2000) and Lazaretou (2005) for the cases of Japan and Greece.

<sup>2</sup>For example, policymaker types could change unobservably over time and a reputation is gradually built as a ‘disciplined’ type separates himself from other types (Mailath and Samuelson, 2001). Alternatively, leadership changes may have a causative effect on economic growth and change perceptions of creditworthiness (see Jones and Olken, 2005).

estimated by the generalised method of moments (GMM). But, as Pesaran et al. (1999) observe, the Arellano-Bond technique can lead to inconsistent estimates in heterogeneous dynamic panels when the time dimension,  $T$ , is large. There is a risk, therefore, of misleading inferences about the true determinants of spreads and the long-run effect of the gold standard on country risk. Moreover, the Arellano-Bond approach does not permit heterogeneity in the short-run adjustment process for yield spreads. So, little can be said about how the gold standard contributed to reputation building by different countries at the ‘core’ and ‘periphery’ of the international financial system.

In this paper, we examine the evolution of sovereign spreads in the pre-World War One era using a similar (though not identical) dataset to Obstfeld and Taylor (2003). We use the Pooled Mean Group (PMG) technique of Pesaran et al. (1999) to derive consistent estimates for the determinants of country risk and examine the (differing) speeds of adjustment towards long-run equilibrium. Our findings are as follows. First, the prime benefit of gold standard membership appears to arise through a more rapid convergence of spreads towards their long-run equilibrium, suggesting that membership conveyed important information to investors that enabled them to price risk more effectively. Second, and in contrast to Obstfeld and Taylor, gold standard adherence was not, by itself, sufficient to drive borrowing terms. Rather, investors appear to have attached significance to fundamental variables such as low inflation, fiscal balances, and levels of economic development when pricing sovereign risk in equilibrium.

Our paper is related to several recent cross-country studies of the gold standard era. Meissner (2005) examines the timing of adoption of the gold standard and shows that countries with large borrowing costs and significant trading relationships with other gold standard members were more likely to fix their exchange rates earlier. Mauro et al. (2002) provide evidence suggesting that investors in the pre-World War One era paid particular attention to country-specific fundamentals when pricing risk. Flandreau and Zumer (2004) argue that the gold standard may have played a much more modest role in the evolution of the global

financial system than hitherto assumed. But their use of feasible generalised least squares (FGLS) with large panel data dimensions also raises the spectre of inconsistent estimates stressed by Pesaran et al. (1999) and Pesaran and Smith (1995).

Ferrucci (2003) also uses the PMG approach to study the determinants of sovereign bond spreads. His focus, however, is on a panel of countries covering the recent period 1992 - 2002. He finds that while bond spreads reflect fundamentals, non-fundamental factors also play a significant role. Unlike the present paper, his analysis of short-run dynamics focuses on the gap between the equilibrium and observed level of spreads and does not consider the role played by the exchange rate regime in influencing the adjustment to equilibrium.

Obstfeld et al. (2005) study the extent to which the gold standard influenced the degree to which countries in the international monetary system followed the world (i.e. the UK) interest rate. Using a similar dynamic equilibrium correction technique to ours (Pesaran et al., 2001), they suggest that the classical gold standard era was a period of limited monetary independence, consistent with the capital market openness of the time. By comparing the adjustment periods of country interest rates to long-run equilibrium periods, they contrast the rapid transmission of interest rate shocks with the greater monetary autonomy of the post-World War Two Bretton Woods era. But their analysis does not consider the fundamental determinants of country risk or how it may have evolved differently in ‘core’ and ‘periphery’ countries.

The structure of the paper is as follows. Section 2 describes the econometric framework and motivates the PMG methodology. Section 3 presents the main results and discusses the preferred parsimonious specification. Section 4 considers the robustness and sensitivity of our results. Section 5 concludes. Details of data sources and diagnostic tests are contained in the Appendix A and B respectively.

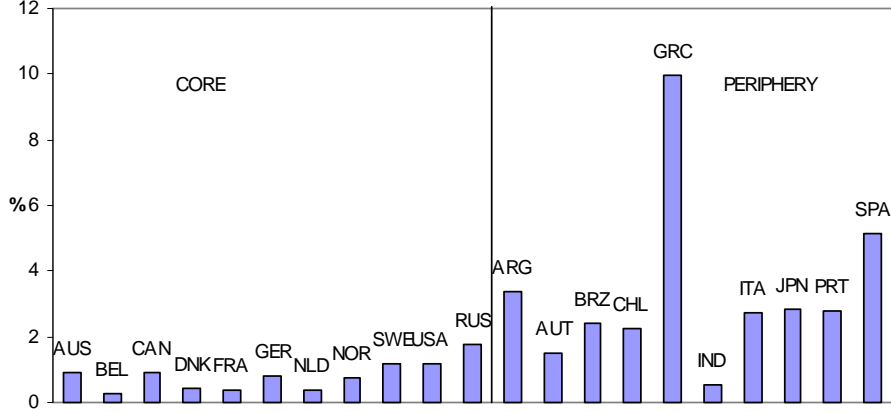
## 2 Econometric Framework

### 2.1 Data

Our data is annual and covers twenty-two countries between 1872 - 1913. The sample begins in 1872 since many countries join the gold standard from the mid-1870s onwards giving a more balanced panel dataset. We follow Obstfeld and Taylor (2003) and separate countries into ‘core’ and ‘periphery’. The ‘core’ countries are predominantly Anglo-Saxon and European, namely Australia, Belgium, Canada, Denmark, France, Germany, the Netherlands, Norway, Russia, Sweden, the United States and the United Kingdom. The ‘peripheral’ or ‘emerging market’ countries in the sample are: Argentina, Austria, Brazil, Chile, Greece, India, Italy, Japan, Portugal and Spain. Our sample differs from Obstfeld and Taylor (2003) in that it excludes South Africa, New Zealand, Turkey, Finland, Hungary, and Mexico, but includes Russia and the Netherlands. The Global Financial Database (GFD) and Mitchell (1992, 1993, 2000) are our primary sources. These data were corroborated against the datasets used by Obstfeld and Taylor (2003) and Flandreau and Zumer (2004). Some missing gaps were filled by consulting other recent studies of the gold standard era, for example, the nominal GDP. We take our measure of country risk to be the bond spread over UK consol yields. The bond yields used are from gold or sterling bonds traded in the London capital market, and their attributes are described in Appendix A along with details of the selected control variables.

Figure (1) illustrates the average cost of external borrowing facing these countries in the pre-World War One period. It is clear that countries belonging to the ‘core’ of the international monetary system experienced much lower yield spreads than their counterparts on the ‘periphery’. The difference in the average spread between the two groups is 261 basis points and there is considerable diversity amongst countries. The ‘core’ countries typically faced borrowing costs that were 74 basis points above the UK rate, while the ‘periphery’ countries faced interest rates around 335 basis points above the UK rate. Figure (2) plots

Figure 1: Average Yield Spreads, 1872 - 1913



sovereign bond spreads for selected ‘core’ and ‘periphery’ countries. They confirm the picture of gradual convergence in bond spreads reported by Obstfeld and Taylor (2003) for the period leading up to 1914. This may be attributable to the rising popularity of the gold standard as a stable and credible regime underpinning the international monetary system and integrating global capital markets (see Figure (3)). From a statistical perspective, it suggests that certain long-run parameters might be identical across groups.

## 2.2 Model

Following Edwards (1986), Bordo and Rockoff (1996) and Obstfeld and Taylor (2003), suppose that yield spreads are determined by the relationship:

$$SPREAD_{it} = \alpha_i + \beta_i WSPREAD_t + \gamma X_{it} + u_{it}, \quad (1)$$

where  $X_{it}$  captures lagged control variables commonly used in country risk analysis such as the inflation rate, the level of real GDP per capita, the M2 to GDP ratio, as well as

Figure 2: Yield Spreads for Selected Core and Peripheral Countries

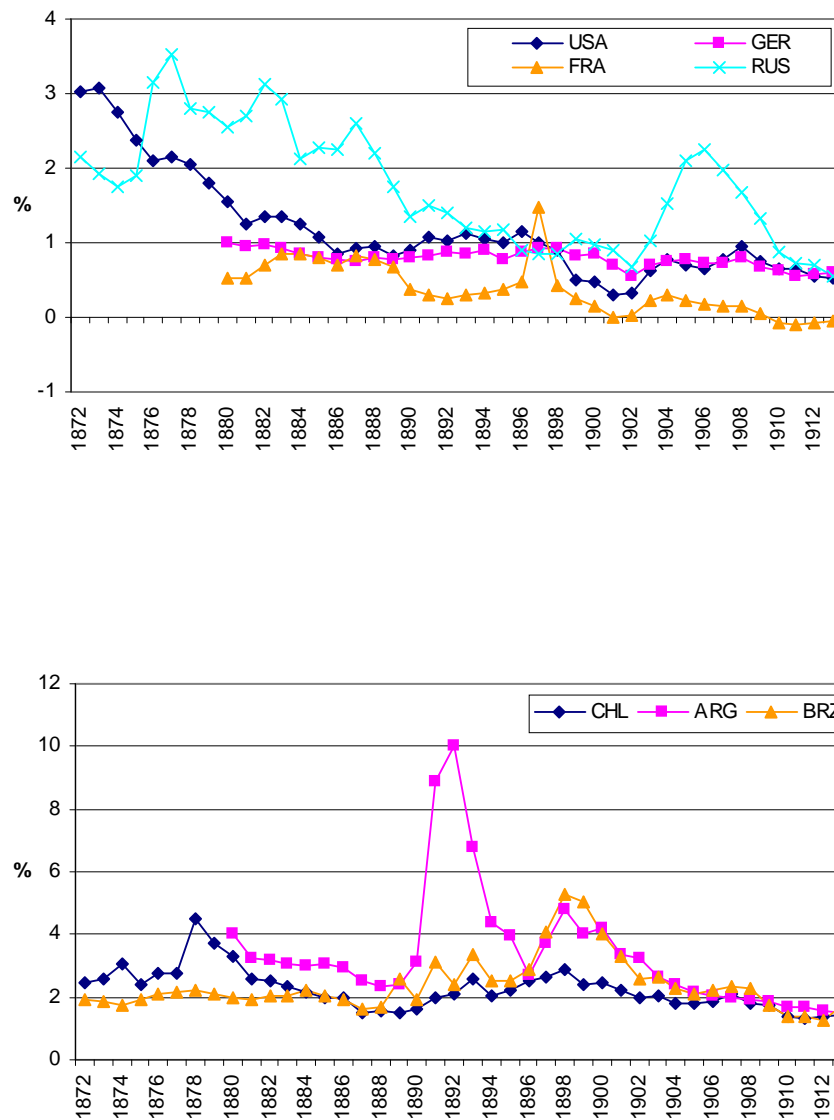
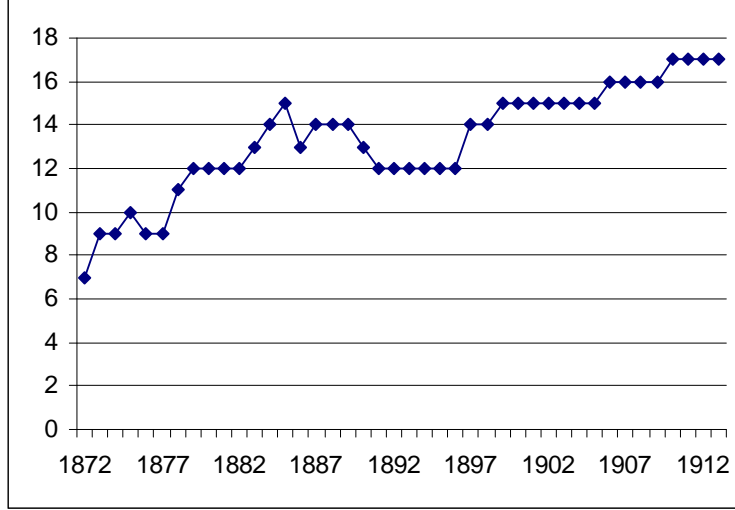


Figure 3: Number of Countries on the Gold Standard, 1872 - 1913



dummy variable for gold standard membership (see Appendix A for details). The variable  $WSPREAD_t$  captures the effects of global interest rate shocks in international capital markets and takes the form:

$$WSPREAD_{it} = \overline{YIELD}_t - YIELD_{UK,t} \quad (2)$$

where  $YIELD_{UK,t}$  is the risk free rate (i.e. the yield on UK consol bonds), and

$$\overline{YIELD}_t = \sum_{i=1}^N \frac{GDP_{it} \times er_{it}}{\sum_{i=1}^N GDP_{it} \times er_{it}} \cdot YIELD_{it}.$$

is the GDP-weighted return on the global bond market portfolio.

The PMG approach to estimating (1) is appropriate for several reasons.<sup>3</sup> First, as explained in Pesaran and Smith (1995) and Pesaran et al. (1999), the procedure yields consistent estimates given the dimension ( $N = 22$  and  $T = 42$ ) of the panel data. Although Arellano and Bond (1991) suggest that the full range of IVs in a GMM procedure helps ad-

<sup>3</sup>Other applications of the PMG approach include Pesaran et al. (1999), Cameron and Muellbauer (2001), and Bassanini and Scarpetta (2002).

dress inconsistency, Pesaran et al. (1999) argue forcefully that the Arellano-Bond technique nonetheless delivers inconsistent long-run coefficient estimates in heterogeneous dynamic panels when  $T$  is large. This is readily illustrated by the fact that different authors derive quite different conclusions on the determinants of yield spreads when using different sub-country and sub-period samples. Second, as a dynamic panel estimation procedure, it explicitly allows for differing speeds of adjustment of the convergence of spreads towards the long-run equilibrium. The different dynamics of convergence of ‘core’ and ‘periphery’ countries suggested by Figure (2) points to the importance of taking this into account. Third, PMG allows estimation of heterogeneous short-run responses of countries to policy changes which could capture the nature of the data.

Our PMG model of sovereign risk is essentially a dynamic equilibrium correction equation of the form:

$$\begin{aligned} \Delta SPREAD_{it} = & \alpha_i + \phi_i SPREAD_{it-1} + \beta'_1 Z_{it-1} + \beta_2 WSPREAD_{t-1} + \\ & \beta_3 GS_{it-1} + \lambda_i \Delta SPREAD_{it-1} + \gamma \Delta WSPREAD_{t-1} + \delta' \Delta Z_{it-1} + u_{it}, \end{aligned} \quad (3)$$

where  $Z_{it-1}$  is the set of explanatory variables,  $\alpha_i$  are individual fixed effects, and  $\phi_i$  the speeds of adjustment to the long-run equilibrium. For lagged dependent variables, we choose a maximum lag of one to preserve sufficient degrees of freedom. The short-run dynamics of the model are governed by the coefficients of  $\lambda_i$ ,  $\gamma$  and  $\delta'$ . As with Pesaran et al. (1999),  $u_{it}$  is assumed to be independently distributed across  $i$  and across  $t$ .<sup>4</sup>

The long-run equilibrium of yield spread for each country is:

$$SPREAD_{it} = - \left( \frac{1}{\phi_i} \right) [\alpha_i + \beta'_1 Z_{it} + \beta_2 WSPREAD_t + \eta_{it}], \quad (4)$$

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<sup>4</sup>Although independence across  $t$  could be violated due to possible serial correlation for long time series, this problem can be addressed by increasing the lagged order on both the dependent and independent variables – we do not follow this approach because of the limited degrees of freedom available to us.

where the  $\eta_{it}$  are stationary processes. For a long-run equilibrium to exist,  $\phi_i \neq 0$ , and for convergence  $\phi_i < 0$ . To obtain the coefficients of the long-run relationships, we follow the assumption of long-run homogeneity discussed in Pesaran et al. (1999). This is done by taking the simple average of  $\phi_i$ 's across countries ( $\bar{\phi}$ ), and multiplying it by  $\alpha_i$ ,  $\beta'_1$  and  $\beta_2$  to obtain:

$$SPREAD_{it} = \theta_{0i} + \theta'_1 Z_{it} + \theta_2 WSPREAD_t, \quad (5)$$

where  $\theta_{0i} = -\frac{\alpha_i}{\bar{\phi}}$ ;  $\theta'_1 = -\frac{\beta'_1}{\bar{\phi}}$ ;  $\theta_2 = -\frac{\beta_2}{\bar{\phi}}$ . In the estimated models of (3) and (4), the vector of explanatory variable enters with a lag order of one. Deviations from equilibrium are possible in the short-run.

In Section 3, we discuss the choice of variables in  $Z_{it-1}$ . Since the choice of suitable explanatory variables in the sovereign debt literature is typically *ad hoc*, we adopt the general to specific approach (Hendry and Richard, 1983) to test the model down to its parsimonious specification. Unless stated otherwise, these variables enter equations (3) and (4) with a timing convention of  $t - 1$ . To address potential endogeneity, we use a lag order of two for the inflation rate and the variables that are expressed as a ratio of nominal GDP. The coding of the gold standard adherence dummy,  $GS_{it}$ , is subjective. To make matters simple, the dummy is assigned the value 1 when countries formally announced their legal commitment to gold (*de jure* status).

## 3 Results

### 3.1 Feasible GLS and PMG Estimation

We begin by initially estimating equation (1) with the feasible generalised least squares method used by Obstfeld and Taylor (2003). We then compare the results against the estimated PMG models of equations (3) and (4). The benefits are two-fold. First, it checks the

quality of our data. Second, it acts as a robustness check of the ‘good housekeeping seal of approval’ hypothesis on a somewhat different set of countries to that used by Obstfeld and Taylor (2003). This may be relevant since some authors (e.g. Flandreau and Zumer, 2004) question the impact of gold standard adherence on yield spreads.

Table (1) presents results for two specifications used by Obstfeld and Taylor (2003). Regressions 1.1 to 1.4 employ variants of the gold standard as the sole explanatory variable. In the first two estimations, we follow Obstfeld and Taylor (2003) and allow for the fact that some countries in the sample may have been in full or partial default. We, therefore, interact the gold standard dummy with a dummy registering defaults to gauge the extent to which investors were able to differentiate between gold standard adherence and debt compliance. The variable  $GS \times no\ default$  is defined as an interaction of the gold standard dummy with a dummy for no default on sovereign bonds, while the variable  $GS \times default$  is the interaction of the gold standard dummy with a dummy for episodes of default. Regressions 1.3 and 1.4 then consider the effects of the gold standard in isolation. By contrast, regressions 1.5 to 1.8 include a range of control variables, such as debt to GDP ratio, inflation, and open economy indicators. Note that the regression results reported in 1.1 and 1.5 are directly reproduced from Obstfeld and Taylor (2003) to facilitate comparison.<sup>5</sup>

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<sup>5</sup>Regression 1.1 is taken from Table 1, regression 3, while regression 1.5 is from Table 2, regression 3.

Table 1: A Comparison of Results with Different Datasets, Gold Standard Definitions and Econometric Methodologies

Methodology	FGLS	FGLS	FGLS	PMG	FGLS	FGLS	FGLS	PMG	FGLS	FGLS	PMG
Specification	1.1	1.2	1.3	1.4	1.5	1.6	1.7	1.8	1.9	2.0	2.1
GS			-0.29 (3.93)**	-0.69 (1.88)+			-0.27 (2.73)**	-0.68 (2.16)*			
GS $\times$ NODFLT	-0.14 (1.96)*	-0.18 (3.09)**			-0.27 (4.05)**	-0.17 (1.91)*					
GS $\times$ DFLT	0.12 (0.40)	-0.28 (0.58)			-0.25 (0.71)	-0.76 (2.17)**					
Default	1.65 (6.72)**	2.23 (6.22)**			0.94 (4.28)**	2.15 (6.97)**					
Partial	0.1 (0.51)	0.31 (0.98)			0.40 (1.95)+	0.76 (3.19)**					
DEBTGDP					0.09 (1.1)	-0.14 (1.57)	-0.13 (1.43)	0.26 (0.54)			
INFL					0.360 (2.45)*	0.0003 (0.21)	0.0004 (0.27)	0.011 (1.08)			
LRGDPPC					-0.45 (4.32)**	-0.52 (4.38)**	-0.53 (4.35)**	-1.75 (2.86)**			
LEXPGRP					-0.11 (0.24)	-0.23 (3.35)**	-0.24 (3.52)**	-0.65 (1.36)			
LTOT					-0.11 (1.03)	0.13 (1.00)	0.10 (0.76)	-0.055 (0.07)			
Observations	895	801	801	759	571	472	472	445			
No. of groups	22	21	19	23	18	17	17	16			

Notes: Absolute value of t-statistics in parentheses

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

Dependent variables:

SPREAD: specifications 1 to 3; 5 to 7.

DSPREAD: specifications 4 and 8.

All regressions contain country's fixed effect and country-specific beta terms, but are not reported.

Regression 1.2 suggests that the coefficient of  $GS \times no\ default$  (-0.18) is comparable to that obtained by Obstfeld and Taylor (2003). Clearly, however, this variable could proxy for other omitted country characteristics. Regression 1.6 shows that, upon adding other variables, the coefficient remains broadly unchanged and is significant at the 5% level.<sup>6</sup> The key determinants of yield spreads for our sample are episodes of debt default, the level of economic development (measured by the log of real GDP per capita), and ability to pay (the log of exports/GDP).

This preliminary experiment is supportive of the findings of Obstfeld and Taylor (2003). As a further check, we consider the *de jure* definitions of the gold standard in isolation instead of jointly accounting for effects of default episodes. In regression 1.3, gold standard adherence lowers yield spreads by 29 basis points. The addition of further control variables in regression 1.7 confirms the robustness of this finding – the level of economic development and ability to pay remain key. Mere changes in the definition of the gold standard do not appear to affect the estimates.

Regressions 1.4 and 1.8 in Table (1) present the results obtained with the PMG estimation method. The coefficients reported are the long-run equilibrium relation between yield spreads and various explanatory variables. As with regression 1.3 based on FGLS, the coefficient capturing the benefits of adopting the gold standard in regression 1.4 is appropriately signed and significant as a stand-alone variable. But the impact on yield spreads (around 70 basis points) is much more marked, however. Upon adding further control variables, the gold standard dummy remains significant (at the 5% level) as does the level of economic development at the 1% level. The good housekeeping effect of the gold standard, therefore, appears to be robust across econometric procedures.

A major shortcoming of the regressions reported in Table (1) is the somewhat *ad hoc* choice of explanatory variables. We therefore use the general to specific approach (Hendry and Richard, 1983) was used to test the model (see equation (3)) down to its parsimonious

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<sup>6</sup>We experimented with dropping the *partial* dummy variable in these specifications. But the results were quantitatively similar.

specification (see Table (2)) using the PMG procedure. These regressions do not include country specific  $\Delta GS$  terms.<sup>7</sup> Since the analysis could be sensitive to the definition of the gold standard in a dynamic setting, we consider three variants. Regression 2.1 uses the simple gold standard dummy. The effect of the gold standard is interpreted as a long-run phenomenon. In regression 2.2, the gold standard variable is interacted with yield spreads,  $GSSPREAD$ . A large and negative coefficient for  $GSSPREAD$  would imply that the adoption of the gold standard shortens the time required for a country to converge to the steady state and reap the benefits of reduced yield spreads. This coefficient could be interpreted as reflecting the learning behaviour of investors as they update their assessments of country risk. Regression 2.3 allows for both the long-run and adjustment effects of  $GS$ .

Several conclusions can be drawn from Table (2). First, there is a significant equilibrium correction term  $\bar{\phi}$  ranging between 0.39 to 0.47. It suggests that yield spreads adjust fairly quickly to their equilibrium levels with an average half life of some 1.5 to 1.8 years following a shock or policy change.

Second, fundamental variables play an important role in spread determination. The level of economic development, as measured by log real GDP per capita, is significant across all specifications. This variable is defined in terms of natural logarithms, so the coefficient associated with it implies that a 1 unit increase in the log of real GDP per capita reduces spreads by 26 basis points (based on regression 2.2).<sup>8</sup> Besides economic development, the ratio of fiscal balances (revenue minus expenditure) to nominal GDP,  $FISCGDP$ , and the inflation rate,  $INFL$  are also significant determinants of yield spreads. A large fiscal surplus will lower a country's cost of borrowing as revenues improve ability to pay as well as serve as collateral. For every 1 percentage point increase in  $FISCGDP$ , yield spreads are lowered by around 7 basis points in the long-run. Countries with a track record of high inflation also face higher borrowing costs – a one percentage point rise in inflation raises spreads by 5 basis

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<sup>7</sup>This is so because most of the countries have adopted the gold standard by the 1900s (see Figure (3)). These coefficients could not be estimated as they have mainly entries of zeros.

<sup>8</sup>Obtained by dividing the coefficient of  $LRGDP$ , -2.015, by the mean of the log real GDP per capita of entire panel data, 7.64.

Table 2: Parsimonious Specification with Various Gold Standard Definitions

<b>Specification</b>	<b>2.1</b>	<b>2.2</b>	<b>2.3</b>
<i>Estimates of Long Run Coefficients</i>			
GS	-0.628 (3.08)**		0.200 (0.39)
WSPREAD	0.350 (1.16)	0.514 (1.48)	0.547 (1.52)
LRGDPPC	-1.761 (3.01)**	-2.015 (3.09)**	-2.086 (3.10)**
FISCGDP	-6.205 (2.67)**	-7.167 (2.69)**	-7.386 (2.70)**
INFL	0.051 (4.45)**	0.054 (4.23)**	0.056 (4.21)**
<i>Average of Short Run Adjustment Coefficients</i>			
$\bar{\phi}$	-0.466 (2.42)*	-0.405 (2.26)*	-0.394 (2.18)*
GSSPREAD		-0.156 (3.45)**	-0.184 (2.12)*
DSPREADL	0.423 (10.16)**	0.427 (10.28)**	0.426 (10.26)**
DWSPREAD	0.464 (2.71)**	0.479 (2.81)**	0.484 (2.83)**
DDINFL	0.011 (4.20)**	0.010 (4.12)**	0.010 (4.11)**
Observations	491	491	491
R-squared	0.37	0.38	0.38
HS Test (F)	3.31	3.82	3.87
p-value	(0.04)	(0.03)	(0.02)
LM $t$ -bar Test, Pooled	-22.61	-22.42	-22.41
Average $t_i$	-4.84	-4.79	-4.79

Notes: Absolute value of t-statistics in parentheses  
+ significant at 10%; \* significant at 5%; \*\* significant at 1%  
LM test: 5% critical value (approx.): -3.96 from Im et al. (2003, Table 2).  
Dependent variable: DSPREAD

points as shown in regression 2.2. Perhaps surprisingly, our results suggest that common shocks to interest rates, *WSPREAD*, do not have a significant impact on spread levels, though the positive coefficient indicates that a 100 basis point increase in the global yield differential with the UK rates raises spreads by around 50 basis points. The lack of perfect correlation with the base rate could reflect a degree of capital immobility or the presence of narrow bands permitting some exchange rate movement between gold points.<sup>9</sup>

Third, short-run dynamic terms are significant determinants of yield spreads.<sup>10</sup> A coefficient of approximately 0.4 on *DSPREADL* points to a fairly high degree of persistence in bond spreads. The size of the coefficient suggests that reputations of creditworthiness in previous periods plays a part on the formation of current reputations and the cost of borrowing. Deviations of a country's yield from world interest rate could also reflect a change in investor's risk assessment about systematic risk. The estimate (of 0.5) for *DWSPREAD* suggests that the transmission of common interest rate shocks to yield spreads is also not instantaneous.

Across the three specifications in Table (2), regression 2.2 is the preferred model. In regression 2.1, the long-run impact of the gold standard on borrowing terms is clear to see. But without an interactive term between *GS* and yield spreads, it is hard to comment on the evolution of country risk. In regression 2.3, which models both long-run and adjustment effects, we would not expect long-run effects to show up strongly if the gold standard was purely a reputational device for investors to learn about a country. As regression 2.3 shows, the long-run coefficient is insignificant (and incorrectly signed), while the adjustment coefficient is significant and of similar magnitude to the one in regression 2.2.

Returning to regression 2.2, the coefficient of the gold standard interacted with the lagged

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<sup>9</sup>See Obstfeld et al. (2005) for a discussion of this issue. In a related study on the determinants of emerging market bond spread in the 1990s, Min (1998) found that external common shock, measured by the three-month US Treasury bill, is not significant. This is also consistent with Antzoulatos (1996) who shows that the US interest rates matter little in the determination of bond flows to the Latin economies in the 1990s.

<sup>10</sup>We constrain the estimates of short run dynamic terms to be identical across countries to preserve sufficient degree of freedom.

yield spreads  $GSSPREAD$  is estimated to be -0.16. On average, a country adopting the gold standard found its borrowing costs converging more quickly to long-run equilibrium levels. In general, ‘peripheral’ countries joining the gold standard experience a more rapid spread convergence than the ‘core’ countries. For the sample as a whole, the imputed half life to achieve lower yield spreads upon adoption of gold standard alone (using regression 2.2)) is 4.4 years.<sup>11</sup> By contrast, with the interaction variable  $GSSPREAD$  of regression 2.2 the half life is 1.8 years.<sup>12</sup> In other words, gold standard adherence may have allowed investors to price country risk much more effectively.

The diagnostic tests for the three models (bottom rows of Table (2)) suggest that the parsimonious specifications may suffer from mild heteroskedasticity problems. As highlighted earlier, the assumption of homoskedastic disturbances gives rise to consistent, but inefficient, estimates. So the presence of potential heteroskedasticity may not be severe enough to invalidate our findings. Furthermore, the p-values for these tests are much too close to us to exercise judgement on the choice of preferred model based on this criterion, especially between regressions 2.2 and 2.3. The LM bar test on the residual shows that there is no serial correlation. As such, a long-run cointegrating relation among the explanatory variables does exist for all three specifications. Regression 2.2 is our preferred specification on grounds of economic intuition and for satisfying the panel cointegration test (which is more crucial amongst the diagnostic tests).

### 3.1.1 Half-lives

Recall that, for the existence of long-run equilibrium, it is necessary to have  $\phi_i \neq 0$ . When  $\phi_i < 0$ , there is convergence to the long-run steady state equilibrium for each country. Table (3) presents the speeds of adjustment  $\phi_i$  and half-life of countries for specifications reported in Table (2). These speeds of adjustment are well behaved in 17 out of 22 countries in

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<sup>11</sup>The value of half-life was imputed from the coefficient of  $GSSPREAD$  in regression 2.2.

<sup>12</sup>This is obtained by  $-\frac{1}{-0.405} \times -0.156$ .

which  $\phi_i \leq 0$ .<sup>13</sup> There is an anomaly for two countries, India and Sweden, where speeds of adjustment are estimated at essentially zero. This could be because the steady state for these two countries is not being well captured by the model, and therefore the speed of adjustment is ill-defined. The overall results are not, in any case, sensitive to the inclusion of India and Sweden.

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<sup>13</sup>Countries with  $\phi_i = 0$  are not reported in Table (3), and these are Canada, Chile and Greece.

Table 3: Speeds of Adjustment and Half-lives of Selected Countries

Specification Country	2.1			2.2			2.3		
	$\theta_i$	t-stat.	Half-life	$\theta_i$	t-stat.	Half-life	$\theta_i$	t-stat.	Half-life
ARG	-0.497	-8.67	1.39	-0.494	-8.75	1.40	-0.490	-8.57	1.41
AUS	-0.662	-0.56	1.05	-0.570	-0.49	1.22	-0.552	-0.47	1.26
AUT	-0.560	-1.38	1.24	-0.587	-1.46	1.18	-0.595	-1.47	1.16
BEL	-0.491	-0.82	1.41	-0.371	-0.62	1.87	-0.350	-0.58	1.98
BRZ	-0.425	-3.73	1.63	-0.418	-3.75	1.66	-0.410	-3.61	1.69
DNK	-1.140	-1.10	0.61	-1.008	-0.97	0.69	-0.986	-0.95	0.70
FRA	-0.754	-2.63	0.92	-0.620	-2.15	1.12	-0.597	-2.03	1.16
GER	-0.995	-1.09	0.70	-0.918	-1.01	0.76	-0.909	-1.00	0.76
IND	0.110	0.08	-	-0.116	-0.08	5.98	-0.197	-0.14	3.52
ITA	-0.187	-1.09	3.71	-0.172	-1.01	4.03	-0.165	-0.96	4.20
JPN	-0.501	-3.88	1.38	-0.461	-3.73	1.50	-0.444	-3.38	1.56
NLD	-0.387	-0.75	1.79	-0.252	-0.49	2.75	-0.227	-0.44	3.05
NOR	-0.418	-1.31	1.66	-0.256	-0.80	2.71	-0.229	-0.69	3.03
PRT	-0.355	-9.61	1.95	-0.358	-9.72	1.94	-0.358	-9.72	1.94
SPA	-0.303	-3.46	2.29	-0.307	-3.53	2.26	-0.308	-3.53	2.25
SWE	-0.031	-0.05	22.36	0.128	0.20	-	0.159	0.25	-
USA	-0.397	-2.07	1.75	-0.263	-1.39	2.64	-0.230	-1.11	3.01
RUS	-0.399	-1.52	1.74	-0.245	-0.93	2.83	-0.214	-0.78	3.24
Core (average)	-0.567	-1.19	3.40	-0.500	-0.98	1.84	-0.477	-0.89	2.02
Peripheral (average)	-0.389	-3.86	2.03	-0.364	-4.00	2.49	-0.371	-3.92	2.22

Notes:

 $\theta_i$  The speed of adjustment of individual country to shocks in the long-run equilibrium relationship.

t-stat. The t-statistic associated with the speed of adjustment.

Half-life The half-life of a shock (in years), computed based on the adjustment speed.

- Negative half-life not reported.

Overall, the results from Table (3) support the hypothesis that there exists a homogeneous long-run relationship across the countries. A small absolute value of the adjustment speed translates to a higher half-life where a country takes longer to converge to the steady state. A common-sense check should indicate that ‘peripheral’ countries take longer to converge compared with ‘core’ countries. Across specifications, the results show that borrowing costs in Argentina took 1.4 years on average to converge; Brazil took around 1.7 years and Italy took approximately 4 years. Meanwhile, yields in ‘core’ countries such as France and Germany took around 0.7 and 1 year on average to converge. Based on regression 2.2, the ‘core’ countries on average have a speed of adjustment of -0.50 which translates to a half-life of 1.8 years, whereas a ‘peripheral’ country has a speed of adjustment of -0.36 which yields a half-life of 2.5 years.

Another noteworthy feature is that Austria and Spain are estimated to have a relatively low half-life even though they did not adopt the gold standard regime at all during the period. This reinforces the findings from Table (2) which show that the gold standard serves as a signaling device in the short-run and that fundamentals matter more in the long run.

### 3.2 Parsimonious Specification

We now document the steps involved in obtaining the parsimonious specification of regression 2.2 in Table (2).<sup>14</sup> For short-run dynamic terms, second differences were considered.<sup>15</sup> These measure the acceleration of the chosen macroeconomic variables, and so attempt to reflect any surprises in trends. Since some of these variables could be I(2), taking second order differences is a conservative practice that creates stationary variables and ensures consistent estimates.

Regression 3.1 in Tables (4) and (5) considers all plausible explanatory variables in the dataset. In addition to those considered in Table (2), we also included: a measure of political

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<sup>14</sup>Regressions 2.1 and 2.3 were tested down in similar manner.

<sup>15</sup>These are debt to GDP ratio, inflation, fiscal balance to GDP ratio, log of export to GDP, custom revenue to GDP, M2 to GDP and log of terms of trade. We also experimented with first differences and obtained quantitatively similar results.

conditions *POLITY2*, the level of financial development *M2GDP*, and customs revenue as ratio of GDP, *CUSREVGDP*. The coefficient of *GSSPREAD* is estimated to be -0.08 which seems plausible (see regression 3.1). In regression 3.2, *DDFISCGDP*, *DDINFL*, *DDLTOT* and *DDM2GDP* were first dropped as they were highly insignificant. *POLITY2*, an index measuring the degree of democracy, was estimated to have a positive, but insignificant coefficient - so we remove the variable. *DDCUSREVGDP* appears to have the wrong sign. The accumulation of custom revenues should typically reduce the cost of a country's borrowing rates in the gold standard era. *DDLTOT*, an open economy indicator for export competitiveness, was also dropped as it was insignificant.

In regression 3.3, we continue to remove wrongly signed and insignificant variables from regression 3.2. We now focus on the long-run equilibrium coefficients. The measure of financial development reflected by *M2GDP* is negative but not significant. A high level of indebtedness of a country, *DEBTGDP*, might be expected to lead to higher borrowing costs. Again, the coefficient is not significant. Finally, open economy indicators like *LEXPGDP* and *LTOT* were also generally insignificant.<sup>16</sup> In regression 3.4, we remove *CUSREVGDP* which correlates positively but not significantly with yield spreads. Since higher custom revenues should enhance ability to repay, a positive effect on yield spread seems counter-intuitive.

High inflation might indicate poor economic performance and signal prospective payments difficulties. In general, one would expect a positive relation between lagged inflation (measured by annual change in CPI levels) and changes in yield spreads. Using the PMG estimator, the coefficient of inflation rate is unstable in regressions 3.1 to 3.4. In regression 3.5, we drop the inflation rate to examine the implications. Since the magnitude and significant of the *GSSPREAD* variable increases, it suggests potential signs of omitted variable bias. So, in regression 3.6, we restore the inflation rate in the regression. A percentage point increase in inflation adds 5 basis points to yield spreads.

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<sup>16</sup>These variables are also not significant in Obstfeld and Taylor.

Table 4: PMG Estimates of Long Run Coefficients: General to Specific

Specification	3.1	3.2	3.3	3.4	3.5	3.6
GSSPREAD	-0.078 (1.94)+	-0.042 (1.32)	-0.114 (2.53)*	-0.094 (2.14)*	-0.182 (4.19)**	-0.156 (3.45)**
WSPREAD	1.243 (3.97)**	1.367 (3.58)**	0.574 (1.35)	0.092 (0.32)	0.686 (2.23)*	0.514 (1.48)
DEBTGDP	-0.739 (1.65)	-1.136 (2.32)*				
INFL	-0.021 (1.12)	-0.020 (0.82)	0.032 (1.23)	-0.004 (0.48)		0.054 (4.23)**
LRGDPPC	0.239 (0.55)	0.662 (1.20)	-0.638 (0.89)	-1.283 (2.45)*	-1.248 (2.38)*	-2.015 (3.09)**
FISCGDP	-1.224 (0.43)	-2.821 (1.04)	-7.275 (2.27)*	-5.522 (2.55)*	-7.582 (2.88)**	-7.167 (2.69)**
ERVAR	2.244 (1.44)	1.925 (0.99)	8.106 (3.36)**	10.031 (7.53)**		
LEXPBGDP	0.058 (0.12)	0.052 (0.10)	-0.490 (1.03)			
CUSREVGDP	9.685 (1.36)	17.813 (2.22)*	8.813 (1.21)			
M2GDP	-0.055 (0.28)	-0.120 (0.49)				
WARINTER	-0.422 (1.89)+					
WARINTRA	-0.126 (0.30)					
LTOT	1.877 (2.48)*					
POLITY2	0.018 (0.82)					
Observations	311	393	470	496	548	491
R-squared	0.38	0.30	0.35	0.43	0.34	0.38
HS Test (F)	2.18	0.51	1.88	3.00	2.79	3.82
p-value	0.11	0.60	0.15	0.05	0.06	0.03
LM $t$ -bar Test						
Pooled	-17.58	-19.15	-21.84	-22.31	-24.93	-22.42
Average $t_i$	-4.26	-4.01	-5.30	-4.77	-5.07	-4.79

Notes: Absolute value of t-statistics in parentheses  
+ significant at 10%; \* significant at 5%; \*\* significant at 1%  
LM test: 5% critical value (approx.): -3.96 from Im et al. (2003, Table 2).  
Dependent variable: DSPREAD

Table 5: PMG Estimates of Short Run Coefficients: General to Specific

<b>Specification</b>	<b>3.1</b>	<b>3.2</b>	<b>3.3</b>	<b>3.4</b>	<b>3.5</b>	<b>3.6</b>
DSPREADL	-0.133 (2.32)*	-0.112 (2.20)*	0.419 (9.92)**	0.442 (11.06)**	0.434 (11.08)**	0.427 (10.28)**
DWSPREAD	0.801 (5.48)**	0.759 (6.35)**	0.501 (3.02)**	0.384 (2.35)*	0.44 (3.01)**	0.479 (2.81)**
DDDEBTGDP	0.142 (0.80)	0.204 (1.40)				
DDINFL	-0.005 (1.78)+	-0.004 (1.46)	0.008 (2.14)*			0.010 (4.12)**
DDFISCGDP	-0.154 (0.21)					
DERVAR	1.415 (3.40)**	1.126 (3.09)**	2.270 (3.69)**	2.906 (5.32)**	0.131 (0.33)	
DDLEXP GDP	-0.076 (0.55)	-0.086 (0.84)				
DDCUSREVGDP	1.063 (0.55)					
DDM2GDP	-0.044 (0.53)					
DDLTOT	0.200 (1.01)					

Notes: Absolute value of t-statistics in parentheses  
+ significant at 10%; \* significant at 5%; \*\* significant at 1%  
LM test: 5% critical value (approx.): -3.96 from Im et al. (2003, Table 2).  
Dependent variable: DSPREAD

## 4 Robustness

To further substantiate our results in regression 2.2 from Table 2, we consider additional robustness and sensitivity checks. Specifically, we vary the degrees of long- and short-run heterogeneity.

In the general PMG setup, the basic specification (3) matches a short-run dynamic term to every long-run coefficient. By adopting the general to specific approach, we restrict the coefficients of some of these short-run dynamic terms to be zero. We also restrict some of the short-run dynamics to be identical across countries. We first add  $DFISCGDP$  into the regression since the long-run level effect of fiscal balance to GDP ratio is significant. But this does not change the result quantitatively. Since  $FISCGDP$  could be a fairly non-stationary process that is often related to business cycle properties, we consider  $DDFISCGDP$ . But this does not change the outcome significantly. As such, fiscal balance is only important in the long-run.

The coefficient of average  $WSPREAD$  is not close to one in regression 2.2 of Table 2. In using the variable, we have assumed that the relation between  $WSPREAD$  and yield spreads remains homogeneous across countries. This may not be realistic given that different countries could have different long-run  $\beta$  terms. So we relax the assumption on homogeneity in the long-run effects of  $WSPREAD$ . In doing so, we obtain point estimates of  $\beta$ s for each country. Ideally, one would expect these  $\beta$  terms to be high for risky countries and low for the more developed countries. We do obtain  $\beta$  terms greater than one for risky countries like Argentina (338 basis points) and Portugal (400 basis points). Whereas, the  $\beta$  terms are less than one for the less risky countries like Belgium (41 basis points) and USA (24 basis points). We also obtain negative  $\beta$  terms for some countries, notably Brazil and Denmark. While it is conceivable that investments in these countries might pay off when the world portfolio does not, it should be noted that these coefficients may be imprecisely estimated – particularly since the introduction of additional variables lowers the effective degrees of

freedom.<sup>17</sup>

As a next step, we add, in turn,  $DSPREAD_i$ ,  $DWSPREAD_i$ ,  $DINFL_i$  and  $DERVAR_i$  to evaluate the effects of heterogenous short-run dynamic terms. This again attempts to identify these effects separately compared to Table 2 where they are restricted to be the homogenous effects. In these separate experiments, the coefficient of core economic variables (i.e. the level of economic development, fiscal balance to GDP ratio and exchange rate variability) remains stable and robust across specifications.

We then add  $WSPREAD_i$  and estimate with the sequential addition of  $DSPREAD_i$ ,  $DWSPREAD_i$ ,  $DINFL_i$  and  $DERVAR_i$  until we reach the general form of PMG specification (3). The gold standard is found to reduce yield spreads by 5.2 to 6.5 basis points during the adjustment process. These variables were marginally insignificant around the 10% level, reflecting the large number of coefficient estimated. This also applies to the rest of the key explanatory variables as well. Given limited data at annual frequency, it is best to fix at least some short-run dynamic terms to be homogeneous across countries.

## 5 Conclusion

This paper has examined the determinants of sovereign yield spreads during the classical Gold Standard era between 1872 and 1913, in a way that takes into account the dynamics of investor perceptions of country risk. Specifically, the Pooled Mean Group methodology is used to consistently estimate differing speeds of adjustment across a panel of twenty-two countries. One rationalisation of these differing speeds of response is that international investors gradually updated their beliefs about the creditworthiness of different countries at different speeds.

Our findings are supportive of the conventional ‘seal of approval’ hypothesis about Gold Standard membership. But by using the PMG methodology, we find that membership of

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<sup>17</sup>As Özler (1989) observes, *betas* estimated with other variables are not strictly *betas* in the CAPM sense. Obstfeld and Taylor (2003) do not report their *beta* terms, so a comparison of results is not readily possible.

the gold standard did not have a direct impact on sovereign spreads to the exclusion of other factors. Rather, the gold standard served as a short-hand device that facilitated a country's reputation in international capital markets. Investors remained concerned about country-specific fundamentals, such as inflation and the fiscal balance, but relied on the gold standard to convey this information more swiftly. By enhancing the speed of adjustment of sovereign spreads to their long-run equilibrium levels, gold standard membership facilitated the efficient pricing of sovereign risk.

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## A Data

The dataset covers the period 1872 to 1913. The twenty-two countries are Argentina (ARG), Australia (AUS), Austria (AUT), Belgium (BEL), Brazil (BRZ), Canada (CAN), Chile (CHL), Denmark (DNK), France (FRA), Germany (GER), Greece (GRC), India (IND), Italy (ITA), Japan (JPN), the Netherlands (NLD), Norway (NOR), Portugal (PRT), Russia (RUS), Spain (SPA), Sweden (SWE), United States (USA) and United Kingdom (GBR). The primary source of data is the Global Financial Database (GFD). The data has also been corroborated against Flandreau and Zumer (2004), Obstfeld and Taylor (2003), and Mitchell (various issues) for consistency. Where not available, other independent studies have been used to fill in missing series and gaps in observations. Variables in logarithm form are prefixed with “ $L$ ” in front of the variable’s name. Lagged variables are suffixed by “ $L$ ” at the end. See Table (6) for summary statistics.

Table 6: Summary Statistics of Key Variables by Countries

Country	SPREAD	ER	INFL	FISCGDP	M2GDP	DEBTGDP	LEXPGRP	LTOT	GDPGBP	LERVAR	LRGDPPC
ARG	3.36	11.71	4.81	-0.038	n.a.	0.58	-1.33	4.61	163.26	-3.36	7.99
AUS	0.93	1.02	0.28	0.001	0.30	0.76	-1.56	4.62	381.01	-4.42	8.42
AUT	1.51	18.56	0.43	-0.012	0.47	0.79	-1.60	4.66	475.83	-3.96	7.86
BEL	0.29	25.62	-0.05	-0.017	0.44	0.45	-1.12	4.64	212.61	-4.66	8.12
BRZ	2.39	16.14	4.75	-0.026	n.a.	0.73	-1.25	4.78	127.68	-2.72	6.64
CAN	0.93	4.86	2.99	-0.008	0.22	0.38	-1.77	4.52	192.79	-4.61	7.78
CHL	2.23	14.19	n.a.	0.242	0.17	0.45	-1.12	4.58	49.95	-2.16	7.72
DNK	0.45	18.43	0.35	-0.005	0.55	0.19	-1.56	4.61	63.87	-7.03	7.86
FRA	0.38	25.57	0.11	-0.006	0.28	0.96	-2.03	4.59	1203.78	-4.58	7.83
GBR	0.00	1.00	-0.44	-0.001	0.55	0.50	-1.73	4.48	1549.26	0.00	8.32
GER	0.79	20.72	0.84	-0.030	0.45	0.41	-1.87	4.58	1296.14	-4.61	7.89
GRC	9.96	19.71	1.52	-0.033	0.34	1.44	-1.86	4.63	40.02	-3.60	7.39
IND	0.54	14.16	0.51	-0.007	0.04	0.19	-2.43	4.62	898.45	-3.13	6.46
ITA	2.74	26.73	0.12	-0.013	0.28	1.05	-2.29	4.97	480.20	-3.93	7.44
JPN	2.85	7.73	4.86	0.051	0.22	0.39	-2.40	4.60	256.79	-3.45	6.99
NLD	0.36	12.27	-0.20	-0.020	0.23	0.79	-0.01	n.a.	121.16	-4.31	8.11
NOR	0.73	18.43	0.52	-0.008	0.49	0.21	-1.80	4.50	51.77	-9.25	7.41
PRT	2.78	5.27	0.66	-0.007	n.a.	0.78	-3.26	4.59	133.14	-3.29	7.21
RUS	1.74	9.74	0.92	-0.018	0.15	0.66	-2.52	n.a.	1142.37	-3.42	7.11
SPA	5.14	29.18	0.40	-0.001	0.15	0.97	-2.28	4.54	349.23	-3.46	7.62
SWE	1.15	18.43	0.44	-0.000	0.59	0.17	-1.71	4.60	104.41	-9.25	7.74
USA	1.17	4.93	0.41	0.021	4.09	0.82	-0.83	4.62	3534.32	-4.61	8.20

Note: n.a. - not available

## **Yield Spread (*SPREAD*)**

The main source is from the GFD. Unless stated otherwise, the yield to maturity are for long term government bonds, usually with a maturity of at least ten years. For missing observations, Flandreau and Zumer (2004), and Obstfeld and Taylor (2003) were consulted.

Argentina: 1880 - 1913 from Flandreau and Zumer (2004).

Australia: 1874 - 1913 from GFD. Based on NSW Government 5% Terminable 1874/1902 (1858 - 87), NSW Government 4% Funded Stock 1912 option (1887 - 1900), all NSW and Commonwealth Government issues maturing more than six months (1901 - 25).

Austria: 1879 - 1913 from GFD. The bonds quoted are the Gold 5s.

Belgium: 1880 - 1913 from Flandreau and Zumer (2004).

Brazil: 1872 - 1913 from GFD. Based on the 5s through 1886, the 4.5% Gold Bonds (1887 - 99) and the 4.5% Bonds of 1883 (1900 - 14).

Canada: 1872 - 1913 from GFD. Based on the 5s (1860 - 74) and the 4s (1874 - 1912).

Chile: 1872 - 1913 from GFD. Based on the 6s (1870 - 74), the 5s (1875 - 86) and the 4.5s (1887 - 1928).

Denmark: 1895 - 1913 from Flandreau and Zumer (2004). The bonds quoted are the 1893 3%.

France: 1880 - 1913 from Flandreau and Zumer (2004).

Germany: 1880 - 1913 from Flandreau and Zumer (2004).

Greece: 1872 - 1913 from GFD. Based on the 5s of 1824/1879 (1863 - 87) and the Monopoly 4% (1887 - 1924).

India: 1872 - 1913 from GFD. Based on India 5s of 1880 (1872 - 73), 4s of 1888 (1874 - 80) and 3.5s (1881 - 1940).

Italy: from GFD. 1872-1908: Italy Maremmiana Railroad 5s payable in sterling pounds and guaranteed by the Italian Government.

Japan: from GFD. Based on the 9s (1872), the 7s (1873 - 97) and the 5s (1898 - 99) and the 4s Japanese sterling bonds (1900 - 14).

Netherlands: 1880-1913 from Flandreau and Zumer (2004).

Norway: 1872 - 1913 from GFD. The bonds quoted are the 4.5s (1875 - 80), the 4s (1881 - 86), the 3.5s (1887 - 93) and the 3s (1894 - 1918).

Portugal: 1872 - 1913 from GFD. Based on the 3s through 1895, 1s (1895 - 1902) and the 3s (1903 - 1913).

Spain: 1872-1913 from GFD. Based on the 3s (1836 - 80), 1s (1881), and the 4s (1882 - 1913).

Sweden: 1872 - 1913 from GFD. Based on the 5s (1868 -78), the 4s (1879 - 94) and the 3s (1894 - 1914).

United Kingdom: 1872 - 1913 from GFD. The bond quoted is the 2.5% consol bonds.

United States: from Friedman and Schwartz (1982:122 - 3).

Russia: 1872 - 1913 from GFD. Based on the 5s (1823 - 1914).

The following table presents the first-order autocorrelations of yields:

#### **Exchange Rate (*ER*)**

All exchange rates are from the GFD. The original series defines exchange rate to be the price of domestic currency per unit price of US dollars. Here, the year-end value is used and the numeraire converted to sterling pounds.

#### **Gold Standard (*GS*)**

The gold standard dummy takes the value of one when the country has officially announced adoption of *de jure* gold standard, otherwise it takes the value of zero. Experimentation was done interacting this dummy variable with full and partial default episodes. These were taken from Obstfeld and Taylor (2003). The main source of gold standard coding is from Bordo and Schwartz (1996), and corroborated against Meissner (2005). The coding for the following countries were found to have discrepancies amongst various authors.

Austria: Flandreau and Zumer (2004) suggest Austria was on gold in 1896. Meissner (2005), and Obstfeld and Taylor (2003) code this country as not having adopted the gold standard until 1925.

Table 7: Autocorrelations of Yield Spreads

<b>Country</b>	<b>Autocorrelation Coefficients</b>
Argentina	0.734
Australia	0.826
Austria	0.553
Belgium	1.020
Brazil	0.749
Canada	0.816
Chile	0.798
Denmark	0.957
France	0.670
Germany	0.902
Greece	0.920
India	0.834
Italy	0.884
Japan	0.851
Netherlands	0.858
Norway	0.864
Portugal	0.749
Russia	0.916
Spain	0.901
Sweden	0.881
United States	0.892
United Kingdom	0.976
Average (core)	0.882
Average (peripheral)	0.797

Brazil: Meissner (2005) excludes 1888 and 1889. We cross-check with Flandreau and Zumer (2004) and find a value for 1889.

Norway: In Braga de Macedo et al (1996), gold was adopted in 1875. The observation for 1873 was obtained from Broz (2002). Thus, we record a dummy with value of one from 1873 onwards.

Spain: We use data from Meissner (2005).

### **Inflation** (*INFL*)

Inflation is measured as the year-on-year change in the CPI index. 1913 is used as the base year. Data are collected from GFD, and missing series are taken from Flandreau and Zumer (2004).

Table 8: Years in which Countries Joined the Gold Standard

<b>Country</b>	<b>Period</b>	<b>Membership*</b>
Argentina	1870 - 75; 1883 - 85; 1899 - 1913	Peripheral
Australia	1870 - 1913	Core
Austria		Peripheral
Belgium	1878 - 1913	Core
Brazil	1888 - 89; 1906 - 13	Peripheral
Canada	1870 - 1913	Core
Chile	1887; 1895 - 98	Peripheral
Denmark	1872 - 1913	Core
France	1878 - 1913	Core
Germany	1871 - 1913	Core
Greece	1885; 1910 - 13	Peripheral
India	1899 - 1913	Peripheral
Italy	1884 - 94	Peripheral
Japan	1897 - 1913	Peripheral
Netherlands	1875- 1913	Core†
Norway	1873 - 1913	Core
Portugal	1870 - 90	Peripheral
Russia	1897 - 1913	Core†
Spain		Peripheral
Sweden	1873 - 1913	Core
United States	1879 - 1913	Core
United Kingdom	1870 - 1913	Core
Notes:		* based on classification in Obstfeld and Taylor (2003). † based on authors' classification.

### **Fiscal Balance/GDP (*FISCGDP*)**

Fiscal balance = government revenue – government expenditure. The main source is Mitchell (various issues). Government revenue refers to the current revenues exclusive of loans, with their main tax constituents. Government (total) expenditures includes capital items which are normally financed by borrowing. As argued in Mitchell, the most reluctant omission has been any disaggregation of government expenditure. But this heterogeneous data changes in nature frequently, and is often not available in a meaningful form.

### **Money Supply**

M1 is defined as the sum of bank note in circulation and deposits in commercial banks. M2 is the sum of M1 and deposits in savings. Source: Mitchell (various issues).

Australia: For currency/banknotes in circulation, the Australian dollar did not replace the pound as the currency unit until the mid-1960s (2 dollars = 1 pound), the present currency unit employed was in the sources used. Statistics to 1900 are at approximately the end of December, and from 1901 to 1945 at the end of June. For demand deposits in commercial banks, figures are the currency deposits of trading banks, and are at approximately the end of December to 1900 and at the end of June from 1901 to 1949. For savings deposits, original statistics are given in pounds. Australian dollars have been converted via the abovementioned exchange rate. Deposits at the end of the financial year of each bank are dated mostly 30 June.

Austria: 1872-1913: National Bank issue (million gulden and million kroner, 2 kroner = 1 gulden). 1872 - 1895 was originally in gulden. This series was converted to kroner. From 1896 onwards, units in millions kroner. For deposits in savings banks, both regulated and post office savings banks have been summed up.

Belgium: For deposits in savings banks, both *Caisse Generale d'Epargne* and post office savings banks have been summed up.

Denmark: Units in million rigsdaler. Deposits in savings are from all savings banks.

France: Units in million francs. For deposits in savings banks, both private savings banks and national savings banks have been summed up.

Germany: Deposits in savings banks.

Greece: Deposits in savings banks.

India: From 1872 to 1873, figures are notes issued. From 1874 to 1913, figures are notes in circulation. For demand deposits in commercial banks, figures are all non-government savings in class A and exchange banks from 1872 to 1913. For savings deposits, figures to 1896 relate to government savings banks, and subsequently they are of the Post Office Savings Bank.

Italy: For deposits in savings banks, both private savings banks and post office savings banks have been summed up.

Japan: For currency/banknotes in circulation, figures to 1959 relate to government notes,

national bank notes, Yokohama Specie Bank, and Bank of Japan notes (exclusive of those held as reserves for other banks' issue. For demand deposits in commercial banks, figures to 1945 are of non-government current account deposits of national banks, special banks (from 1880), private ordinary banks (from 1893), and savings banks (from 1896).

Netherlands: For deposits in savings banks, both general savings banks and post office savings banks have been summed.

Portugal: There are missing values for 1876 and 1877. Linear interpolation was used to fill up the gaps.

Spain: Deposits in savings banks are from all savings banks.

Sweden: For deposits in savings banks, both private and post office savings have been summed up. M2 is the sum of money and quasi-money. Observations from between 1873 and 1875 were interpolated to arrive a value of M1 for 1974. .

United Kingdom: For deposits in commercial, data is only available for the UK. For deposits in savings banks, both trustee savings banks and post office savings banks have been summed up. Thus for currency circulation, the UK data series was used.

Russia: For deposits in savings banks, this reflects all savings banks.

### **Total nominal public debt (*DEBT*)**

The main source is Flandreau and Zumer (2004) for 1880 - 1913. These debt figures have a wide definition - covering both short and long term debt. Missing series are supplemented by the following.

Australia: From Obstfeld and Taylor (2003).

Canada: From Stateman's Year Book (various issues).

Chile: From Obstfeld and Taylor (2003).

India: From Obstfeld and Taylor (2003).

Japan: From Obstfeld and Taylor (2003).

United Kingdom: From Mitchell (1962:403). Table on Aggregate Gross Liabilities of the State: 1870-1913.

United States: From Obstfeld and Taylor (2003).

Russia: From Feis (1930). Units: million rubles and measures the directly incurred debt of the Russian government. The author also provides per cent held by foreigners.

### **Export**

Data collected from GFD's Annual Files on World Trade. Note that these figures correspond to the export of manufactured goods. Original units in national currency.

### **Terms of Trade (*TOT*)**

From Obstfeld and Taylor (2003) dataset. Missing for Netherlands and Russia.

### **Nominal GDP (*GDP*)**

Sources include GFD, Mitchell (various issues) and Flandreau and Zumer (2004). Missing series supplemented by:

India: from Goldsmith (1983).

Portugal: Real GDP from Mata and Valerio (1996). This original series is converted into nominal terms using the GDP deflator (1994 = 100).

### **Exchange Rate Volatility (*ERVAR*)**

Exchange rate variability is constructed using methodology following Flandreau and Zumer (2004). The raw exchange rate (end of month) data was collected from GFD which provides daily data in terms of open, high, low and close. The closing price of exchange rates was used.

$$ERVAR_{it} = \frac{\sum_{t=1}^{12} \ln\left(\frac{ER_t}{ER_{t-1}}\right)}{12} \times 10,$$
$$\text{where } \ln\left(\frac{ER_t}{ER_{t-1}}\right) = \begin{cases} \ln\left(\frac{ER_t}{ER_{t-1}}\right) & \text{if } \ln\left(\frac{ER_t}{ER_{t-1}}\right) > 0 \\ 0 & \text{if } \ln\left(\frac{ER_t}{ER_{t-1}}\right) < 0 \end{cases}$$

### **Real GDP per capita (*RGDPPC*)**

From Maddison (1995). Similar to Obstfeld and Taylor (2003), linear interpolation is used as the frequency of observations is not regular. The years, in which there are observations,

are reported.

<b>Country</b>	<b>Observations</b>
Argentina	1870, 1890, 1900-1913
Australia	1870-1913
Austria	1870 - 1913
Belgium	1870 - 1913
Brazil	1870, 1890, 1900 - 13
Canada	1870 - 1913
Chile	1900 - 13
Denmark	1870 - 1913
France	1870 - 1913
Germany	1870 - 1913
Greece	1913
India	1870, 1890, 1900 - 13
Italy	1870 - 1913
Japan	1870, 1885 - 1913
Netherlands	1870 - 1913
Norway	1870 - 1913
Portugal	1870, 1890, 1900, 1913
Spain	1870, 1890, 1900 - 13
Sweden	1870 - 1913
United Kingdom	1870 - 1913
United States	1870 - 1913
Russia	1870, 1890, 1900, 1913

### **Custom Revenue** (*CUSREV*)

Custom revenue is a component of central government revenue. Sources: Mitchell (various issues).

### **Political Democracy** (*POLITY2*)

The version of data used is from POLITY IV project available from <http://weber.ucsd.edu/~kgledits/Polity.html>. Polity IV contain coded annual information on regimes and authority characteristics for all independent states (with greater than 500,000 total population) and covers the period 1800 – 2002. Originally, political democracy is measured by ‘Polity’ taking values between -10 and 10, where -10 denotes high autocracy and 10 denotes high democracy. When ‘polity’ could not function properly, it is represented by (-88) for transition period; (-77) for periods of interregnum (collapse of central political authority);

and (-66) for a period of interruption. However such classification is cumbersome for time series analysis. As such, the new variable *POLITY2* has been adopted, which adjusts these dummies to generate a smoother time series.

### **War**

War dummies are taken from the Correlates of War 2 (COW2) database available from <http://cow2.la.psu.edu/>. *WARINTER* refers to inter-state war while *WARINTRA* refers to intrastate war.

## **B Diagnostic Tests**

This appendix explains various diagnostic tests essential for the PMG procedure. Given that this is a historical dataset and there are missing observations, this unbalanced panel dataset is a challenge for the implementation of standard diagnostic tests. Instead of brute-force applications, the following discusses how some tests were modified.

### **B.1 Heteroskedasticity**

As discussed in Baltagi (1995), the assumption of homoskedastic disturbances in the presence of heteroskedasticity gives consistent but not efficient estimates. This means that the standard errors of estimates will be biased, leading to inaccurate statistical inference. The problem can be rectified by robust standard errors like in Obstfeld and Taylor (2003). But these authors stop short of testing whether heteroskedasticity exists in the first place. The presence of heteroskedasticity is assumed.

One of the tests for panel level heteroskedasticity involves conducting panel specific iterative GLS procedure on the disturbance terms.<sup>18</sup> This estimates panel level heteroskedasticity and saves the likelihood function. Next, the estimated disturbances of the model are assumed to be homoskedastic. A log-likelihood ratio (LR) test is performed to test these likelihood

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<sup>18</sup>See <http://www.stata.com/support/faqs/stat/panel.html>.

functions with the degrees of freedom determined by the number of constraints imposed, that is, the number of groupings. When this test was applied to the data, the estimation under panel level heteroskedasticity yielded positive log-likelihood values. This reflects the fact that the log-likelihood function is a probability density function instead of a probability value. On the other hand, the log-likelihood estimate for homoskedastic errors is negative. This results in unconventional negative LR test statistics.

In what follows, panel level heteroskedasticity is tested following the White test:

$$\hat{u}_{it} = \beta_1 \hat{y}_{it} + \beta_2 \hat{y}_{it}^2$$

If the null hypotheses for  $\beta_1 = \beta_2 = 0$  are not rejected, the disturbances are homoskedastic.

## B.2 Unit Root Tests

The common tests are Levin and Lin (1992) and Im, Pesaran and Smith (2003, IPS hereafter). Maddala and Wu (1999) compare these tests with the Fisher's tests and favor this simple test to the abovementioned tests in terms of practical use. One of the restrictions of the Levin and Lin test, and the IPS test is that they are only applicable to balanced panel dataset. Similar to the tests here on heterogeneity, synthetic versions of unit root tests are devised, based on discussion in Levin-Lin and IPS. The Levin-Lin test was conducted by pooling OLS regressions to test for serial correlation, the following is estimated via OLS technique:

$$x_{it} = \rho x_{it-1} + \alpha_i + \epsilon_{it},$$

where  $x_{it}$  is the variable of interest, and  $H_0 : \rho = 0$ . Rejection of the null hypothesis means the variable is non-stationary and is  $I(1)$  in levels. To test for stationarity in differences,

$$\Delta x_{it} = \rho x_{it-1} + \alpha_i + \epsilon_{it}.$$

Acceptance of the null hypothesis suggests the variable is stationary and is  $I(0)$  in difference. Table (9) shows the Levin-Lin tests for the set of variables considered in our regressions. The tests could not reject the null of panel unit root for log of terms of trade, log real GDP per capita and log nominal GDP. One of the limitations of the Levin-Lin test is that  $\rho$  is assumed to be the same for all groups, that is, all countries' variables converge at the same rate. Since the dataset here includes countries with diverse experiences, it is not surprising that unit root is violated for growth related variables.

As the Levin-Lin test is restrictive, the data is tested with the IPS test (see Table (10)) which relaxes the assumption that  $\rho_1 = \rho_2 = \dots = \rho_N$ . Instead of pooling the data, separate unit root tests were performed on the  $N$  countries in the dataset. A test statistic based on the average of an augmented Dickey-Fuller regression is computed. This is also known as the LM  $t$ -bar test. Furthermore, this method is superior to Levin-Lin since it allows for residual serial correlation, and heterogeneity of dynamics and error variances across groups. For both tests, only fixed effects without time trend are considered. Lag length was fixed at one.

Table 9: Dynamic Panel Unit Root Tests, Levin and Lin (1992)

<b>Variables</b>	<b>Levels</b>	<b>Differences</b>
SPREAD	68.21	-7.24
INFL	5.06	-22.36
FISCGDP	35.28	-8.97
M2GDP	74.23	-7.86
DEBTGDP	58.69	-5.85
LEXPGRP	93.11	-5.36
LTOT	415.87	-0.91*
LRGDPPC	983.38	1.64*
LGDP	461.17	0.54*
WSPREAD	98.31	-4.34

Note: \* not significant at 5% level.

From Table (10), the null of panel unit root is not rejected for M2 to GDP ratio, debt to GDP ratio, log real GDP per capita and log nominal GDP. Due to the heterogenous nature of the IPS alternative hypothesis (that is, each  $\rho_i$  is different), IPS (2003:74) suggest a rejection of the null hypothesis does not necessarily imply that the null is rejected for all groups. So,

the presence of serial correlation may not affect the PMG analysis severely.

Table 10: Dynamic Panel Unit Root Tests, Im et al. (2003)

<b>Variables</b>	<b>Levels</b>	<b>Differences</b>
SPREAD	10.12	-2.47*
INFL	0.61	-5.48*
FISCGDP	3.49	-3.71*
M2GDP	20.44	-1.50
DEBTGDP	12.35	-1.33
LEXP GDP	8.82	-2.12*
LTOT	9.24	-2.61*
LRGDPPC	44.60	-1.31
LGDP	35.01	0.47
WSPREAD	7.46	-2.35*

Note: Critical values based on IPS (2003, table 2)

10%: -1.78; 5%: -1.85

\* significant at 5%

### B.3 Panel Cointegration

Pedroni (1999) is the standard test but is designed for a balanced panel. Again, in view of the unbalanced nature of the data, testing is done for the non-stationarity of residuals of each regression. This takes the form of an augmented Dickey-Fuller test on the estimated residuals following Cameron and Muellbauer (2001). Rejection of the null hypothesis suggests the absence of serial correlation, and the existence of a long-run equilibrium.