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Emerging from the War: Gold Standard Mentality, Current Accounts and the International Business Cycle 1885-1939

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Abstract

We study international business cycles and capital flows in the UK, the United States and the Emerging Periphery in the period 1885-1939. Based on the same set of parameters, our model explains current account dynamics under *both* the Classical Gold Standard *and* during the Interwar period. We interpret this as evidence for Gold Standard mentality: the expectation formation mechanism with respect to major macroeconomic variables driving the current account – output, exchange rates and interest rates – has remained fundamentally stable between the two periods. Nonetheless, the macroeconomic environment changed: Volatility increased generally, but less so for international capital flows than for GDP. This pattern is consistent with shocks in the Interwar period becoming more persistent and more global.

JEL-Codes: F32, F36, F40, F41, N1

KEYWORDS: Current Accounts, Capital Flows, Business Cycles, Great Depression, Gold Standard, Emerging Markets, Present-Value models

1 Introduction

After World War I, the return to the prosperity of the pre-war period – the "return to normal"[1] – seemed to require the re-introduction of the Gold Standard as the primary goal of monetary policy : "Gold standard belief placed the maintenance of the currency's gold value as the

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[1]See Nurkse, 1944, p. 7

highest priority for monetary policy; Gold Standard rhetoric argued that doing so would provide a stable system adjusting naturally to economic changes in order to recover and maintain equilibrium." (Mouré, 2002, p. 271). This "Gold Standard mentality" (Eichengreen and Temin, 2000) is our point of departure. While this mentality led to continuity in terms of the policy framework, it is also well known that the nature of the shocks hitting the economy changed dramatically in this period (Eichengreen, 1992).

These ideas are well established as a historical narrative. Surprisingly, however, there is very little formal econometric evidence on the relative importance of continiuity and change in the transition from the classical Gold Standard to the Interwar period. We provide such evidence in this paper. We focus on a set of eight economies in the period 1885-1939 for which uninterrupted data is available: Australia, Canada, Italy, Japan, Sweden, the United Kingdom and the United States. This group of countries offers wide variation with respect to the roles they played as (developed) center and (emerging) periphery countries during the respective Gold Standard arrangements of the era before 1913 and during the Interwar period.

To capture Gold Standard mentality empirically, we focus on the stability of agents' expectation formation mechanisms with respect to key macroeconomic variables. We document the stability of this mechanism using an empirical implementation of the present value model of the current account (PVMCA).[2] As a forward-looking variable, the current account should contain information about the expectations agents hold concerning our variables of interest: interest rates, real exchange rates, output. The current account should also contain information about how common across countries and how persistent shocks to these variables are. We therefore follow the approach advocated by Campbell and Shiller (1987) and include the current account as an additional regressor in a VAR that contains our variables of interest. We than show that the same set of VAR (and deep structural) parameters allows us to replicate the dynamics of the current account in both the classical Gold Standard and the Interwar periods.

The fact that this simple model is able to replicate current account dynamics in both periods is an indication that the expectation formation mechanism – and with it macroeconomic transmission – indeed have remained largely constant. It is also interesting for an-

[2]See e.g.Otto (1992),Bergin and Sheffrin (2000),Hoffmann (2001*b*), Nason and Rogers (2006)and Kano (2009, 2008)for empirical implementations of this type of model.

other reason: as has been widely discussed, the classical Gold Standard was characterized by a period of unprecedented international capital mobility and of financial globalization that led to huge net international capital flows (see Obstfeld and Taylor (2004)). Obstfeld (2004) refers to these unidirectional flows as 'development finance' to suggest that these capital flows should predict future growth in the receiving economies.[3] It is exactly this type of globalization – through development finance – that is stressed in the simple bond-only economies which underlie the PVMCA. This makes it particularly interesting to confront the intertemporal approach with historical data from the classical Gold Standard period. Again, the fact that the same model also replicates current account data even after this globalization period had ended, is an indication that key aspects of the macroeconomic transmission mechanism seem to have remained quite stable across our entire sample period (1885-1939).

One interesting aspect of this stability is that current account imbalances at the onset of the Great Depression, contain considerable information for the international spread of the crisis. Estimating our models up to 1931, we are able to show that the countries with the lowest output growth expectations – as implied by their current account positions – eventually did experience the most severe downturns.[4]

While these findings highlight the enormous stability of the fundamental mechanisms shaping macroeconomic expectations, they do not mean that the economic environment has remained stable. Quite to the contrary, we document important changes in some key business cycle moments between the Gold Standard and Interwar periods: first, as other have done before, we show that the volatility of most macroeconomic aggregates increases in the Interwar period. Importantly, however, the volatility of the current account across the board has increased much less than that of GDP.

From a theoretical perspective, the intertemporal approach predicts that these changes should be particularly informative about changes in the nature of the underlying shocks to the economy: on the one hand, a series of papers (e.g. Glick and Rogoff (1995), Hoffmann (2001*b*), Hoffmann (2003)and Kano (2009)) have emphasized that global shocks should not affect the currrent account for the average country. Hence, changes in the relative variabil-

[3]Conversely, he refers to the present-day pattern of globalization - characterized by huge cross-holdings of financial assets – as 'diversification finance' since it mainly enables countries to diversify macroeconomic risk.

[4]We emphasize that this does not mean that the model predicts the Great Depression itself, which the literature has largely found to be non-predictable (See e.g.Dominguez, Fair and Shapiro (1988), Hamilton (1992) and Obstfeld and Taylor (2003) for prominent contributions). Our finding is about the relative position of countries, conditional on the common shock of the Great Depression.

ity of output and the current account may contain information with respect to the relative importance of global and country-specific shocks. On the other hand, shocks to output are only smoothable to the extent that they reflect cyclical variation. If the random-walk component in GDP becomes more important, this also may lower the variability of current account relative to net output. We propose a novel identification procedure that exploits the heteroscedasticity implied by the increased volatility of macroeconomic aggregates in the Interwar period and links it with cointegrating information to show that both mechanisms – more global and more persistent shocks – have contributed to the pattern we see in the data.

Hence, while our results suggest that a simple intertemporal model of the current account is in principle able to explain the data for our entire sample period, they also suggest that the structure of shocks hitting economies has changed dramatically after the war. In particular, it seems that the Interwar period is characterized by a prevalence of much more global shocks. These shocks, however (unlike in the prewar period), are not cyclical in nature. Rather, the global factors at work during the Interwar period reflect continual structural change. This pattern is reminiscent of a stylized fact that has been prominently documented for modern day emerging market economies (Aguiar and Gopinath, 2007*a*): as in emerging market economies today, the cycle in the world economy emerging from the war is the trend.

2 Current Accounts and Business Cycle Patterns in the Pre-WWII Period

The key message we wish to convey in this paper is summarized in Figures 1 and 2: i) a simple intertemporal model of the current account in which there are no a-priori restrictions to international capital mobility or other frictions in financial markets, explains the current account in virtually all economies in our sample for the entire period 1885-1939. ii) at the same time, key business cycle moments have changed between the pre-war period of the classical Gold Standard and the Interwar period. As we will argue, the first part of this findings is evidence of the fundamental stability in the formation of expectations about key macroeconomics drivers of the current account, including expectations of future output and exchange rates across the two apparently fundamentally different macroeconomic regimes of the Classical Gold Standard period and the Interwar period. The second part of our findings, however, suggests that this fundamental continuity is accompanied by secular change

in the structure of the underlying macro-economic shocks.

The first part of this claim is illustrated in Figure 1, where the red line presents the current account of the seven economies in our sample: Australia, Canada, Italy, Japan, Norway, Sweden, the United Kingdom and the United States. The blue (dashed) line gives the prediction from our model as we will discuss it in detail below. The model is estimated based on the entire sample period from 1885-1939 and seems to do a remarkable job in replicating actual current account patterns. This suggests a considerable amount of stability in our model across what would usually to be considered as secular break, i.e. the First World War (WWI) and the breakdown in international capital mobility during the war and in the Interwar period.

Figure 2 illustrates the second part of our claim by showing that key business cycle moments have changed: Specifically, the figure reports the relative standard deviations of the current account and (net) output, i.e. GDP less investment and government spending along with the predictions from our model. Across the board, there is a decline in this ratio between the period of the Classical Gold Standard (1885-1913) and the Interwar period (1919-39). This stylized fact will be our point of reference in identifying *how* the structure of shocks has changed between the periods: current accounts can react only to 'smoothable' shocks. Therefore, shocks that affect all countries symetrically or that change the size of the random-walk component in net output can increase the variabilty of net output without increasing the variability of current accounts in the same proportion. Note that the current accounts predicted by our (reduced-form) model replicate this drop in the relative volatility of the current account.

To get an impression of the robustness of the stylized facts presented in the figures we also use GDP growth data that are detrended using an HP-filter with a smoothing weight of 100. Standard deviations of the filtered data are displayed in Table 1. The first impression is that business cycles in the Interwar period are much more volatile. Notably, the fluctuations of per capita GDP become more volatile in the period after WWI, except for Australia. The same is true for net output and the current account to net output ratio, with the exception of Japan. Note also however, that the current-account / net output ratio generally becomes relatively less volatile than output, consistent with the findings in Figure 1.

These stylized facts represent our point of reference. We now look at them through the lens of a simple present-value model of current-account behavior (PVMCA).

3 Theoretical and econometric setup

3.1 Theoretical background

As the theoretical backdrop for our analysis we use a simple intertemporal model of the current account in which the representative consumer maximizes

$$\sum_{t=0}^{\infty} \beta^t E_0 \left[\frac{X\left(C_{Nt}, C_{Tt}\right)^{1-\gamma}}{1-\gamma} \right]$$

where C_N is non-tradeables consumption, C_T is tradeables consumption and $X(.)$defines a Cobb-Douglas consumption bundle. The intertemporal budget constraint is

$$B_t = (1 + r_t^W)B_{t-1} + Y_t - I_t - G_t - C_t$$

where $B_t$is the stock of foreign assets, $r_t$the world real interest rate and $Y_t, I_t,$ and G_t denote real output, investment, government consumption and $C_t$denotes private consumption expenditure expressed in terms of tradeable goods, i.e.

$$C_t = C_{Tt} + PC_{Nt}.$$

Here, P is the relative price of non-tradeable goods. In this model, the current account balance is given by

$$CA_t = \Delta B_t = r_t^W B_{t-1} + NO_t - C_t$$

where r_t^W is the world interest rate and where we have introduced the notation $NO_t = Y_t - I_t - G_t$ to denote net output, i.e. the national cash flow available for consumption in period t.

Imposing the usual transversality constraint, this law of motion for the current account can be solved forward, to yield the non-linear intertemporal budget constraint.

$$B_{t-1} = \sum_{k=0}^{\infty} E_t \left\{ R_{t+k} \left[C_{t+k} - NO_{t+k} \right] \right\}$$

where $R_{t+k} = \left[\prod_{l=0}^{k} (1 + r_{t+l}^W) \right]^{-1}$. We follow Kano (2008) and log-linearize this expression to

obtain a formula for the current-account / net output ratio

$$\frac{\widetilde{CA_t}}{NO_t} = b\widetilde{r}_t^W + c\sum_{k=1}^{\infty}\kappa^k E_t\left\{\Delta\widetilde{c}_{t+k} - \widetilde{r}_{t+k}^W\right\} + \sum_{k=1}^{\infty}\kappa^k E_t\left\{\widetilde{r}_{t+k}^W - \Delta\widetilde{no}_{t+k}\right\} \tag{1}$$

Here, Δno and Δc are the growth rates of net output and consumption expenditure respectively and the tilde denotes deviations from the unconditional mean. The parameters b, c, are the long-term means of B/NO, C/NO respectively and $\kappa = exp\left[E(\Delta no_t) - E(r_t)\right]$. Note that the approximation above follows directly from the intertemporal budget constraint. The condition is therefore consistent with arbitrary processes for investment and output and would also hold in a production economy.

Also, we have not yet made use of our specific assumptions on the form of utility or the presence of traded and non-traded goods. We now do so by assuming that $X_t = C_{Tt}^{\alpha} \times C_{Nt}^{1-\alpha}$ is a unit-elasticity-of-substitution aggregate of traded and non-traded goods where α is the expenditure share of traded goods. It is well known that in this case the intertemporal allocation of consumption can be solved for independently from the intratemporal allocation of consumption between tradeable and non-tradeable goods. Specifically, we can define a price index of aggregate consumption by recognizing that for any such index P^*it must be true that $P_t^* X = C_{Tt} + P_t C_{tNt} = C_t$ for all P_t. Substituting for X in the utility function we obtain the first order condition

$$E_t\left(\left(\frac{C_t}{C_{t+1}}\right)^{\gamma}\left(\frac{P_t^*}{P_{t+1}^*}\right)^{1-\gamma} \times \left(1 + r_{t+1}^W\right)\right) = 1 \tag{2}$$

As shown in Obstfeld and Rogoff (1995) and Bergin and Sheffrin (2000), the aggregate price index for consumption is an expenditure-weighted CES aggregate of the tradeable and non-tradeable goods prices so that $P_{t+1}^*/P_t^* = \left(P_{t+1}/P_t\right)^{1-\alpha}$. Hence, (2) links aggregate consumption expenditure growth to the consumption-based real interest rate, which is the world-real interest rate corrected for real exchange rate changes (defined as the change in the relative price of non-traded goods). Assuming that consumption growth, the real exchange rate, and the real interest rate are jointly log-normal, Bergin and Sheffrin (2000) show that this condition can now be log-linearized to obtain

$$E_t(\Delta c_{t+1}) = \frac{1}{\gamma}E_t\left(r_{t+1}\right) + const.$$

where $r_{t+1} = r^W_{t+1} + (\gamma - 1)(1-\alpha)\Delta p_{t+1}$ is the consumption-based real interest rate.

We can use this expression for expected consumption growth to impose more structure on the log-linearized budget constraint (1). Plugging in from the previous equation and rearranging, we obtain the solution for the current-account / net output ratio that is the focus of our empirical analysis here:

$$\frac{\widetilde{CA}_t}{NO_t} = b\widetilde{r^W_t} + \left[1 - c\left(1 - \frac{1}{\gamma}\right)\right] \sum_{k=1}^{\infty} \kappa^k E_t \widetilde{r^W}_{t+k} + c\left[1 - \frac{1}{\gamma}\right] \sum_{k=1}^{\infty} \kappa^k E_t \widetilde{\Delta q}_{t+k} - \sum_{k=1}^{\infty} \kappa^k E_t \Delta\widetilde{no}_{t+k} \tag{3}$$

where $\Delta q_t = (1-\alpha)\Delta p_t$ denotes the change in the real exchange rate.

The first term in (3) measures the impact of net asset income on the current account: an increase in the world interest rate (or a depreciation of the real exchange rate) increases the value of non-tradeable factor income from abroad. *Ceteris paribus*, the current account of a debtor country will deteriorate following an increase in the world interest rate or a real depreciation wheres that of creditor country will improve. The second and third terms are consumption tilting terms: first, an increase in the world real interest rate above its long-run mean lowers consumption today and increases the current account. Second, an expected appreciation of the real exchange rate increases the future relative price of non-tradeables. With $1/\gamma < 1$, this equally provides an incentive to save tradeable goods, increasing the current account. We refer to the first channel as global tilting (because it is driven by global variation in interest rates) and to the second as domestic tilting (since it is driven by idiosyncratic variation in the consumption-based real interest rate, i.e. by real exchange rate changes). Clearly, the more willing the representative household is to substitute consumption today for consumption tomorrow (i.e. the higher is the intertemporal elasticity of substitution), the stronger will be the global tilting effect. Finally, the last term is the typical consumption smoothing term in these models: if the sum of future expected output increases is positive, this should induce the country to borrow in order to increase consumption to its permanent level.

In analyzing our historical data set we take guidance from several properties of the model above: first, to the extent that the world interest rate is covariance-stationary and that net output and the real exchange rate are integrated of order no higher than one, the above equa-

tion implies that the current account should be covariance-stationary itself. This is a special instance of cointegration in the context of present-value model as first noted by Campbell and Shiller (1987) and it is worth emphasizing here since it has an important bearing on the econometric specification of the model. We will discuss this issue in the next section.

Secondly, the model implies that the current account should add very important information concerning agents' expectations about our variables of interest: real exchange rates, (net) output and the world real interest rate. Including CA/NO in a VAR for these variables should therefore help us assess the (reduce-form) stability of the expectation-formation mechanism.

Third, the empirical literature on the PVMCA has emphasized the importance of distinguishing between country-specific and global shocks. [5] Clearly, countries will only be able to smooth fluctuations in net output through borrowing and lending to the extent that these are country-specific. In the context of our model, that should imply that global (expected) variation in interest rates should not affect current account variability for the average country.[6] As shocks become more global, this could increase the variability of output without increasing the variability of the current account to the same extent. Furthermore the model implies that current accounts should react only to variation in the *predictable* components of our variables of interest. For example, an increase in the variability of the random walk component of net output or real exchange rates could increase the volatility of the latter without increasing the variability of CA/NO to the same extent. As we will argue, both mechanisms help explain the general decline in the relative standard deviation of the current account and net output that we documented in Figure 2.

[5] See Glick and Rogoff 1995; Hoffmann 2001*a*,*b*, 2003; Nason and Rogers 2002; Engel and Rogers 2006.

[6] Note that in our analysis, this does not preclude the possibility that global interest rate shocks could affect a country's current account through international interest payments as well as through its impact on intertemporal substitution– the global tilting term — and the effect that interest rates will have on the present-value of future cash-flows. However, in global equilibrium this will be possible only to the extent that countries' initial net foreign asset positions or their reaction to a common shock are at least somewhat heterogeneous.

4 Empirical implementation

We study the empirical dynamics of the world interest rate, the current account and net output in a vector auto regressive model (VAR):

$$\mathbf{A}(L)\begin{bmatrix} r^W_{t+1} \\ CA_t/NO_t \\ q_t \\ no_t \end{bmatrix} = \epsilon_t \tag{4}$$

where $A(L)$ is a 4×4 matrix polynomial in the lag operator with no roots inside the unit circle and ϵ_t is a 4×1vector of white noise.

It is well-known that present-value relations such as (3) impose cointegrating restrictions on the data. In the present setup, we assume that the world real interest rate is stationary and that $no_t$and q_t are integrated of order at most one (*I(1)*). [7] Then, (3) implies that CA/NO, as the discounted sum of expected future realizations of a process that is integrated of order zero (I(0)) , is equally I(0). These restrictions allow us to interpret (4) as a cointegrated VAR with two trivial cointegrating relations – the current-account (i.e. CA/NO) and the world interest rate are themselves stationary so that the cointegrating space is spanned by the first two unit vectors.

We can then write the level-VAR in error correction form so that

$$\mathbf{\Gamma}(L)\Delta X_t = \boldsymbol{\alpha}\boldsymbol{\beta}' X_{t-1} + \epsilon_{\mathbf{t}} \tag{5}$$

where

$$\boldsymbol{\beta} = \begin{bmatrix} 1 & 0 \\ 0 & 1 \\ 0 & 0 \\ 0 & 0 \end{bmatrix}$$

is the matrix of cointegrating vectors, $\boldsymbol{\alpha}$ is a vector of adjustment loadings and $\mathbf{\Gamma}(L)$ is a lag polynomial with all roots outside the unit circle.

Our empirical analysis is based on this VECM-specification. Once $\boldsymbol{\beta}$ is known, the other parameters can straightforwardly be estimated by OLS. Most earlier analyses of the present-

[7] In fact, these assumptions are necessary for the log-linearization leading up to (3)

value model of the current account (with the exception of Hoffmann 2001*a,b*, 2003) have worked with a mixed levels differences specification of the VAR in which the stationary variable appears in levels and the non-stationary variable in differences. The advantage of working with the VECM-formulation is that the long-term dynamics of the cointegrated system can easily be expressed in closed form in terms of the three parameter matrices $\mathbf{\Gamma}(L)$, $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$. Specifically, as we will show, the permanent and transitory shocks to the system can directly be inferred from knowledge of ϵ_t and the adjustment loadings $\boldsymbol{\alpha}$. Clearly, this is particularly convenient in our setting here since our interest is in studying the impact of shocks of different orders of persistence on the current account and net output. An additional advantage of exploiting the cointegrated structure of the model in this way is that the just-identification of structural shocks – to the extent that they can be classified as either permanent or transitory – is determined by the data themselves, thus requiring the researcher to impose fewer *a priori* restrictions from economic theory. We will illustrate these points in turn. Before turning to the identification of structural shocks, however, we assess the reduced-form fit of our model.

4.1 Data description

We analyze annual data in the observation period 1885-1939. The countries under analysis are Australia, Canada, Japan, Norway, Sweden, United Kingdom, and United States. The main data source is Jones and Obstfeld (2001) (current account, GDP, fixed investment)[8] and Backus and Kehoe (1992) (government consumption, prices),[9] population data are from Maddison (2004). As proxy for the world interest rate, we use the discount rate for the United Kingdom.[10] Real effective exchange rates are calculated as trade weighted averages of the real exchange rates vis-à -vis the partner countries. For the Japanese real effective exchange rate, we use the data from Shimazaki and Solomou (2001). For the other countries, we determine the main trading partners based on the availability of direction of trade statistics in Mitchell (2003*b*,*c*,*a*) ,Table E.2, and calculated the weights as averages of import and export weights.[11]

[8]http://www.nber.org/databases/jones-obstfeld/.

[9]dge.repec.org/BK92.html.

[10]NBER Macro History Database, /www.nber.org/databases/macrohistory/, file m13016.csv.

[11]Australia: Japan, UK, USA; Canada: Germany, Japan, UK, USA; Japan; China, UK, USA; Norway: Canada, Denmark, France, Germany, Netherlands, Sweden, UK, USA; Sweden: Denmark, France, Germany, Netherlands, Norway, UK, USA; UK: Argentina, Australia, Canada, France, Germany, India, Netherlands, New Zealand, Russia, USA; USA: Canada, France, Germany, Japan, Mexico, UK.

The consumer price indices are mainly from Mitchell (2003*b*,*c*,*a*), Table H.2;[12] The exception is India, where we use Mukherjee (1969), Table A2.11. The nominal exchange rates are from Schneider, Schwarzer and Denzel (1991, 1992, 1994, 1997). For Norway and Sweden, we use the data base provided by the Riksbank and the Norges Bank.[13]

4.2 Fit of the present value model

To assess the general fit of our model, we follow the approach by Campbell and Shiller (1987) and use the estimated VAR to back out the expectations on the right hand side of the current account equation (3). We rewrite the VECM in companion form as[14]

$$Z_{t+1} = \boldsymbol{G} Z_t + \boldsymbol{u}_{t+1}$$

where Z_t is the vector of current and past realizations of ΔX_t and $\boldsymbol{\beta}' X_t$, $\boldsymbol{G}$ the associated companion matrix and $\boldsymbol{u}_{t+1}$ a disturbance term.[15] We then use the Hansen-Sargent prediction formula to proxy the expectations on the right hand-side of the current-account equation (3). Specifically, we have

$$E_t \{r_{t+i}\} = \mathbf{e}'_{\boldsymbol{r}} \boldsymbol{G}^{\boldsymbol{i}} Z_t,\ E_t \{\Delta q_{t+i}\} = \mathbf{e}'_{\boldsymbol{\Delta q}} \boldsymbol{G}^{\boldsymbol{i}} Z_t \text{and } E_t \{\Delta no_{t+i}\} = \boldsymbol{e}'_{\boldsymbol{\Delta no}} \boldsymbol{G}^{\boldsymbol{i}} Z_t$$

where e_r, $e_{\Delta q}$ and $e_{\Delta no}$ are the unit vectors associated with the $r-$, $\Delta q_t-$and Δno-equations in the companion-form of the VECM. Plugging into the current account equation (3) we then obtain the predicted current account - net output ratio:

$$\frac{\widehat{CA_t}}{NO_t} = b\widetilde{r}_t^W + \left[\left(\left(\frac{1}{\gamma} - 1 \right) c + 1 \right) \boldsymbol{e}'_{\boldsymbol{r}} - c \left(\frac{1}{\gamma} - 1 \right) \boldsymbol{e}'_{\boldsymbol{\Delta q}} - \boldsymbol{e}'_{\boldsymbol{\Delta no}} \right] \kappa \boldsymbol{G} \left[\boldsymbol{I} - \kappa \boldsymbol{G} \right]^{-1} Z_t$$

[12] For Japan, we had to use a wholesale price index Mitchell (2003*a*), Table H.1.

[13] The data can be found at:

- www.riksbank.se/templates/Page.aspx?id=27394
- www.norges-bank.no/templates/article____42331.aspx

[14] Since Campbell and Shiller (1987) it is conventional to obtain a companion form based on mixed levels-differences VAR in which the cointegrating relations (here: the current account and the real interest rate) appear in levels and the non-stationary variables in differences. We obtain a similar representation, with the important difference that we write the companion form directly as a function of the VECM-parameters $\boldsymbol{\alpha}$, $\boldsymbol{\beta}$ and $\boldsymbol{\Gamma}(L)$. We discuss this issue in the technical appendix.

[15] In the case of a a VECM with one lagged adjustment term, i.e. $\boldsymbol{\Gamma}(L) = \boldsymbol{\Gamma}_1$, we show that. $\boldsymbol{G} = \begin{bmatrix} \boldsymbol{\Gamma}_1 & \boldsymbol{\alpha} \\ \boldsymbol{\beta}'\boldsymbol{\Gamma}_1 & \boldsymbol{\beta}'\boldsymbol{\alpha} + \boldsymbol{I}_h \end{bmatrix}$. Generalization to higher-order VECM's is straightforward. See the technical appendix for details.

The predicted current account - net output ratio is a function of the parameters c (the consumption / net output ratio), b (the steady-state foreign asset position), and the intertemporal elasticity of substitution ($1/\gamma$). As mentioned above, the real exchange rate is $\Delta q_t = (1-\alpha)\Delta p_t$ and we use data on Δq in our estimation directly, so that we do not have to estimate α. We fix c/no as the sample average from the data. While b could in principle also be obtained from the data, good data on foreign asset positions are very sparse and unreliable for the historical period we are studying here.[16] We therefore estimate b and $1/\gamma$ using a GMM procedure similar to Bergin and Sheffrin (2000) and Kano (1998): to minimize the sum of squared deviations between the actual and the predicted value of CA/NO we perform a grid search over b and $1/\gamma$, letting $1/\gamma$ vary between zero and unity. To initialize the grid search over b, we first obtain an initial measure b_0 of net foreign assets by cumulating the current account and dividing this value by NO_t and averaging over the entire sample period 1885-1939. We then perform the search over the range $b_0 \pm 1$. We also investigate the possibility that the war could have affected steady state asset positions, allowing for a discrete jump Δb in b after 1919.[17] We determine Δb as a third parameter in the grid search procedure that is then performed over the *entire* sample period, 1885-1939.

For each of our eight countries, the panels in Figure 1 plot the predicted and the actual current account (over net output) ratio against each other. The first two columns of Table 1 report correlations between $\widehat{CA/NO}$ and its real counterpart in the data as well as relative standard deviations of the two variables. The last columns report the estimates of $1/\gamma$ and of the steady-state net foreign asset position b (before 1919) and $b+\Delta b$ for the period after 1919.

As is apparent, the model does a remarkable job in replicating the dynamics of historical current accounts. For all countries except Sweden, we obtain correlations around 0.9 and (with the exception of the United Kingdom) the relative standard deviations are all close to unity.

Secondly, our results appear particularly remarkable since they have been obtained over a sample period that covers the Classical Gold Standard as well as World War I and the post-war period inclusive of the Great Depression and its aftermath. It would appear that this was a period of severe parameter instability. However, from the graphs it also seems that the pa-

[16] Lane and Milesi-Ferretti (2007) emphasize that cumulated current accounts are a very imprecise measure of foreign assets in modern data due to valuation effects and measurement error. Meissner and Taylor (2006) study the role of valuation effects in historical data.

[17] Note that our VAR model deliberately does not contain any deterministic controls for structural breaks. Note also that $1/\gamma$ is kept fixed for the entire sample period.

rameters of the cointegrated VAR that govern the dynamics of the model – $\boldsymbol{\alpha}$, $\boldsymbol{\beta}'$ and $\boldsymbol{\Gamma}(\mathbf{L})$ – appear remarkably stable. Otherwise we would expect to see a severely lower performance of the model in some subperiods. No such deterioration is, however, generally apparent. This is our first main empirical point: between the Classical Gold Standard and the Interwar periods, there is considerable stability in the expectation formation mechanism underlying the right hand side of equation (3). The same simple model – without any controls for structural breaks etc. — can explain most of the dynamics of the current account in both periods!

Channels of external adjustment

But did the relative importance of the channels of external adjustment – interest rates (global tilting), exchange rate changes (domestic tilting), output changes (smoothing) and net factor income flows – change over time? We examine what fraction of the variance of the current account can be explained by each of these channels respectively. To this end, we decompose the variance of CA/NO as follows: first, write the current account as the sum of its predicted value and its residual, res, so that $CA/NO = \widehat{CA/NO} + res$. Then plug in from (3), take the variance on both sides and divide by $var(CA/NO)$ to obtain

$$1 = \beta_b + \beta_r + \beta_{\Delta q} + \beta_{\Delta no} + \beta_{res} \tag{6}$$

where

$$\begin{aligned}
\beta_b &= \frac{cov(b\,(\boldsymbol{e}_r' Z_t, CA/NO)}{var(CA/NO)} \\
\beta_r &= \frac{cov\left((1-\phi)\boldsymbol{e}_r'\kappa\boldsymbol{A}\left[\boldsymbol{I}-\kappa\boldsymbol{A}\right]^{-1} Z_t, CA/NO\right)}{var(CA/NO)} \\
\beta_{\Delta q} &= \frac{cov\left(\phi\boldsymbol{e}_{\Delta q}'\kappa\boldsymbol{A}\left[\boldsymbol{I}-\kappa\boldsymbol{A}\right]^{-1} Z_t, CA/NO\right)}{var(CA/NO)} \\
\beta_{\Delta no} &= \frac{cov\left(-\boldsymbol{e}_{\Delta no}'\kappa\boldsymbol{A}\left[\boldsymbol{I}-\kappa\boldsymbol{A}\right]^{-1} Z_t, CA/NO\right)}{var(CA/NO)} \\
\beta_{res} &= \frac{cov(res, CA/NO)}{var(CA/NO)}
\end{aligned}$$

where $\phi = \left(1 - \frac{1}{\gamma}\right) c$. Here, β_b is the contribution of net factor income to the variance of the current account, β_r the contribution of (expected) variation in the world real rate of interest (the global tilting factor), $\beta_{\Delta q}$ the contribution of expected changes in the real exchange rate (the domestic tilting factor), and $\beta_{\Delta no}$ the contribution of output variation (consumption smoothing). The coefficient β_{res} is the fraction of the variance of the current account that remains unexplained by the model.

We refer to these coefficients β_x (where $x = res$, Δno,Δq, r and b in turn) as the pattern of external adjustment. In principle, the coefficients β_x can be estimated country-by-country. However, to obtain a better impression of how these patterns vary across time and across countries, we turn to estimating them from panel regressions

$$x_t^k = \alpha + \tau_t + \mu_k + \beta_x^k(t) \times \left[\frac{CA}{NO}\right]_t^k + \boldsymbol{z}_t^{k\prime}\, \boldsymbol{\delta} + \nu_t^k \tag{7}$$

where x_t^k stands in turn for the VAR-implied expectations of real interest rates, exchange rates etc. on the right hand side of (3), α is a constant and τ_t and $\mu_k$reflect time- and country effects and and the vector $\boldsymbol{z}_t^k$ stacks additional controls. In our baseline specification, we first keep $\beta_x^k(t) = \beta_{x0}$ constant across countries and time. We then let $\beta_x^k(t)$ vary as a function of time and country characteristics. To check whether there are any changes over time in the pattern of external adjustment that are common to all countries., we specify

$$\beta_x^k(t) = \beta_{0x} + \sum_{l=1}^{p} \beta_{lx} PeriodDummy_{lt}$$

where $PeriodDummy_{lt}$ is a sequence of dummies that capture plausible breakpoints in in $\beta_x^k(t)$. We distinguish between WWI, the early Interwar period (1919-28) and the period of the Great Depression (1929-39). In a third specification, finally, we allow $\beta_x^k(t)$ to vary across countries by using two dummies, $OnGS_t^k$ and $OffGS_t^k$, that become one from the point in time at which country k returns to ($OnGS_t^k$) or goes off ($OffGS_t^k$) the Interwar Gold Standard, so that

$$\beta_x^k(t) = \beta_{0x} + \beta_{On,x} OnGS_t^k + \beta_{On,x} OffGS_t^k$$

Plugging these parametrizations for $\beta_x^k(t)$ back into equation (7) and multiplying out, we obtain a sequence of interaction terms between the dummies and the current account. In

the case of the specification that controls for a country's accession to the Gold standard , we also add a first-order terms of the gold-standard dummies to control for first order effects so that so that $z_t^k = \left[\begin{array}{cc} OnGS_t^k, & OffGS_t^k \end{array} \right]'$.

We estimate (7) by panel OLS. The coefficients β_{0x} reflect the pattern of external adjustment during the pre-1913 period, whereas the estimates of $\beta_{l,x}$ measure how this pattern changes in each superiod (relative to the pre-1913 baseline period). The coefficients $\beta_{On,x}$ and $\beta_{Off,x}$ capture the marginal impact of the return to or leaving of the Interwar Gold Standard. Results are presented in Table 3. For each channel, the first cloumn (I) gives the baseline estimate, the second column reports results for the specifcation with common period dummies and the last (III) for the specification with dummies for the Gold Standard.

Across all three specifications it seems that intertemporal smoothing and domestic tilting are the main channels of external adjustment. This is particularly true for the baseline specification (column I) in which all other channels appear completely insignificant.

The specification with period dummies suggests that the end of the classical Gold Standard did not mark an across-the-board change in the patterns of external adjustment: with the exception of tthe WWI-dummy for the net foreign imcome channel, all period dummies are insignificant for all channels. This pattern is consistent with our interpretation of Gold Standard mentality in that it confirms our earlier conclusion that there is a considerable degree of stability between the classical Gold Standard era and the Interwar period. Again, the point estimates suggest that smoothing and domestic tilting explain the bulk of the dynamics (as indicated by β_{x0}), though both would appear only marginally significant in the specification with period dummies.

The specifications in column III, however, reveal that the transition to and from (Interwar) gold did matter for the patterns of external adjustment: a country's transition to the Gold Standard seems to be associated with a sharp decline in the role of intertemporal smoothing and a concomitant increase in the contribution of expected real exchange rate changes (though the latter is only significant at the 10 percent level). This finding illustrates the de-stabilizing effect that the return to gold had on international capital flows: the standard smoothing role of the current account moves to the background in favor of capital flows that appear to be driven by speculative motives and, in particular, by expected realignments of exchange rates. However, once a country leaves the Interwar Gold Standard, the previous pattern of adjustment is restored: the contribution to current account variability of expected

variation in net output increases again while the role of expected real exchange rate changes drops sharply.

Continuity and Change: predicting the spread of the Great Depression

Our results here suggest that current accounts should predict changes in net output and real exchange rates. Given the high degree of stability in our model, current account balances should therefore also have contained a considerable amount of information about medium-term real exchange rate adjustments and net output declines during the Great Depression and after the demise of the Gold Standard.

Figure 3 illustrates this point. The left panel of the figure presents the (negative) smoothing component of the current account ($\sum_{k=1}^{\infty} \kappa^k E_t \Delta \widetilde{no}_{t+k}$) obtained from a re-estimate of our VAR model ending in 1931 against a country's average net output growth rate over the following years, from 1931-1935.[18] There is a very strong positive relation, with countries with highest (smoothing-related) surpluses seing the biggest cumulative (net) output losses. The orders of magnitude line up very well with all points scattered more or less along the 45-degree line.[19] This implies that much of how the global slump would affect certain countries in the medium-run (and how quickly they would recover) was anticipated by the markets at the end of the Gold Standard period and found its reflection in the directions and magnitudes of international capital flows.

The right panel of Figure 3 shows that a similar pattern of out-of sample predictability is apparent for real exchange rates. However, in line with our finding that current accounts after the end of the Gold Standard are determined mainly by expected variation in net output and no longer so much by real exchange rate changes, the link appears somewhat less tight than for net output.

Most researchers find the Great Depression was essentially non-predictable.[20] Ritschl and Wolf (forthcoming) conclude that a breakdown of the Gold Standard was not expected before 1929. They also argue that agents did not anticipate the currency blocks that eventually took shape after 1931. We emphasize that we do not take issue with these findings

[18] We focus on 1931 as the de facto end of the Gold Standard for most countries: five of the eight countries in our sample left the Gold Standard in 1931, including the UK. A sixth, Australia had alredy abandoned gold in 1929. Only the US (1933) and Italy (1936) were to follow later.

[19] A regression reveals a coefficient of close to unity with a t-statistics higher than two and an insignificant constant.

[20] See the literature overview in Ritschl and Wolf (forthcoming): Hamilton (1987, 1992); Dominguez, Fair and Shapiro (1988); Obstfeld and Taylor (2003)

and that we do *not* claim that the Great Depression was predictable. Rather, our exercise here takes the common shock of 1929-1931 as given and asks whether we can use international capital flows during that period to predict how strong the Depression would eventually come to affect a country's output and relative price levels. Hence, our concern here is with whether current accounts contain information about how different countries' reaction to the common shock was. Our results suggests that capital flows contain a considerable amount of such information and that there is a considerable degree of out-of-sample predictive stability in our empircal model, even across the watershed which was the Great Depression.

These findings tie in with our earlier results that document a considerable degree of stability in the factors driving international capital flows in both the classical Gold Standard and the Interwar periods. Given this high degree of continuity, what then can account for the shifts in the moments that we presented in the introductory part of the paper — notably the increase in the volatility of output relative to the current account? We argue: different shocks, not a different transmission mechanism. We turn to exploring this possibility in more detail in the next section, where we present the results from an agnostic identification of the shocks driving current accounts and business cycles.

4.3 Identification of structural shocks: cointegration and heteroskedasticity

Our results so far convey a notion of the stability of the macroeconomic expectation formation mechanism. We interpret this as a reflection of Gold Standard mentality, the fact that the overiding goal of policy makers in the Interwar period was to return to gold.[21] At the same time, the changes in important business cycle moments also document that the nature of shocks to the economcy had changed: the extension of the franchise, the emergence of labor parties, the growing public sector, and the collapse of international co-operation have been prominently discussed in the literature as potential sources of this increase in the volatility of shocks (Eichengreen 1992, Feinstein et al. 1997).

In identifying structural shocks from the model, we exploit both the continuity in expectation formation and the change that is reflected in the changes in key business cycle moments. Specifically, we propose a novel approach that bridges the gap between two — so far

[21] "A further aspect of great significance was the widespread belief in financial and political circles that it was essential to return to the pre-war Gold Standard if the growth and prosperity of the pre-1914 era were to be re-established, whatever the sacrifices their countries would have to make in oder to force down wages and prices so that the pre-war value of the currency could be restored." (Feinstein et al. 1997, p. 1)

quite distinct — literatures: the first is the literature on the identification of permanent and transitory components in cointegrated systems (Johansen, 1995; Hoffmann, 2001*a*). The important insight we take from here is that the cointegrated structure of our empirical model enables us to identify the space spanned by the permanent and transitory shocks without further restrictions from economic theory. Importantly, the stability of the expectation formation mechanism suggests that key parameters governing the stationary and non-stationary dynamics in the system have remained stable across periods. In our four-variable system, there are two cointegrating relationships – reflecting the stationarity of current accounts and real interest rates — which allows us to isolate two permanent shocks (the innovations in the two common trends in real exchange rates and net output) and two transitory shocks, with the two types of shocks orthogonal to each other.

The second part of our approach is based on the literature on identification through heteroskedasticity (Normandin and Phaneuf (2004); Rigobon (2003)). Specifically, the fact that some key business cycle moments *did* change between periods suggests that the structure of underlying shocks may have changed. This, in turn, should show up as heteroskedasticity in the reduced-form residuals. We exploit the heteroskedasticity across regimes to further disentangle the permanent and transitory shocks.

To see, first, how the permanent shocks can be identified from the VECM, let $\boldsymbol{\alpha}_\perp$be the orthogonal complement of $\boldsymbol{\alpha}$. Then premultiply (5) with $\boldsymbol{\alpha'}_\perp$ to obtain

$$\boldsymbol{\alpha'}_\perp \boldsymbol{\Gamma}(\boldsymbol{L})\Delta X_t = \boldsymbol{\alpha'}_\perp \boldsymbol{\epsilon_t}$$

In general, if X_t is of dimension n and if there are h cointegrating relationships, then $\boldsymbol{\alpha'}_\perp$ will be of dimension $(n-h)\times n$ with full rank. Hence, $\boldsymbol{\alpha'}_\perp \boldsymbol{\Gamma}(\boldsymbol{L})\boldsymbol{X_t}$ will be an $(n-h)$-dimensional random walk. Since, according to the Stock-Watson representation there are exactly $n-h$ common trends in X_t, the permanent shocks in the system are given by $\boldsymbol{\alpha'}_\perp \boldsymbol{\epsilon_t}$. By requiring the elements of $\boldsymbol{\alpha'}_\perp \boldsymbol{\epsilon_t}$ to be mutually orthogonal and to have unit variance we obtain the orthogonalized permanent shocks as

$$\boldsymbol{\pi_t} = (\boldsymbol{\alpha'}_\perp \boldsymbol{\Omega} \boldsymbol{\alpha}_\perp)^{-1/2} \boldsymbol{\alpha'}_\perp \boldsymbol{\epsilon}_t$$

where $\boldsymbol{S}_\pi = \boldsymbol{\alpha}'_\perp \boldsymbol{\Omega} \boldsymbol{\alpha}_\perp$ is the variance-covariance matrix of $\boldsymbol{\alpha}'_\perp \boldsymbol{\epsilon}_t$ and $(.)^{1/2}$denotes some matrix root of $\boldsymbol{S}_\pi$. Clearly, any root of $\boldsymbol{\alpha}'_\perp \boldsymbol{\Omega} \boldsymbol{\alpha}_\perp$ will satisfy the orthogonality restriction $var(\boldsymbol{\pi}_t) =$

$\boldsymbol{I}$, reflecting the fact that $\boldsymbol{\alpha}_\perp$ is determined only up to multiplication with a non-singular $(n-h)\times(n-h)$-matrix. Hence additional restrictions will generally be needed to achieve just-identification. In our case here, $n=4$ and $h=2$,so that $\boldsymbol{\pi}_t$ is a two-dimensional vector. Before we turn to identifying these permanent shocks further, we first identify the vector $\tau_{\mathbf{t}}$ of the two remaining transitory shocks by requiring $\tau_{\mathbf{t}}$ to be orthogonal to $\pi_{\mathbf{t}}$. It is easily verified that this leads us to

$$\boldsymbol{\tau}_t = (\boldsymbol{\alpha}'\boldsymbol{\Omega}^{-\mathbf{1}}\boldsymbol{\alpha})^{-1/2}\boldsymbol{\alpha}'\boldsymbol{\epsilon}_{\mathbf{t}}$$

where again the factor $\boldsymbol{S}_\tau^{-1/2} = (\boldsymbol{\alpha'}\boldsymbol{\Omega}^{-\mathbf{1}}\boldsymbol{\alpha})^{-1/2}$ arises due to the requirement that $var(\boldsymbol{\tau_t}) = \boldsymbol{I}$.

We now have two pairs of shocks: one permanent, one transitory. While all four shocks are constructed to be mutually orthogonal, the two types of shocks are not yet uniquely identified among themselves: any matrix root of $\boldsymbol{S}_\pi$ and $\boldsymbol{S}_\tau$ respectively will achieve orthogonalization — the orthogonality conditions that $var(\pi) = var(\tau) = I$ impose only three non-redundant restrictions on $\boldsymbol{S}_\pi$ and $\boldsymbol{S}_{\boldsymbol{\tau}}$ respectively. To single out a particular choice of normalization, we therefore need to impose one further restriction on each of these two matrices.

We obtain these restrictions by recognizing that the reduced-form model parameters $\boldsymbol{\Gamma}(L)$, $\boldsymbol{\beta}$ and $\boldsymbol{\alpha}$ that govern the conditional expectations in (3) seem stable across time while key second moments of the vector X – such as the relative volatility of output and the current account or the correlation between the two variables — seem to have changed. It may therefore be reasonable to assume that the variance of shocks has not been stable across time, while the transmission mechanism as such has been stable. To see how this assumption imposes the required restrictions, let $\boldsymbol{\Omega}^{GS}$ be the reduced-form residual covariance matrix in the Gold Standard period and $\boldsymbol{\Omega}^{IW}$ during the Interwar period.

Let $\boldsymbol{S}_\pi^{GS}$ be the covariance matrix of the permanent shocks in the Gold Standard period. Then the set of orthogonality conditions.

$$var(\boldsymbol{\pi}_t) = \left(\boldsymbol{S_\pi^{GS}}\right)^{\mathbf{-1/2}}\boldsymbol{\alpha}_\perp'\boldsymbol{\Omega}^{\boldsymbol{GS}}\boldsymbol{\alpha}_\perp\left(\boldsymbol{S_\pi^{GS}}\right)^{\mathbf{-1/2\prime}} = \boldsymbol{I}$$

will be satisfied for any matrix root $(.)^{1/2}$ of $\boldsymbol{S}_\pi^{GS}$. Note that

$$\left(\boldsymbol{S}_\pi^{GS}\right)^{-1/2}\boldsymbol{\alpha}_\perp'\boldsymbol{\Omega}^{IW}\boldsymbol{\alpha}_\perp\left(\boldsymbol{S}_\pi^{GS}\right)^{-1/2} = \boldsymbol{\Sigma}$$

will be a positive definite symmetric matrix. Hence there exists an orthonormal basis of Eigenvectors of $\mathbf{\Sigma}$, so that in the spectral decomposition,

$$\mathbf{\Sigma} = \mathbf{Q}\mathbf{\Lambda}_{\pi}^{IW}\boldsymbol{Q}'$$

the matrix $\boldsymbol{Q}$ is orthogonal, i.e. $\boldsymbol{Q'Q} = \boldsymbol{I}$ and $\mathbf{\Lambda}^{IW}$ is diagonal with positive entries. Since $\boldsymbol{Q}$ is orthogonal, the matrix

$$\mathbf{S}_{\pi}^{-1/2} = \mathbf{Q}'\left(\mathbf{S}_{\pi}^{GS}\right)^{-1/2}$$

for any initial choice of $\left(\mathbf{S}_{\pi}^{GS}\right)^{-1/2}$ will satisfy the orthogonality constraint for the Gold Standard period, but it will also satisfy the condition that

$$\boldsymbol{S}_{\pi}^{-1/2}\boldsymbol{\alpha}_{\perp}'\mathbf{\Omega}^{IW}\boldsymbol{\alpha}_{\perp}\boldsymbol{S}_{\pi}^{-1/2} = \mathbf{\Lambda}_{\boldsymbol{\pi}}$$

which means, it achieves orthogonalization of the permanent shocks *also* in the Interwar period. Furthermore, because they are positive, the diagonal entries of $\mathbf{\Lambda}_{\boldsymbol{\pi}}{}^{IW}$ can directly be interpreted as the variances (relative to the Gold Standard period, where they were normalized to unity) of the two permanent shocks in the Interwar period.

The restriction that achieves the identification of $\mathbf{S}_{\pi}^{-1/2}$ here is that $\mathbf{S}_{\pi}^{-1/2}\boldsymbol{\alpha}_{\perp}'\mathbf{\Omega}^{IW}\boldsymbol{\alpha}_{\perp}\mathbf{S}_{\pi}^{-1/2}$ must be diagonal. Hence, the off-diagonal zero in $\mathbf{\Lambda}^{IW}$ is the source of this restriction (*not* the diagonal elements, which are allowed to be freely determined). Clearly, this additional zero restriction must be non-redundant, i.e. it must be different from the zero-restriction which arises from the set of orthogonality restrictions for the first period. This will be the case, whenever $\mathbf{\Omega}^{IW}$ and $\mathbf{\Omega}^{GS}$ are not exact multiples of each other or, equivalently, if the relative increase in variance of the underlying shocks is not uniform across structural shocks, i.e. whenever the diagonal elements of $\mathbf{\Lambda}_{\pi}$ are not equal.

The transitory shocks can now be identified following a completely analogous approach. Here, the respective orthogonality restrictions are to chose the matrix root of $\boldsymbol{S}_{\tau}$ such that

$$\boldsymbol{S}_{\boldsymbol{\tau}}^{-1/2}\boldsymbol{\alpha}'\mathbf{\Omega}^{\boldsymbol{GS}^{-1}}\boldsymbol{\alpha}\boldsymbol{S}_{\boldsymbol{\tau}}^{-1/2'} = \boldsymbol{I}$$

and

$$\boldsymbol{S}_{\boldsymbol{\tau}}^{-1/2}\boldsymbol{\alpha}'\mathbf{\Omega}^{\boldsymbol{IW}^{-1}}\boldsymbol{\alpha}\boldsymbol{S}_{\boldsymbol{\tau}}^{-1/2\prime} = \mathbf{\Lambda}_{\boldsymbol{\tau}}$$

so that $\mathbf{\Lambda}_\tau$ is diagonal with positive diagonal entries.

Once we have identified $\boldsymbol{S}_\tau^{-1/2}$ and $\boldsymbol{S}_\pi^{-1/2}$ in this way, we can now invert the relation between the permanent and transitory shocks $[\boldsymbol{\tau}_t, \boldsymbol{\pi_t}]'$ and ϵ_t so that

$$\epsilon_t = \boldsymbol{P}(\boldsymbol{\Omega}^{\boldsymbol{R}}) \begin{bmatrix} \tau_{\mathbf{t}} \\ \pi_t \end{bmatrix}$$

where $R = GS, IW$ stands for the respective regime and where (as shown in Hoffmann (2001)), the matrix $\boldsymbol{P}$ is given by

$$\boldsymbol{P}(\boldsymbol{\Omega}^{\boldsymbol{R}}) = \begin{bmatrix} \boldsymbol{\alpha S_\tau^{-1/2}}, & \boldsymbol{\Omega^r \alpha_\perp S_\pi^{-1/2}} \end{bmatrix}$$

The variance of the structural shocks in the first period is then just the identity matrix, wheres in the second (Interwar) period, it will be given by

$$var\left(\begin{bmatrix} \boldsymbol{\tau}_{\mathbf{t}} \\ \boldsymbol{\pi}_t \end{bmatrix}\right) = \begin{bmatrix} \mathbf{\Lambda}_\tau & \mathbf{0} \\ \mathbf{0} & \mathbf{\Lambda}_{\boldsymbol{\pi}} \end{bmatrix}$$

This completes our identification procedure. Note that $\mathbf{\Lambda}_\tau$ and $\mathbf{\Lambda}_\pi$ (and therefore the corresponding matrices of eigenvectors, $\boldsymbol{Q}_{\boldsymbol{\pi}}$and $\boldsymbol{Q}_\tau$) are unique only up to the permutation of the diagonal elements. For normalization, we therefore assume that the Eigenvalues on the diagonal appear in decreasing order. We therefore refer to the first shock in each group as the high-volatility shock. Below, we discuss the economic interpretation of these shocks in more detail.

Table 4 presents the shock variances for the Interwar era relative to the pre-war period as we obtain them from this identification procedure. Our procedure reveals huge shifts in relative volatilities: while transitory shocks become less volatile overall , there is a dramatic increase in the relative variance of trend shocks, and in particular for the high-volatility trend shock. Again, this supports our claim that the world economy of the Interwar period shared important features with modern emerging markets: the trend here is clearly the cycle — permanent shocks appear as the dominant source of variability in all eight economies.[22]

The same message is borne out by Table 5 that presents the share of permanent shocks in the forecast error variance, again for the two periods 1885-1913 and 1919-39. While output is

[22] See Aguiar and Gopinath (2007*a*) who document this very pattern for modern-day emerging markets.

dominated by permanent shocks in both periods, the importance of these shocks for no increases further in the second period. Also, while transitory shocks played an important role for the dynamics of real interest rates, the real exchange rate and, notably, the current account, under the Classical Gold Standard, the increase in the variance of trend output shocks implies that all these variables become determined mainly by permanent shocks in the Interwar period.

Interpretation of shocks

It is beyond the scope of this paper to give a tight structural interpretation to the individual shocks. In fact, based on our deliberately agnostic identification scheme there is no reason to believe that the high- and low volatility shocks should actually capture the same structural factors in all countries.

Still, there are interesting patterns that allow us to attempt a broad interpretation of these shocks. Table 6 provides the share of the variance of for the two permanent shocks for the 1919-39 period. In all eight countries, the more volatile shock accounts for virtually all of the variability in the real exchange rate. In seven countries – with the UK being the only exception – it also fully explains the variability in the real interest rate. Conversely, the relatively less volatile shock hardly matters in countries such as Australia, Canada, Japan and Norway. For others however – Italy, Sweden, the UK and the US – it is the main driver of net output variability. For the UK 'low' volatility shock also accounts for for a sizable share of the variation in the current account and for the bulk of real interest rate variability.

Based on these patterns, we suggest to interpret the high-volatility shock as global interest rate disturbance, originating in the instability of the Interwar exchange rate regime and, notably, in the persistent undervaluation of the British pound during the Interwar Gold Standard. The second permanent, 'low volatility' shock displays much more idiosyncratic behavior and we refer to it as a country-specific disturbance.

We underpin this interpretation of one shock as global and the other as country-specific by turning to a key proposition from economic theory: global shocks should affect the world real interest rate, whereas idiosyncratic shocks should not be directly related to interest rate variability. To test this proposition, for each of the four types of shocks we stack the shocks from all eight countries in our sample. We then extract the first three principal components from each of this eight-dimensional vector time-series of shocks. Finally, for all four types of

shocks, we regress the world real interest rate on these cross-country principal components. Our results are given in Table 7. The regressions of the interest rate on the principal components of the high-volatility, transitory and the low-volatility permanent shocks respectively do not yield significant coefficients and have low R². Conversely, the regression of r^w on the first three principal components of the low-volatility transitory and the high-volatility permanent shock yield highly significant coefficients and quite high R^2 . This pattern, first, underpins our interpretation of the permanent high-volatility shock as a global disturbance and of the low-volatility permanent shock as a largely idiosyncratic disturbance. Second, the findings also suggest that the the same distinction between global and country-specific can also be applied to the transitory shocks, with the (transitory) low-volatility shock reflecting global variation in interest rates and the (transitory) high-volatility shock reflecting idiosyncratic variation.

Third, tying our findings here with the previous results in Table 4 and 5 shows that both global and country-specific shocks have become a lot more permanent in the Interwar period – but the former more so than the latter: both the global and the country-specific transitory shock drop in volatility in the Interwar period, but the decline for the global transitory shock is stronger – it is largely associated with the low-volatility transitory disturbance. Conversely, among the permanents shocks, it is mainly the high-volatility disturbance that is associated with global interest rate variation, whereas the lower-volatility shock is more idiosyncratic.

Changes in the moments of current account and net output

This shift towards more permanent and more global shocks can also explain the decline in the variability of the current account relative to net output growth that we documented initially. To illustrate this point, we calculate the historical decomposition of net output growth and the current account and ask what the relative variability of these variables would have been in the Interwar period if only permanent shocks had occurred (i.e. we re-assemble the two variables, switching off the transitory shock in the historical decomposition). Figure 4 provides the results of this exercise. The horizontal axis shows the relative variability in the Gold standard period, the vertical in the Interwar period. Blue dots are reproduced from Figure 2, showing the decline of the relative variability in the data. Red circles illustrate what the relative variability would have been if only permanent shocks had occur ed in the Interwar

period (taking pre-1913 relative variability as given). It is clearly apparent that permanent shocks explain the decline in the relative variability of CA/NO and Δno. In fact, for all countries except the UK, the predicted decline is actually slightly bigger (red circle are below the corresponding blue dot). This is what we would expect since, clearly, there is some transitory variability in the current account that is unrelated to output variability and which would tend to increase the variability of the current account ceteris paribus, but the picture is clear: the shift towards more permanent and global shocks accounts for the bulk of the decline in the the variability of international capital flows relative to national cash flows.

5 Conclusions

This paper has applied a simple intertemporal, present-value model of the current account to study capital flows and international business cycles in the period between 1885 and 1939. To our knowledge, we are the first to rigorously apply such a model to historical data. The period of the classical Gold Standard with its high levels of international capital mobility and uni-directional capital flows (Obstfeld (2004)) would appear as an ideal testing ground for such a model. Our main result, however, is not that this model fits data from the Gold-Standard period well. More importantly, the very same model — that does not have any hard-wired frictions or limitations on international capital flows — with the same set of parameters explains the data for *both* the Classical Gold Standard and the Interwar period. We interpret this finding as a sign of Gold Standard Mentality: the drivers behind international capital flows seems to have stayed remarkably stable over the entire period, reflecting the stability of agent's expectation with respect to how monetary policy actions might affect future variation in national cash flows (output), exchange rates and interest rates.

At the same time, we document that key business cycle moments changed between the Interwar period and the pre-war Gold Standard. Specifically, we have focused on the stylized fact that the volatility of the current account relative to output generally decreased. What can explain this simultaneous pattern of continuity and change? We argue that a) the predictive stability of our model is an indication of a fundamental stability in the macroeconomic transmission mechanism but that b) the shifting correlations and volatilities highlight the importance of changes in the structure of underlying shocks.

Our explanation follows a recent literature in modern-day emerging market macroeco-

nomics in arguing that basic models without frictions match the data from emerging market economies quite well once the underlying shocks are allowed to be more persistent than usually specified for industrialized economies. This does *not* mean that financial or goods market frictions are unimportant in these economies. Rather, these frictions manifest themselves in the structure of the underlying shocks — their volatility and persistence — rather than in a breakdown of the fundamental model of the transmission mechanism. Following this logic, we argue that the change in the moments that we document here can be explained by more permanent and volatile shocks to trend output and exchange rates. Our results also suggest that these shocks seem to have become much more global. These structural shifts seem to have interacted to lower the variability of the current account relative to output. In particular, the Interwar period saw the emergence of a global business cycle that had its roots in the global instability of the era. The fact that macroeconomic fluctuations appear permanent in nature therefore is likely to reflect continual structural change. This pattern of 'the cycle being the trend' has been prominently documented for modern-day emerging economies (see Aguiar and Gopinath (2007*b*)), which suggests that, in important ways, the world economy emerging from the war shared important features with today's emerging markets.

Our findings complement more narrative evidence on the relative roles of continuity and change in explaining the experience of the Interwar period: first, our finding that shocks have changed — more volatile, more persistent and more global — is consistent the view that that World War I was the watershed for international capital mobility (Obstfeld and Taylor (2004)) and that the Interwar period saw goods and financial markets that were much more segmented. Secondly, our result that the transmission mechanism has stayed remarkably constant lines up with the view (Eichengreen, 1992) that there was remarkable continuity in policymakers' mindset and in their policy and institutional responses (such as the ill-fated return to the gold exchange standard) to what were effectively very different shocks.

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Technical Appendix

We estimate a vector error correction model of the form

$$\mathbf{\Gamma}(L)\Delta X_t = \boldsymbol{\alpha}\boldsymbol{\beta}' X_{t-1} + \epsilon_{\mathbf{t}}$$

To obtain the conditional expectations on the right hand side of (3), we then stack the VECM in a companion form so that

$$Z_{t+1} = \boldsymbol{G} Z_t + \boldsymbol{u}_{t+1}$$

with $Z_{t+1} = \left[\begin{array}{ccc} \Delta X_t' & ... & \Delta X_{t-p+1}' \quad [\beta' X_t]' \end{array} \right]'$. Defining Z_t in this particular way, allows us to use the Hansen-Sargent prediction formula to obtain the discounted sum of expected changes in net output and real exchange rates as well as levels of r on the right hand side of (3). However, while it is usually straightforward to write a VECM as VAR in levels, our particular way of defining Z_t as containing a mix of levels and differences leaves the companion matrix $\boldsymbol{G}$ to be determined. Here, we discuss how we construct this matrix in the case of one lagged adjustment in the VECM. Generalization to higher order VECM's is straightforward. First write

$$\left[\begin{array}{cc} \boldsymbol{I}_n & \mathbf{0}_{n\times h} \\ -\boldsymbol{\beta}' & \boldsymbol{I}_h \end{array} \right] \left[\begin{array}{c} \Delta X_t \\ \beta' X_t \end{array} \right] = \left[\begin{array}{cc} \mathbf{\Gamma}_1 & \boldsymbol{\alpha} \\ \mathbf{0}_{h\times n} & \boldsymbol{I}_h \end{array} \right] \left[\begin{array}{c} \Delta X_{t-1} \\ \boldsymbol{\beta}' X_{t-1} \end{array} \right] + \left[\begin{array}{c} \boldsymbol{I}_n \\ \mathbf{0}_{h\times n} \end{array} \right] \epsilon_t$$

where n is the dimension of the VECM and h the number of cointegrating relationships. Then, recognizing that

$$\left[\begin{array}{cc} \boldsymbol{I}_n & \mathbf{0}_{n\times h} \\ \boldsymbol{\beta}' & \boldsymbol{I}_h \end{array} \right] \left[\begin{array}{cc} \boldsymbol{I}_n & \mathbf{0}_{n\times h} \\ -\boldsymbol{\beta}' & \boldsymbol{I}_h \end{array} \right] = \boldsymbol{I}_n$$

we obtain

$$\begin{aligned} \left[\begin{array}{c} \Delta X_t \\ \boldsymbol{\beta}' X_t \end{array} \right] &= \left[\begin{array}{cc} \boldsymbol{I}_n & \mathbf{0} \\ \boldsymbol{\beta}' & \boldsymbol{I}_h \end{array} \right] \left[\begin{array}{cc} \mathbf{\Gamma}_1 & \boldsymbol{\alpha} \\ \mathbf{0}_{h\times n} & \boldsymbol{I}_h \end{array} \right] \left[\begin{array}{c} \Delta X_{t-1} \\ \boldsymbol{\beta}' X_{t-1} \end{array} \right] + \left[\begin{array}{cc} \boldsymbol{I}_n & \mathbf{0}_{n\times h} \\ \boldsymbol{\beta}' & \boldsymbol{I}_h \end{array} \right] \left[\begin{array}{c} \boldsymbol{I}_n \\ \mathbf{0} \end{array} \right]_{h\times n} \epsilon_t \\ &= \left[\begin{array}{cc} \mathbf{\Gamma}_1 & \boldsymbol{\alpha} \\ \boldsymbol{\beta}'\Gamma_1 & \boldsymbol{\beta}'\boldsymbol{\alpha} + \boldsymbol{I}_h \end{array} \right] \left[\begin{array}{c} \Delta X_{t-1} \\ \boldsymbol{\beta}' X_{t-1} \end{array} \right] + \left[\begin{array}{c} \boldsymbol{I}_n \\ \boldsymbol{\beta}' \end{array} \right] \epsilon_t \end{aligned}$$

so that

$$\boldsymbol{G} = \left[\begin{array}{cc} \mathbf{\Gamma}_1 & \boldsymbol{\alpha} \\ \boldsymbol{\beta}'\mathbf{\Gamma}_1 & \boldsymbol{\beta}'\boldsymbol{\alpha} + \boldsymbol{I}_h \end{array} \right]$$

is the desired companion matrix.

Table 1: Volatility of Business Cycle Components

	GDP	NO	CA/NO	r_w
1885-1913				
AUS	0.05	0.03	0.04	1.64
CAN	0.04	0.03	0.02	1.66
JAP	0.04	0.05	0.03	1.72
NOR	0.02	0.01	0.02	1.89
SWE	0.02	0.01	0.01	1.74
UK	0.02	0.02	0.01	1.05
USA	0.04	0.03	0.01	1.76
1919-1939				
AUS	0.05	0.04	0.08	1.29
CAN	0.09	0.06	0.03	1.4
JAP	0.04	0.03	0.02	1.54
NOR	0.04	0.05	0.03	1.26
SWE	0.04	0.03	0.02	1.16
UK	0.03	0.04	0.02	1.56
USA	0.09	0.06	0.01	1.61

Standard deviations of HP-filtered data for output (GDP), net output (NO) and the world real interest rate (r^w). (GDP, NO in logs, smoothing weight: 100)

Table 2: Statistics for the Predicted and Actual Current Account (1885-1939)

	Correlation $\rho(\widehat{CA/NO}, CA/NO)$	Rel. Std. Dev. $\frac{\sigma(\widehat{CA/NO})}{\sigma(CA/NO)}$	Subst. Elasticity $(1/\gamma)$	Net Foreign Assets (b) before 1919	Net Foreign Assets (b) after 1919
AUSTRALIA	0.88	0.93	0.01	-0.37	-0.27
CANADA	1.00	0.91	1.01	-0.85	-0.75
ITALY	0.98	0.99	0.91	-0.92	-1.02
JAPAN	0.95	1.24	0.31	0.09	-0.01
NORWAY	0.87	1.20	0.51	0.18	0.68
SWEDEN	0.70	1.26	0.21	0.50	0.50
UK	1.00	2.33	0.01	0.16	0.16
USA	0.99	1.01	0.51	-0.07	0.03

Table 3: Beta Decomposition of CA/NO

Regressors		Interest income			Global Tilting			Domestic Tilting			Smoothing			Unexplained		
		Channel of External Adjustment														
		I	II	III	I	II	III	I	II	III	I	II	III	I	II	III
$\left[CA_t^k/NO_t^k\right]$	(β_{x0})	-0.09	0.02	-0.10	-0.15	**-0.18**	*-0.16*	**0.37**	0.44	**0.36**	**0.88**	*0.72*	**0.92**	-0.02	-0.00	-0.03
		(-1.06)	(0.81)	(-1.22)	(-1.80)	(-3.67)	(-1.87)	(2.57)	(1.52)	(2.45)	(3.89)	(1.95)	(4.54)	(-0.30)	(-0.01)	(-0.45)
Interactions of $\left[CA_t^k/NO_t^k\right]$...																
.. with period dumies																
WWI	(β_{x1})		**-0.28**			0.07			-0.07			0.37			-0.09	
			(-2.55)			(0.49)			(-0.27)			(1.11)			(-0.55)	
$1919-1928$	(β_{x2})		-0.05			0.00			-0.09			0.09			0.05	
			(-0.47)			(0.02)			(-0.25)			(0.33)			(0.42)	
$1929-39$	(β_{x3})		0.01			0.08			-0.19			0.17			-0.06	
			(0.06)			(0.77)			(-1.18)			(1.31)			(-0.57)	
... Gold Standard																
$OnGS_t^k$	$(\beta_{x,on})$			0.03			-0.08			*0.28*			**-0.53**			0.14
				(0.16)			(-0.42)			(1.80)			(-3.95)			(1.59)
$OffGS_t^k$	$(\beta_{x,off})$			0.17			-0.00			**-0.42**			**0.53**			-0.18
				(0.96)			(-0.55)			(-2.80)			(2.41)			(-1.22)
$\overline{R}^2$		-0.13	-0.08	-0.13	-0.06	-0.04	-0.04	0.17	0.19	0.19	0.58	0.61	0.60	-0.16	-0.01	-0.16

The Table reports the results of panel regressions of the form

$$x_t^k = \boldsymbol{d_t^k} + \beta_t^k \times \left[\frac{CA}{NO}\right]_t^k + \boldsymbol{z}_t^{k\prime}\,\boldsymbol{\delta} + v_t^k$$

for all channels of external adjustment. Here, x_t^k stands, in turn, for the VAR-implied expectations of the interest income, global tilting, domestic tilting and consumption smoothing channels and for the unexplained component $CA_t - \widehat{CA}_t^k$. The vector $\boldsymbol{d_t^k} = [\ \tau_t \quad \mu_k \quad \alpha_0\ ]'$ stacks the constant as well as country-fixed and time effects and $\boldsymbol{z}_t^k$ contains additional controls. In columns labelled I, we report the baseline specification, i.e. $\beta_t^k = \beta_0$ is constant over time and countries. In columns labelled II, interactions with the period dummies, are added to the baseline regression, so that $\beta_t^k = \beta_0 + \left[\sum_{l=1}^p \beta_{lx} PeriodDummy_{lt}\right]$. Columns labelled III present regressions that control for a countries return to and exit from the interwar gold standard, i.e. $\beta_t^k = \beta_0 + \beta_{on,x} OnGS_t^k + \beta_{off,x} OffGS_t^k$. In specification III, we also control for first-order effects by choosing $\boldsymbol{z}_t^k = [\ OnGS_t^k, \quad OffGS_t^k\ ]'$. Estimates of $\boldsymbol{\delta}$ not reported. t-statistics, based on (heteroskedasticity- and autocorrelation-consistent) standard errors clustered by country appear in parentheses.

Table 4: Volatility of structural shocks 1919-39 vs. 1885-1913

1919-39 relative to Gold Standard period

	transitory		permanent	
	high vol.	low vol.	high vol.	low vol.
AUSTRALIA	0.81	0.07	2.43	0.66
CANADA	1.20	0.15	3.98	0.57
JAPAN	0.28	0.19	5.28	3.65
ITALY	3.41	0.45	2.01	0.83
NORWAY	0.47	0.18	13.77	1.40
SWEDEN	0.70	0.30	15.08	4.11
UK	0.26	0.16	4.04	2.23
USA	0.60	0.17	3.54	1.97

Table 5: **Variance contribution of Permanent shocks**

Horizon/yrs	1885-1913 r	1885-1913 CA/NO	1885-1913 q	1885-1913 no	1919-1939 r	1919-1939 CA/NO	1919-1939 q	1919-1939 no
AUSTRALIA								
1	0.54	0.39	0.44	0.81	0.98	0.41	0.95	0.92
2	0.43	0.36	0.33	0.85	0.96	0.34	0.95	0.97
5	0.42	0.32	0.61	0.92	0.95	0.40	0.98	0.99
8	0.42	0.32	0.76	0.95	0.95	0.40	0.99	0.99
CANADA								
1	0.02	0.22	0.66	0.86	0.99	0.14	0.99	0.98
2	0.02	0.22	0.68	0.87	0.99	0.25	0.99	0.98
5	0.02	0.21	0.75	0.90	0.99	0.41	1.00	0.98
8	0.02	0.21	0.80	0.91	0.99	0.45	1.00	0.98
ITALY								
1	0.19	0.14	0.92	0.70	0.99	0.94	1.00	1.00
2	0.34	0.35	0.89	0.54	1.00	0.97	1.00	1.00
5	0.34	0.39	0.92	0.57	1.00	0.98	1.00	1.00
8	0.34	0.39	0.95	0.73	1.00	0.98	1.00	1.00
JAPAN								
1	0.29	0.13	0.75	0.66	0.98	0.00	0.75	0.15
2	0.21	0.12	0.78	0.66	0.95	0.01	0.82	0.33
5	0.17	0.12	0.90	0.79	0.91	0.08	0.95	0.74
8	0.17	0.12	0.94	0.85	0.91	0.09	0.97	0.82
NORWAY								
1	0.19	0.15	0.92	0.32	0.97	1.00	1.00	1.00
2	0.27	0.21	0.96	0.53	0.96	1.00	1.00	1.00
5	0.33	0.21	0.98	0.70	1.00	1.00	1.00	1.00
8	0.32	0.23	0.98	0.78	1.00	1.00	1.00	1.00
SWEDEN								
1	0.14	0.05	0.97	0.55	0.95	0.99	1.00	1.00
2	0.17	0.09	0.79	0.65	1.00	0.98	1.00	1.00
5	0.15	0.11	0.76	0.79	1.00	0.99	1.00	1.00
8	0.15	0.11	0.84	0.86	1.00	0.98	1.00	1.00
UK								
1	0.01	0.38	0.99	0.10	0.98	0.93	1.00	0.92
2	0.02	0.38	1.00	0.18	0.98	0.93	1.00	0.95
5	0.03	0.38	1.00	0.47	0.98	0.94	1.00	0.98
8	0.03	0.38	1.00	0.66	0.98	0.94	1.00	0.99
US								
1	0.18	0.19	0.79	0.99	1.00	0.78	1.00	1.00
2	0.18	0.21	0.83	0.99	1.00	0.89	1.00	1.00
5	0.19	0.24	0.90	1.00	1.00	0.94	1.00	1.00
8	0.19	0.24	0.94	1.00	1.00	0.95	1.00	1.00

Table 6: **Variance contribution of the permanent shocks – breakdown by volatility**

	1919-39							
	high-volatility shock				low-volatility shock			
Horizon/yrs	r	CA/NO	q	no	r	CA/NO	q	no
AUSTRALIA								
1	0.97	0.02	0.94	0.71	0.01	0.40	0.01	0.21
2	0.93	0.08	0.94	0.90	0.03	0.26	0.01	0.07
5	0.92	0.23	0.97	0.93	0.03	0.17	0.00	0.06
8	0.92	0.23	0.98	0.94	0.03	0.17	0.00	0.05
CANADA								
1	0.98	0.13	0.99	0.96	0.01	0.01	0.00	0.01
2	0.98	0.25	0.99	0.96	0.01	0.01	0.00	0.02
5	0.98	0.40	0.99	0.96	0.01	0.00	0.00	0.02
8	0.98	0.45	1.00	0.96	0.01	0.00	0.00	0.02
ITALY								
1	0.98	0.21	0.99	0.13	0.01	0.73	0.01	0.87
2	0.89	0.39	0.99	0.12	0.11	0.58	0.01	0.88
5	0.89	0.70	0.99	0.10	0.11	0.28	0.01	0.90
8	0.89	0.71	1.00	0.13	0.11	0.27	0.00	0.87
JAPAN								
1	0.97	0.00	0.52	0.03	0.00	0.00	0.23	0.13
2	0.95	0.00	0.61	0.22	0.00	0.00	0.20	0.11
5	0.90	0.08	0.86	0.62	0.00	0.00	0.09	0.12
8	0.90	0.08	0.90	0.69	0.00	0.00	0.07	0.13
NORWAY								
1	0.92	1.00	0.99	1.00	0.04	0.00	0.01	0.00
2	0.90	1.00	1.00	1.00	0.07	0.00	0.00	0.00
5	1.00	1.00	1.00	1.00	0.00	0.00	0.00	0.00
8	1.00	1.00	1.00	1.00	0.00	0.00	0.00	0.00
SWEDEN								
1	0.16	0.97	1.00	0.44	0.80	0.02	0.00	0.55
2	0.96	0.94	1.00	0.34	0.04	0.05	0.00	0.66
5	0.96	0.93	1.00	0.26	0.04	0.05	0.00	0.74
8	0.96	0.93	1.00	0.24	0.04	0.05	0.00	0.76
UK								
1	0.01	0.00	1.00	0.25	0.97	0.93	0.00	0.67
2	0.01	0.00	1.00	0.19	0.97	0.93	0.00	0.76
5	0.01	0.00	1.00	0.11	0.96	0.93	0.00	0.87
8	0.01	0.00	1.00	0.09	0.96	0.93	0.00	0.90
USA								
1	1.00	0.59	0.93	0.07	0.00	0.19	0.06	0.93
2	1.00	0.80	0.94	0.05	0.00	0.10	0.05	0.95
5	1.00	0.89	0.96	0.02	0.00	0.05	0.04	0.98
8	1.00	0.90	0.97	0.02	0.00	0.04	0.03	0.98

Table 7: World interest rates and structural shocks

Panel A: 1885-1913

	transitory		permanent	
	high vol.	low vol.	high vol.	low vol.
1. PC	-0.0109	0.009	0.004	-0.003
	(-5.19)	(9.01)	(1.68)	(-1.06)
2. PC	0.00	-0.002	0.009	-0.0002
	(0.12)	(-1.12)	(3.21)	(-0.06)
3. PC	0.004	0.003	-0.005	0.003
	(1.72)	(1.08)	(-1.53)	(0.75)
const.	0.02	0.03	0.02	0.022
	(5.73)	(11.86)	(6.79)	(5.58)
R^2	0.54	0.77	0.38	0.06

Panel B: 1919-39

	transitory		permanent	
	high vol.	low vol.	high vol.	low vol.
1. PC	-0.01	0.01	-0.11	0.01
	(-1.98)	(6.36)	(-5.13)	0.30
2. PC	0.00	0.00	0.0048	0.001
	(0.051)	(2.20)	(0.10)	(0.11)
3. PC	-0.00	-0.004	-0.02	-0.02
	(-0.25)	(-2.02)	(-0.37)	(-0.83)
const.	0.04	0.02	0.04	0.04
	(2.52)	(2.29)	(3.36)	(2.17)
R^2	0.22	0.71	0.65	0.04

The table shows regressions of the world real interest rate, r_t^w , on the first three (cross-country) principal components of each type of shock: $r_t^w = \alpha + \sum_{i=1}^{3} \gamma_i PC_t^i(k) + \epsilon_t^i$ where $PC_t^i(k)$ is the $i-th$ principal component extracted from the eight-dimensional (i.e. stacked across countries) series of shocks of type $k = \{high\ volatility,\ low\ volatilty\} \times \{permanent, transitory\}$.

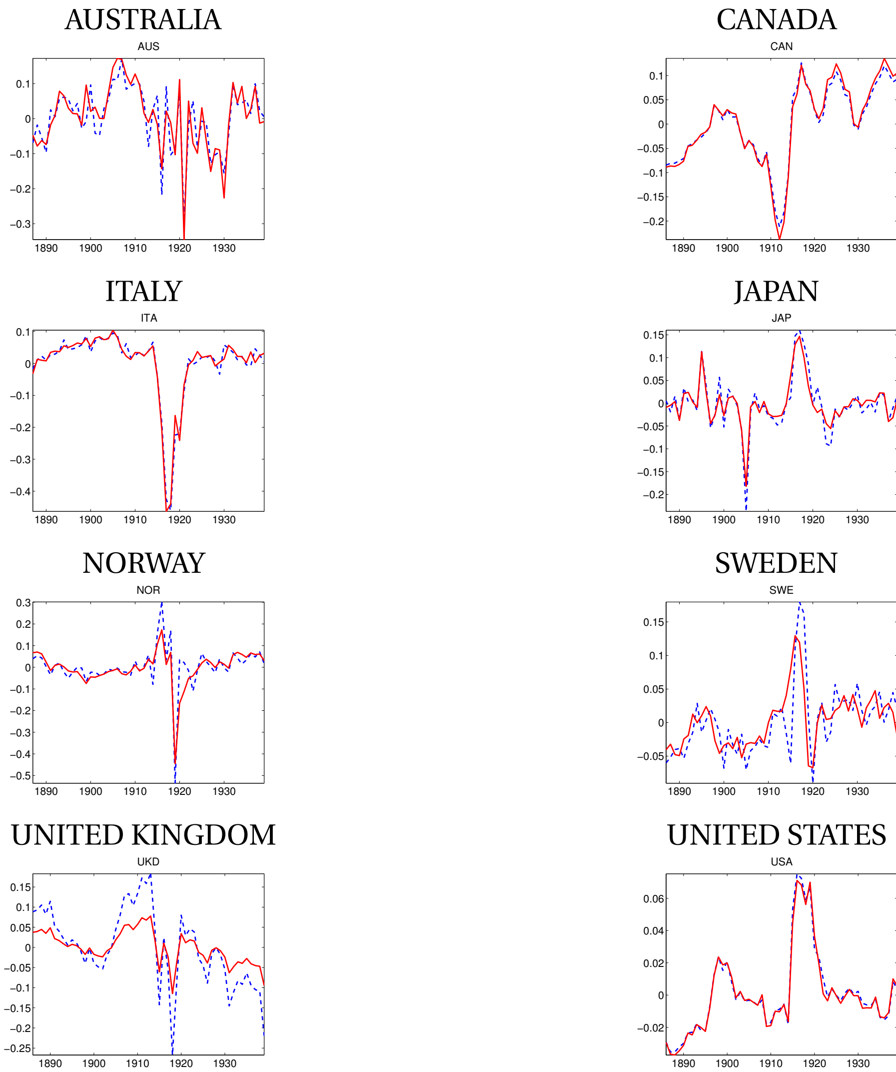

Figure 1: Actual current account/ net output ratio (solid, red line) vs. predicted current account (dashed line).

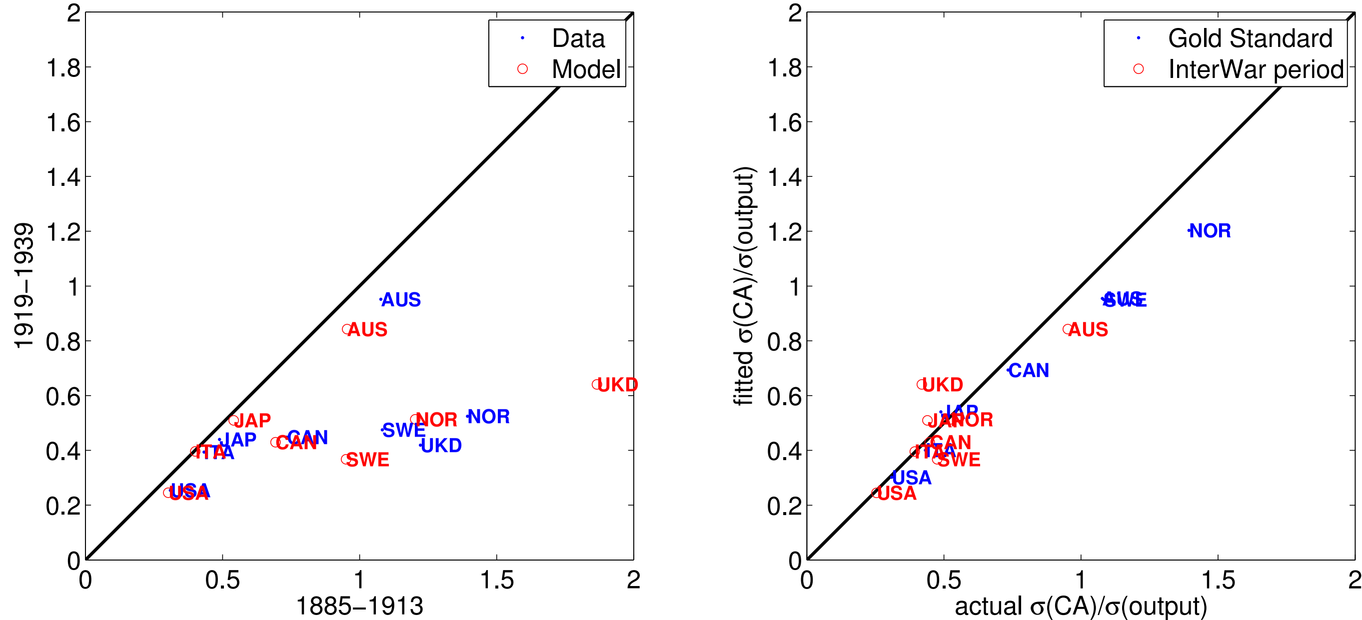

Figure 2: Relative standard deviations of the Current account / net output ratio and net output growth ($\sqrt{(var(CA/NO)/var(\Delta no)}$).
The left panel offers a comparison across the two periods 1885-1913 and 1919-39, with a point below the 45-degree line indicating a lower relative standard deviation in the interwar period. Blue dots indicate the data, red circles are the fitted moments predicted from the model.
The right panel plots the actual relative standard deviation against the one predicted from the model, where blue dots indicate the Gold Standard period (1885-1913) and red circles the interwar period (1919-39). The black line is the 45-degree line.

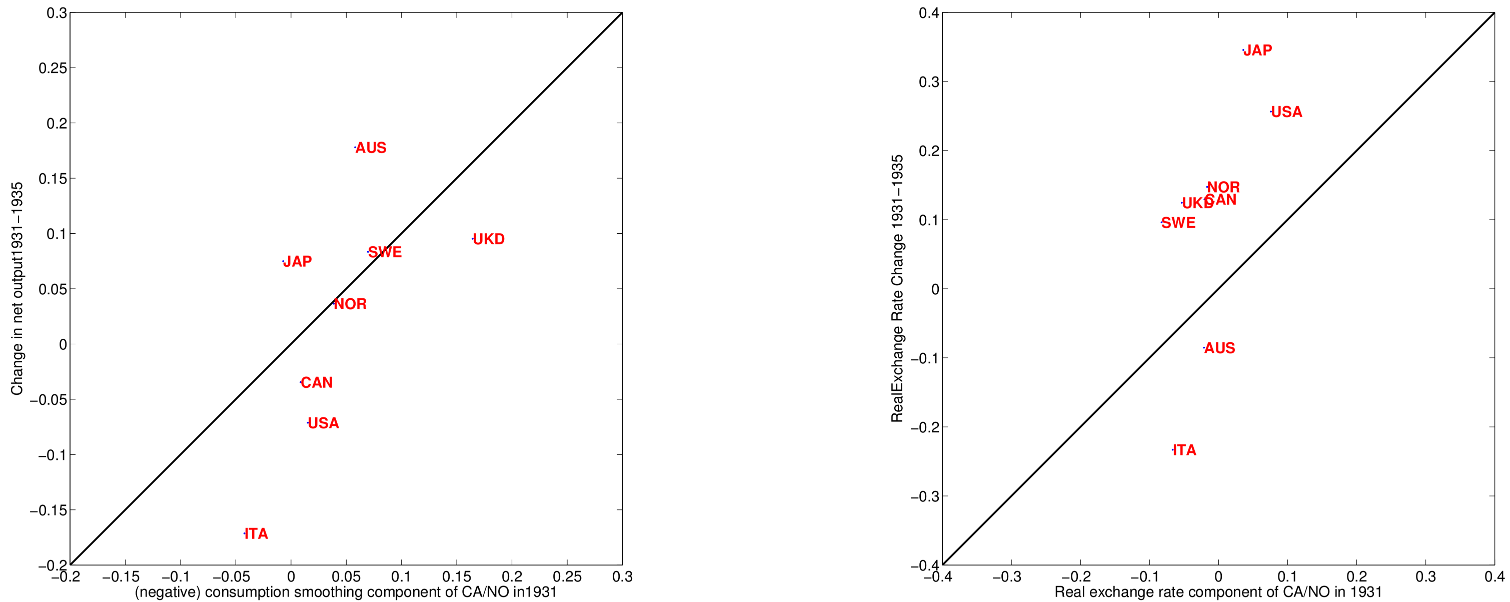

Figure 3: The left panel plots the annualized rate of change of net output during 1931-37 $(NO_{1935}) - log(NO_{1931}))$ against the (negative) smoothing component of the current account $(\sum_{k=1}^{\infty} \kappa^k E_t \Delta \widetilde{no}_{t+k})$ based on a VAR estimated from the sample period 1885-1931. The right panel plots the (annualized) depreciation of the trade-weighted real exchange rate $(P_{1935}) - log(P_{1931}))$ over the same period against the domestic tilting (exchange rate) component of the current account $(\sum_{k=1}^{\infty} \kappa^k E_t \widetilde{\Delta q}_{t+k})$ implied by the same VAR.

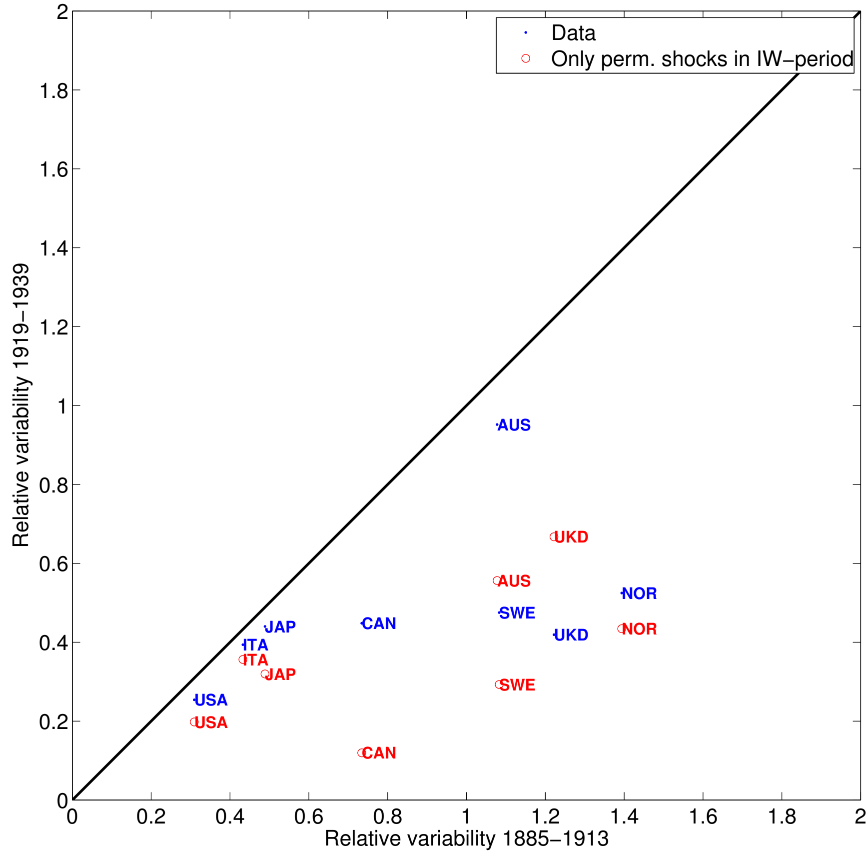

Figure 4: The decline in relative volatility of CA/NO and Δno revisited. (Blue) dots show the decline of $\sigma_{CA/NO}/\sigma_{\Delta no}$ in the data between the Gold Standard period (horzontal axis) and the Interwar Period (vertical axis). (Red) circles show what the development of $\sigma_{CA/NO}/\sigma_{\Delta no}$ would have been, if only permanent shocks had occured during the interwar period (taking the volatilities in the Gold Standard period as given).