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Effective Exchange Rates, Current Accounts and Global Imbalances

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Joscha Beckmann and Robert Czudaj[1]

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Abstract

This study analyzes the dynamics between real effective exchange rates and current accounts from a novel perspective. We start by dissecting long-run and time-varying short-run dynamics as well as causalities between both variables. Following this, we extend our framework by including short-term interest rates. Finally, we examine common exchange rate and current account dynamics across countries based on common factors. Our results show that a real appreciation coincides with a worsening of the current account in most cases. The adjustment pattern is time-varying but suggests that the causality mainly runs from effective exchange rates to current accounts. However, an extension of our framework based on monthly data shows that trade balance adjustment is observed less frequently, suggesting that valuation effects play an important role for the relationship between current accounts and exchange rates. From a global point of view, cross-country trends which drive exchange rates and current accounts also share similar dynamics over the long-run, which is an important finding in the context of global imbalances.

JEL Classification: F31, F32

Keywords: Current account; global imbalances; Markov-switching; multivariate cointegration; real exchange rates

March 2016

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1 Introduction

Current account imbalances are a key feature of the current international monetary system and have triggered controversial discussions among policymakers and economists in recent times. This is true not only for the origins and consequences but also for any potential mechanisms to reverse sustainable current account deficit or surpluses. Even the importance of the exchange rate as possibly the most intuitive adjustment tool is still subject to controversy.[1] For instance, the idea that a flexible exchange rate regime generally facilitates current account adjustment has yet to be convincingly demonstrated (Chinn and Wei, 2013). A reacquisition aimed at generating an impact of exchange rate changes on the current account via trade effects is an exporters' producer-currency pricing strategy (Campa and Goldberg, 2005). However, such a pass-through is frequently found to be incomplete due to local-currency pricing and heterogeneity across sector firms and exporters (Berman *et al.*, 2012; Brun-Aguerre *et al.*, 2012; Auer and Schoenle, 2016). Against this background, the overall empirical link between exchange rates and current accounts is far from unambiguous.

This paper offers a comprehensive and novel analysis of the relationship between effective exchange rates and current accounts. We undertake several steps to clarify the issue of causality, both for a broad number of single countries and from a global perspective. More precisely, we make three key contributions: **firstly**, we dissect long-run and short-run relationships between real effective exchange rates and current accounts for eleven major economies. In doing so, we make use of monthly and quarterly data for both, the current account relative to GDP and the real trade balance. This procedure provides an implicit robustness check for our results and a direct evaluation of the impact of exchange rate changes via the trade channel.[2]

Our **second** contribution stems from the consideration of time-varying short-run dynamics which are related to disequilibria from the underlying long-run relationships. Such a framework has turned out to be useful for examining exchange rate dynamics since it allows for the existence of a stable relationship between current accounts and exchange rates while causality is allowed to change over time (Sarno *et al.*, 2004; Sarno and Valente, 2006). The degree and character of the current account or trade balance adjustment allows an evaluation of the exchange rate channel for the correction of global imbalances. Leaving the country dimension, our **third** contribution stems from focusing on

[1] As an example, U.S. dollar adjustment has been discussed as a solution to the twin deficit in the federal budget and the current account that has been observed recently and also during the eighties (Krugman, 1985; Obstfeld and Rogoff, 2009).

[2] Owing to the large number of estimated models, we focus solely on real effective exchange rates in the following, and thus do not discriminate between nominal exchange rates and price dynamics. We leave this issue for future research.

the question of whether effective exchange rates and current accounts or trade balances share similar dynamics across countries. This issue has not been explicitly considered in previous studies although it is important when it comes to (global) policy recommendations regarding global imbalances and the role of the exchange rate. Similar to the country level, we also examine long-run and time-varying short-run dynamics for our global factor model. As will be shown our results indicate that a real effective depreciation coincides with a worsening of the current account relative to GDP for most economies under observation. We also find that the effects are significantly weaker for the trade balance in most case, suggesting that the trade channel is not able to explain the obtained result completely.

When discussing the links between exchange rates and current accounts, a distinction between positive understanding and normative evaluation is necessary (IMF, 2013). In this paper, we are neither interested in calculating a "fair" exchange rate which is designed to "correct" global imbalances nor are we aiming at providing a general answer to the question of whether effective exchange rates should adjust or not. Conducting a positive rather than a normative approach, our key questions are: if, to what extent and during which times current accounts (relative to the GDP) and/or trade balances have been linked to effective exchange rates in the past. Our paper is therefore in line with several studies focusing on the relationship between exchange rates and current accounts from different perspectives (Lee and Chinn, 2006; Chinn and Lee, 2009; Chinn and Wei, 2013; Shibamoto and Kitano, 2012). We suggest an error correction model with Markovian shifts to model the relationship between current accounts and exchange rates as a key difference to the existing empirical literature which is notably silent regarding nonlinearities in this relationship although previous studies have provided broad evidence in favor of nonlinear exchange rate dynamics.

The remainder of this paper is organized as follows. The following section discusses the motivation for our analysis from several perspectives: after a brief recap of the theoretical background, we categorize and summarize previous empirical findings and justify our own empirical approach. Section 3 provides our dataset, a detailed description of our empirical approach and a discussion of our findings. In the empirical part of the paper we start with an analysis of the relationship between current accounts and exchange rates for eleven major economies. To this end, we apply a Markov-switching vector error correction model (MS-VECM) which is well-suited for this purpose, considering that the relationship between effective exchange rates and current accounts is potentially time-varying and not necessarily clear-cut. After testing for the long-run relationship, a consideration of a time-varying adjustment mechanism allows us to look at time-varying causality patterns. As a next step, we analyze the common exchange rate and current account dynamics on a global level based on common factors derived from a principal component analysis (PCA). We then turn to the relationship between those common factors according to our MS-VECM. As an extension and robustness check, we then re-run our analysis for a broader model at a monthly frequency including industrial production and short-term interest rates. This step also contains a study of long-run impacts of shocks between the variables and therefore allows for a direct assessment of the trade channel for exchange rate effects. Figure I presents an overview of our modeling cycle.[3] Finally, we discuss the implications of our results for the reduction of global imbalances through changes in exchange rates. Section 4 concludes.

*** Figure I about here ***

2 Effective exchange rates and the current account: background, previous studies, and methodological issues

2.1 Background and previous studies

This section starts with a brief reconsideration of theoretical and empirical linkages between effective exchange rates and current accounts. Our empirical framework allows for the possibility that causalities in terms of (time-varying) adjustment to long-run disequilibria go into both directions. We therefore elaborate on both possibilities and also include studies which deal with the role of exchange rate adjustment for restoring global imbalances.

To explain a potential causality running from effective exchange rates[4] to current accounts, consider the following simplified model equation introduced by Milesi-Ferretti (2008). The current account is expressed as the sum of trade balance (tb_t), net exports of services (nse_t), net receipts from interest, dividends and profits (nir_t), and net unilateral receipts (nur_t)

$$ca_t = tb_t + nse_t + nir_t + nur_t. \tag{1}$$

[3] It should be noted that the analysis of the trade balance is also considered as a sub-analysis of the current account.

[4] In our investigation, we rely on an external definition of the real exchange rate. Other studies correspond to the real exchange rate in internal terms as the ratio between prices of tradable and non-tradable goods with a relative increase in the price of tradable goods corresponding to a real depreciation of the domestic currency.

If we disregard net exports of services and net unilateral receipts, the current account can be written as the sum of trade balance and net investment income which is driven by the interest rate differential between interest on domestic assets (a) and loans (l) (Milesi-Ferretti, 2008).

$$ca_t = tb_t + ir_t^a a_{t-1} - ir_t^l l_{t-1}. \tag{2}$$

The change in the net foreign asset position (Δb_t) is given by the sum of the current account (ca_t), the capital gain or loss on the net foreign asset position (kg_t), capital account transfers and measurement errors (v_t)

$$\Delta b_t = b_t - b_{1-1} = ca_t + kg_t + v_t. \tag{3}$$

Based on these definitions, an impact of the exchange rate on the evolution of the current account and the net foreign asset position can work through a trade channel and a valuation channel (Milesi-Ferretti, 2008). On the one hand, an exchange rate appreciation could worsen the trade balance (possibly with a lag according to the J-Curve effect). In addition, the percentage change in the exchange rate influences the relative return earned on foreign assets, and can either worsen or improve the net external position through validation effects.[5] Our empirical setup in Section 3 picks up the distinction between the trade and the valuation channel by analyzing two settings where the effective exchange is either examined in connection to the current account or the trade balance.

In contrast, the idea of a reversed causality (running from the current account to the exchange rate) has been introduced in an early paper by Dornbusch and Fischer (1980) who emphasize the role of the current account within an asset market model of the nominal exchange rate. The main line of reasoning here is that asset markets determine the exchange rate at a point in time while the current account determines the path of the exchange rate through the net foreign asset position. Referring to an extension of the traditional monetary approach, Hooper and Morton (1982) provided the first empirical study that suggests that the current account is a useful determinant of the exchange rate.

[5] If for example, U.S. foreign liabilities are mainly denominated in dollars, while most U.S. foreign assets are denominated in a foreign currency, a depreciation of the dollar will improve the net foreign asset position (Milesi-Ferretti, 2008). Hence, trade and evaluation effects stemming from exchange rates should be qualitatively equivalent. Recent studies by Gourinchas and Rey (2007) and Lane and Shambaugh (2010) analyze the effects of real exchange rates on relative valuations of foreign assets and liabilities in greater detail. However, including these dynamics explicitly in our time-series framework is beyond the scope of our study since the most appropriate measure of net foreign asset positions is only available at an annual frequency.

A simple empirical equation for the nominal exchange rate e^n could for example be written as

$$e_t^n = (m_t - m_t^*) - (y_t - y_t^*) + (i_t - i_t^*) + tb_t, \tag{4}$$

with m as money supply, y as industrial production, i as the interest rate, and tb as the trade balance (Beckmann *et al.*, 2011). Variables with an asterisk refer to the foreign economy, whereas variables without refer to the domestic economy. While an in-sample link between exchange rates and the current account in the spirit of the above equation has partly been observed, current account dynamics fail to provide systematic superior exchange rate forecasts against the random walk benchmark (Rossi, 2013). In line with the so-called exchange disconnect puzzle, the general evidence suggests that the causal relation from current accounts to exchange rates is fragile and more frequently observed in terms of a long-run relationship in the spirit of our cointegration framework.

The question of whether exchange rates bear the potential to remove current account imbalances is also controversially discussed. The origin of underlying shocks is considered to be crucial: an exchange rate response to global imbalances is only likely to occur if the underlying shocks to exchange rates simultaneously lead to a closing of global imbalances (Rogoff, 2007). In this context, recent studies based on a new open-economy macroeconomics framework in the spirit of Obstfeld and Rogoff (1996) focus on the first causality from exchange rates to current accounts and suggest that current account improvements should be associated with a real exchange rate depreciation.[6] However, clear empirical evidence for this has not been established, although some studies have found predictive power of the change in the U.S. current account for exchange rate movements (Gourinchas and Rey, 2005; Rogoff, 2007). A recent study by Fratzscher *et al.* (2010) also raises doubts regarding an exchange rate adjustment since relative global asset prices rather than exchange rates are considered to be the key source of adjustment. Our framework does not allow direct predictions regarding the removal of imbalances but is nevertheless useful for providing policy implications against the background of the nonlinear and global dynamics between effective exchange rates and current accounts in the past.

In a nutshell, current accounts and real effective exchange rates should both be considered as endogenous in an empirical investigation. Both are simultaneously determined and a function of other variables such as interest rates or output gap (IMF, 2013). In this vein, most recent literature has

[6] In a series of papers, Obstfeld and Rogoff (2001, 2005, 2009) calibrate different scenarios of exchange rate and net foreign asset adjustments for reducing the U.S. current account deficit prior to the crisis based on a new open-economy macroeconomics framework. Depending on parameter choices such as the elasticity of substitution between tradables and non-tradables, the effective dollar exchange rate is expected to fall between 21 and 33%, according to their calculations (Obstfeld and Rogoff, 2009).

emphasized the importance of distinguishing between short-run and long-run shocks in modeling the real effective exchange rate, the current account and the relationship between them. The idea that the correlation between the real exchange rate and the current account depends on the source of shock mirrors the theoretical insights of Backus *et al.* (1994) and has been empirically analyzed by Chinn and Lee (2009) and Shibamoto and Kitano (2012). Based on a structural VAR for the G7 economies, the former argue that a theory-conform combination of real exchange rate depreciation and a current account surplus is more likely to be observed if temporary monetary policy shocks are the main driver of exchange rates.[7] Identifying a structural change during the nineties, Shibamoto and Kitano (2012) report slightly different results for some G7 economies but share the finding that permanent shocks drive the U.S. current account while temporary shocks drive the real exchange rate of the dollar.

As outlined in the Introduction, our study contributes to this string of the literature by considering the role of shocks from a cross-country and time-varying perspective. Rather than focusing on the source of shocks in separate countries, we analyze whether common shocks to real effective exchange rates and current accounts across countries share similar long- and short-run dynamics. This issue of restoring global imbalances through exchange rates boils down to the question of whether shocks to (global) real exchange rates and current account shocks share similar dynamics. However, it should be noted that our cross-section global perspective does not include the long-run constraint that the global sum of current account balances is zero. The application of a cointegration framework is motivated by the fact that we find current account imbalances to be sustainable. This result violates the long-run intertemporal budget constraint (solvency constraint), which implies a stationary current account, but it has been confirmed by other studies (Herwartz and Xu, 2008).

2.2 Methodological issues and contribution

The empirical assessment of exchange rate behavior in general is an extensively studied topic and this section will only briefly review a small part of the literature that is relevant to our study. Unsurprisingly, nonlinearities are a key ingredient when it comes to modeling exchange rate behavior. In a nutshell, two different kinds of framework have turned out to be useful in the context of recurring regime switches and cointegration: smooth transition and Markov-switching models.[8] Both

[7] Against this background, the pattern of dollar depreciation and a worsening current account observed in the United States is due to mostly permanent factors driving the real exchange rate.

[8] As another alternative, models with structural breaks or time-varying coefficients, which allow for different regimes, have been applied, for instance, by Goldberg and Frydman (2001, 2007) and Beckmann *et al.* (2011).

frameworks focus on regime switches in the adjustment mechanism and the short-run dynamics while relying on a constant long-run relationship.

The key difference is that Markov-switching models apply an exogenous stochastic switching process, while smooth transition models rely on endogenously determined switching, which is, for example, triggered by the degree of deviations from a fundamental value such as purchasing power parity (PPP).[9] A distinction between different exchange rate regimes and adjustment patterns through Markov-switching models has, for example, been applied in the studies by Sarno *et al.* (2004) and Sarno and Valente (2006) when analyzing different exchange rate adjustment to fundamentals deviations according to the canonical monetary exchange rate model and PPP.[10] The core idea is that the exchange rate is driven by fundamentals in the long-run while non-fundamental regimes can be identified in the short-run. More generally, Cheung and Erlandsson (2005) provide unambiguous general evidence for the presence of Markov-switching dynamics in exchange rates.[11] As outlined in the previous subsection, exchange rate and current account shocks might be triggered by several factors not included in our empirical setting, such as productivity shocks, changes in the exchange rate regime or demographical factors.[12] In this case, an endogenously determined switching approach seems inappropriate owing to the lack of an adequate transition variable. In addition, such an approach has already been adopted by Arghyrou and Chortareas (2008) when analyzing the relationship between current account imbalances and real exchange rates in the euro area, providing evidence for a nonlinear relationship.

For these reasons, we adopt a Markov-switching approach which is, for example, able to disentangle periods with and without adjustment to long-run equilibrium relying on a model with two regimes. One regime could thus be classified as an adjustment regime where errors are corrected and the other one into a bubble accumulation regime in which long-run deviations are not corrected. Overall, such a choice seems appropriate from both an economic and econometric point of view.[13] This enables us to disentangle the dynamics between real effective exchange rates and the current account into periods with and without adjustment to an existing long-run equilibrium.

[9] Models of this kind have been applied, for example, by Taylor *et al.* (2001) and Wu and Hu (2009).

[10] In two related studies, Frömmel *et al.* (2005a,b) reformulate the monetary model in annual changes and allow for changes in the long-run coefficients itself through a Markov-switching process.

[11] The first study to adopt a Markov-switching model in the context of exchange rates is provided by Engel (1994) and deals with exchange rate forecasting.

[12] Recent research has also emphasized the role of credit markets for the determination of current accounts (Kunieda and Shibata, 2005).

[13] In order to justify our choice, we have computed the logarithms of marginal likelihoods for our models by allowing for two and three regimes, respectively. As will be shown in Section 3.3 the fit appears to be better when two regimes are used in most of the cases.

The usefulness of our common factor approach in assessing cross-country dynamics applied in the second step of our analysis has also been verified in various studies. Engel *et al.* (2015) have provided evidence that accounting for common factors provide improved forecasts against the benchmark of traditional exchange rate models. The intuitive explanation is that such a framework efficiently exploits the cross-country dimension and co-movements between exchange rates. Berg and Mark (2015) provide a theoretical explanation for such third-country effects. One frequent finding is that common factors in exchange rates and fundamentals are cointegrated, with their relationship even matching theoretical predictions (Beckmann *et al.*, 2012). Including current accounts and trade balances, respectively, into this approach is a natural extension of such frameworks, considering our research question.

3 Data, empirical methodology and results

3.1 Data and preliminary tests

Our sample period contains quarterly and monthly data running from January 1980 to March 2013. We use trade weighted real effective exchange rates provided by the Bank for International Settlements (BIS).[14] The major advantage of this is that these series are available at a relatively long time period and that the corresponding time-varying weights are published. This allows the calculation of country-specific foreign quantities as a robustness check. The current-account-to-GDP ratio is taken from the World Bank, while short-term interest rates with a maturity of three months, industrial production indices, CPIs, and trade balances are taken from the OECD. We use nominal rather than real interest rates, since the former is directly influenced through monetary policy. For Hong Kong, industrial production is approximated by GDP through interpolation of the quarterly into a monthly series. In a similar fashion to Rose (1990), in order to obtain a measure of the real trade balance, we deflate the nominal trade balance by the product of the CPI and the nominal exchange rate. For our time series approach, we consider the following eleven economies, with the most of them being part of the G13: Australia, Canada, France, Germany, Hong Kong, Italy, Japan, Korea, Mexico, UK, and the USA.[15]

A crucial first step in our analysis is the application of unit root tests. While there is little doubt

[14] See Chinn (2006) for a detailed overview of different calculations for real effective exchange rates and a comparison of different weighting criteria depending on the topics under investigation.

[15] We do not consider the terms of trade as a possible determinant of current accounts or trade balances. Such an analysis is beyond the scope of our study, since it would correspond to disentangling prices and nominal exchange rate dynamics.

that real effective exchange rates, industrial production, trade balances, and interest rates are nonstationary, previous research has often considered the current account to be stationary. However, our results in most cases clearly suggest nonstationarity and therefore sustainability of current account imbalances.[16] Although this finding violates the theoretical intertemporal budget constraint, it is in line with the actual observations and with previous empirical findings (Herwartz and Xu, 2008). The full results of the unit root tests for the current account and effective exchange rate data are presented in Table I. All remaining results are available upon request.

*** Table I about here ***

It should also be mentioned that the accumulated real trade balances, a possible proxy of the net foreign asset position, are integrated of order two. Hence, using changes in the accumulated real trade balances (i.e. the real trade balance) is an adequate procedure in the context of cointegration.

3.2 MS-VECM

Figure II provides both current accounts and effective exchange rates for each economy. A quick glance suggests that both series are related over the long-run in many cases. We now tackle this question empirically without pre-assuming any causality. The framework we apply for each economy is an M-regime pth order MS-VECM, which in general allows for discrete regime shifts in the vector of deterministic terms D_t, the autoregressive part $\Gamma(L)(s_t)\Delta Y_{t-1}$, the long-run matrix $\Pi(s_t)$, and the variance-covariance matrix of the errors:

$$\Delta Y_t = \Gamma(L)(s_t)\Delta Y_{t-1} + \Pi(s_t)Y_{t-1} + \Phi(s_t)D_t + \varepsilon_t, \quad t = 1, \ldots, T, \tag{5}$$

where Δ denotes the difference operator and Y_t represents a K-dimensional vector of the observed time series consisting of a subset of the following elements as a starting point: $Y_t \subset [e_t, ca_t, ir_t]'$, depending on the model under observation. Details are provided in the next subsection. $\varepsilon_t = [\varepsilon_{1t}, \ldots, \varepsilon_{Kt}]'$ describes a K-dimensional vector of error terms with regime-dependent variance-covariance matrix $\Sigma(s_t)$, $\varepsilon_t \sim NIID(0, \Sigma(s_t))$. The $K \times K$ matrix lag polynomial $\Gamma(L)(s_t)$ of order p denotes the state-dependent short-run dynamics of the model. D_t gives the d-dimensional vector of deterministic terms

[16] More precisely, in neither case the unit root null is rejected at the 1% level for each of the three tests conducted. At the 5% level the current account appears to be stationary for Mexico and the UK. See Table I for details.

(i.e. dummy variables). The stochastic regime-generating process is assumed to be an ergodic, homogenous, and irreducible first-order Markov chain with a finite number of regimes, $s_t \in \{1, \ldots, M\}$, and constant transition probabilities

$$p_{ij} = Pr(s_{t+1} = j | s_t = i), \quad p_{ij} > 0, \quad \sum_{j=1}^{M} p_{ij} = 1 \quad \forall \quad i, j \in \{1, \ldots, M\}. \tag{6}$$

The first term in Equation (6) gives the probability for switching from regime i to regime j at time $t+1$, which is independent of the history of the process. p_{ij} is the element in the ith row and the jth column of P, the matrix of the transition probabilities with dimension $M \times M$.

We make use of a reduced rank ($r < K$) restriction of the state-dependent $K \times K$ long-run level matrix $\Pi(s_t)$ to account for the non-stationarity of the series. Thus $\Pi(s_t)$ can be fragmented into two $K \times r$ matrices $\alpha(s_t)$ and β such that $\Pi(s_t) = \alpha(s_t)\beta'$. β' gives the coefficients of the variables for r long-run relations, which are assumed to be constant over the whole sample period, while $\alpha(s_t)$ contains the regime-dependent adjustment coefficients describing the reaction of each variable to disequilibria from the long-run relations given by the r-dimensional vector $\beta' Y_{t-1}$. Thus, in our model, the most interesting distinction between regimes is the speed at which deviations from long-run equilibria are corrected, given by $\alpha(s_t)$. This allows for the possibility of asymmetries and changes in the adjustment process between exchange rates and current accounts, which is indicated in Figure II.

*** Figure II about here ***

Firstly, in order to identify the rank of $\Pi(s_t)$, i.e. the number of cointegrating relations r, and to estimate the coefficients of the r cointegrating vectors in β', we employ the maximum likelihood framework developed by Johansen (1988, 1991). Secondly, conditional on these cointegrating vectors, the regime-dependent adjustment parameters $\alpha(s_t)$, deterministic terms $\Phi(s_t)$, autoregressive coefficients $\Gamma(L)(s_t)$, and variance-covariance matrix $\Sigma(s_t)$ as well as the transition probabilities, are estimated using a multi-move iterative Gibbs sampling procedure. Saikkonen (1992) and Saikkonen and Luukkonen (1997) showed that the Johansen procedure provides consistent estimates for the cointegrating vectors even in the presence of regime-switching.[17]

[17]From a Bayesian perspective, it is possible that the posterior variance of all quantities involved is underestimated since β is treated as known. In this vein, Jochmann and Koop (2015) provide a full Bayesian approach to model a VECM with Markovian shifts. This framework is ideal to estimate a MS-VECM which allows for shifts not only in the short-run parameters such as the adjustment coefficients and the variance of the errors (like in our case) but also in the

In order to estimate our regime-dependent parameters, we define a $(1+Kp+r)M$-dimensional vector:

$$\begin{aligned}Z_t = [&\Delta Y_{t-1}\mathbb{1}(s_t = 1) \quad \dots \quad \Delta Y_{t-1}\mathbb{1}(s_t = M) \quad \dots \quad \Delta Y_{t-p}\mathbb{1}(s_t = 1) \quad \dots \quad \Delta Y_{t-p}\mathbb{1}(s_t = M) \\ &\beta' Y_{t-1}\mathbb{1}(s_t = 1) \quad \dots \quad \beta' Y_{t-1}\mathbb{1}(s_t = M) \quad D_t\mathbb{1}(s_t = 1) \quad \dots \quad D_t\mathbb{1}(s_t = M)]', \end{aligned} \tag{7}$$

where $\mathbb{1}(s_t = i)$ denotes an indicator function which equals 1 for regime i and 0 otherwise. Therefore, Equation (5) can be written in compact form as

$$Y = \Xi Z + \varepsilon, \tag{8}$$

with

$$Y = [\Delta Y_1 \quad \dots \quad \Delta Y_T], \quad Z = [Z_1 \quad \dots \quad Z_T], \quad \varepsilon = [\varepsilon_1 \quad \dots \quad \varepsilon_T]. \tag{9}$$

The coefficient matrix of order $K \times (1 + Kp + r)M$ is given by

$$\Xi = [\mu(s_t) \quad \Gamma_1(s_t) \quad \dots \quad \Gamma_p(s_t) \quad \alpha_1(s_t) \quad \dots \quad \alpha_r(s_t)]. \tag{10}$$

Given Equation (8), the cointegrating matrix β, and a series of states $\tilde{s}_T = \{s_1, \dots, s_T\}$, coefficient values are drawn from the normal-inverse Wishart posterior distribution with uninformative priors ν_{01}, …, ν_{0M}, N_0, F_0, W_{01}, …, W_{0M}.[18] We apply uninformative priors, in order to model the cointegrating vectors explicitly and to achieve parameter estimates that do not depend on the prior information (Francis and Owyang, 2005).[19]

In each iteration step, we draw Ξ and $\Sigma(s_t)$ for $s_t \in \{1, \dots, M\}$ from a distribution with $\nu_1, \dots, \nu_M$ degrees of freedom, precision matrix N, parameter means F, and variance-covariance matrices W_1, …, W_M, which are computed as follows for each regime i:

$$\nu_i = \nu_{0i} + \hat{T}_i, \quad N = N_0 + Z'Z, \quad F = N^{-1}\left(N_0 F_0 + Z'Z\hat{F}\right),$$

long-run relationships between the variables (i.e., β) and the cointegrating rank r. However, in line with previous studies we have followed the idea that cointegrating vectors are stable over the entire sample period (therefore also referred to as long-run relationships) and the potential nonlinearity between the variables is absorbed by the possibility that the adjustment to deviations from these long-run relationships are allowed to vary over time. A related approach to the one followed in this paper has been applied by Beckmann *et al.* (2014) to examine the relationship between global liquidity and commodity prices.

[18] We have used the following choices for the hyperparameters: $\nu_{0i} = 6 \; \forall \; i$, F_0 as a vector of zeros and N_0 as well as $W_{0i} \; \forall \; i$ as matrices of zeros. Those are intended to be weakly informative to prevent the Gibbs sampler from falling into an absorbing state.

[19] We also refer to Koop *et al.* (2009) for a different identification strategy of the prior for β, where the cointegration space is typically restricted to be a member of an orthonormal space called the Stiefel manifold.

$$W_i = \frac{\nu_0}{\nu} W_{0i} + \frac{\hat{T}_i}{\nu_i}\hat{\Sigma} + \frac{1}{\nu}\left(\hat{F} - F_0\right)' N_0 N^{-1} Z'Z \left(\hat{F} - F_0\right), \tag{11}$$

with $\hat{F} = (Z'Z)^{-1} Z'Y$ and $\hat{\Sigma} = \left(Y - Z\hat{F}\right)' \left(Y - Z\hat{F}\right)$. $\hat{T}_i$ denotes the number of periods spent in state i.

Then, conditional on the data series $\tilde{Y}_T$ and the drawn parameters Ξ and $\Sigma(s_t)\ \forall\ s_t$, the series of states $\tilde{s}_T$ is drawn from the posterior distribution $p\left(\tilde{s}_T|\tilde{Y}_T, \Xi, \Sigma(s_t)\forall s_t\right)$,[20] which is obtained from:

$$p\left(s_t|\tilde{Y}_t, \Xi, \Sigma(s_t)\forall s_t\right) = \frac{f\left(Y_t|\tilde{Y}_{t-1}, s_t, \Xi, \Sigma(s_t)\forall s_t\right) p\left(s_t|\tilde{Y}_{t-1}, \Xi, \Sigma(s_t)\forall s_t\right)}{\sum_{s_t} f\left(Y_t|\tilde{Y}_{t-1}, s_t, \Xi, \Sigma(s_t)\forall s_t\right) p\left(s_t|\tilde{Y}_{t-1}, \Xi, \Sigma(s_t)\forall s_t\right)}, \tag{12}$$

where

$$p\left(s_t|\tilde{Y}_{t-1}, \Xi, \Sigma(s_t)\forall s_t\right) = \sum_{s_{t-1}} p\left(s_t|s_{t-1}\right) p\left(s_{t-1}|\tilde{Y}_{t-1}, \Xi, \Sigma(s_t)\forall s_t\right), \tag{13}$$

and $p\left(s_{t-1}|\tilde{Y}_{t-1}, \Xi, \Sigma(s_t)\forall s_t\right)$ is given by each previous iteration step (Hamilton, 1989; Kim and Nelson, 1999). As common practice in such a case, the transition probabilities p_{ij} are derived within this algorithm by drawing from posteriors formed from Dirichlet conjugate distributions (Kim and Nelson, 1999; Francis and Owyang, 2005).

3.3 Long-run and short-run dynamics from a country perspective

We start this section with the results of our bivariate models for the economies under observation.[21] The main diagnostics for each model are shown in Table II.

*** Table II about here ***

According to the trace test, a long-run relationship is detected in each case except for Canada. The findings also show that each configuration provides satisfying results in terms of autocorrelation tests. As a next step, the character of the long-run relationships is considered. The estimated coefficients are provided in Table IV.

[20] For the states a Dirichlet prior with hyperparameter vector $(8, 2)$ is used in our application, where two regimes are applied (i.e. $M = 2$).

[21] In order to save space we do not report the skewness and kurtosis of the residuals of each equation, the tests for lag length determination, and the simulated values for the trace test statistic. All those diagnostics are available upon request.

*** Table IV about here ***

For each economy except Germany, real effective exchange rates and the current account are negatively related after rearranging the equation. This implies that a real appreciation coincides with a worsening of the current account, which is in line with the theoretical considerations mentioned in Section 2. In some cases, even the hypothesis that both series are exactly inversely related in the long-run cannot be rejected. The findings for Germany are the only exception, but the outcome might simply be driven by the fact that an appreciation of the euro has coincided with an improvement of the current account since the Millennium, as shown in Figure II.[22] Considering that our framework considers all variables to be endogenous, the adjustment dynamics need to be assessed in the next stage to clarify the issue of causality for each economy. In doing so, we have estimated the MS-VECM with two regimes (i.e. $M = 2$) as the most reasonable choice. This enables us to disentangle the dynamics between real effective exchange rates and the current account into periods with and without adjustment to an existing long-run equilibrium. One regime could thus be classified as an adjustment regime where errors are corrected and the other one into a bubble accumulation regime in which long-run deviations are not corrected.[23]

*** Table III about here ***

In most cases, the current account adjusts to disequilibria in one regime while there is no theory-conform adjustment in the second regime. Hence, the causality runs from exchange rates to current accounts but is only observed during specific time periods. Exceptions are the United States and Hong Kong, where the exchange rate adjusts in one regime while the current account adjusts in the other. Although this result suggests a more complex nature of the underlying causalities, an encouraging finding is that adjustment is observed throughout the sample period.

The approach up to this point has been restricted, owing to the fact that only effective exchange rates and current accounts have been considered. In the following we also introduce domestic nominal short-

[22]The results for Germany might simply be affected by its membership in the euro area. Therefore, exchange rate dynamics are determined somewhat independently of its domestic situation. The results might also be influenced by the reunification.

[23]In order to justify our choice, we have also computed the logarithms of marginal likelihoods of our trivariate models (including the effective exchange rate, the current account, and the interest rate as described below) with Markovian shifts by allowing for two and three regimes, respectively. According to Table III the fit appears to be better for most of our models with only two regimes since the logarithms of marginal likelihoods are higher for $M = 2$ compared to $M = 3$ in most of the cases. To allow comparability of our results, we have chosen $M = 2$ for all models.

term interest rates with a maturity of three months into our systems. On the one hand, this step is useful as an implicit robustness check regarding the results obtained up to this point. In addition, previous findings suggest that monetary policy shocks influence the relationship between exchange rates and current accounts (Chinn and Lee, 2009).[24] From a theoretical point of view, the long-run relationships identified should continue to hold in larger systems (Juselius, 2006). Except for France, where a cointegrating relationship is no longer detected,[25] the character of the relationship between current accounts and effective exchange rates and the time-varying causality pattern does not change. However, both the adjustment patterns and the effect of the interest rate suggest some differences: a rise in domestic interest rates either appreciates or depreciates the domestic exchange rate and improves or worsens the current account. For Korea and Mexico, current account adjustment turns out be insignificant for the extended model. In the latter case, including interest rates increases the number of long-run relationships and provides evidence of exchange rate adjustment.[26]

*** Tables V and VI about here ***

Summing up the results up to this point, we have shown robust evidence that exchange rates and current accounts share similar dynamics over the long-run for most economies. The fact that causality in terms of adjustment is only observed during specific periods mirrors the observation that countries might accumulate imbalances even if the exchange rate seems to move in the "correct" direction for removing imbalances. Another possible interpretation is that short-run exchange rate fluctuations, which display high volatility, do not trigger current account adjustment.

[24]We rely on nominal rather than real interest rates since the real effective exchange rate already includes price dynamics. Analyzing the real interest rate and the real effective exchange rate in one system might therefore produce misleading results.

[25]If the interest rate is restricted to zero in the long-run relationship, the results for France are equivalent to the previous setting.

[26]In addition, Table V provides the transition probabilities for both types of model. These show that both regimes are relative persistent since the probability of staying in a given regime lies around 0.9 in most of the cases. We have also used the regime classification measure (RCM) suggested by Ang and Bekaert (2002) to show that the regimes have been identified correctly. This measure is defined as $RCM(M) = 100M^2 \frac{1}{T_j} \sum_{t=1}^{T_j} \prod_{j=1}^{M} \tilde{p}_{j,t}$, where $\tilde{p}_{j,t}$ stands for the smoothed probability for regime j. The $RCM \in [0, 100]$ provides a degree of accuracy with which a model identifies regime switching behavior over the sample period under observation, with 0 representing a perfect regime classification performance and 100 denoting that the model fails to exhibit any information about the regime-dependence. According to Table VI regimes have been identified correctly for most of the models since the RCM is below 50 in most of the cases.

3.4 Common dynamics across countries

Having focused on the country perspective, we now turn to the global analysis of common dynamics across countries. Previous literature has focused on the question of whether the source of shocks drives the relationship between exchange rates and current accounts for particular economies. We focus on the more essential question of whether common exchange rate and current account shocks across countries are related in the long-run and in terms of causality.[27]

Once again, we consider a setting with two and three variables. Common factors for the effective exchange rate, the current account, and the interest rate are estimated by principal component analysis, according to Bai and Ng (2004). As is common practice, each series is taken as (logarithmic) first difference and then standardized so that each has a zero mean and a variance of unity. Otherwise, the results would be systematically affected by cross-country differences in variability. The first principal component derived from all individual quantities for each country explains the largest fraction of the total variance of the dataset in comparison to the remaining principal components and, therefore, suitably qualifies as a factor capturing international co-movement.

At this stage, we no longer restrict our sample to the G13 economies that are included in our dataset.[28] Instead, we apply the common factor approach to all economies, for which the narrow effective exchange rate measure is provided by the BIS. Starting with the bivariate approach, the results mirror the findings obtained on a country base: common shocks to effective exchange rates and the current account are negatively related over the long-run. This implies that common exchange rate shocks triggering an appreciation are strongly related to common shocks which worsen the trade balance. Hence, a transmission of global common shocks between both series is likely to occur. Unsurprisingly, a clear adjustment pattern is not detected, so that the causality in terms of shocks cannot be dissected.

Including common factors of interest rates again does not change the overall conclusion: the positive long-run relationship between the shocks of exchange rates and current accounts prevails, while interest rate shocks are also related to exchange rate shocks according to a second long-run relationship. Interestingly, exchange rate shocks react to interest rate shocks according to the adjustment coefficient. Hence, shocks introduced by monetary policy might influence the current account adjustment

[27] Previous estimations also aimed at distinguishing between fixed and flexible exchange rate regimes in the context of a global analysis in the spirit of Chinn and Wei (2013). This part of the analysis has been dropped owing to the fact that the classification of countries frequently differs across time.

[28] The full list of economies is provided under `http://www.bis.org/statistics/eer/`. We use the narrow index.

through exchange rates, although such an influence is unlikely to be systematic.

Altogether, our new perspective on shocks which drive exchange rates and current accounts provides some interesting insights: firstly, we find that global shocks are the main source of nonstationarity in current accounts and exchange rates. In line with the findings of Chinn and Lee (2009), the underlying shocks of exchange rates and current accounts are related. In terms of policy recommendations, the key question that arises is whether any systematic influence on these common factors can be obtained. In the best case, this is only possible through sustainable long-term global policies, which suggests that the exchange rate channel is unlikely to be a trigger by current account adjustment in the near future. In this regard, an important question is whether the observed effects occur through direct trade or valuation effects as discussed in Section 2. We examine this question as a next step by taking the trade balance into account.

3.5 Monthly analysis and robustness checks

Up to this point, we have considered the current account relative to GDP. However, to dissect the underlying dynamics and shocks, a more in-depth analysis might be necessary. For this reason, in the following we consider, within our system, trade balances, industrial production, effective exchange rates, and short-term interest rates. More precisely, we use interest rates relative to the United States and industrial production relative to OECD production.[29] For our common factor analysis, we solely rely on country measures of industrial production and interest rates since drawing common factors for similar relative differentials is not plausible.

By analyzing this system, we are able to compare the impact of monetary policy and exchange rate changes on exchange rates and the trade balance. In addition, a distinction between trade and valuation channels, as briefly described in Section 2, might be provided. The opposite causality, from trade balances to exchange rates, can also be analyzed in a broader context, with industrial production and interest rates as possible exchange rate determinants also included. Finally, we also examine the long-run impact of interest rates and industrial production shocks on effective exchange rates and the current account. This enables us to compare our findings to the studies like the one provided by Chinn and Lee (2009) who have applied VAR models to analyze the impact of shocks.[30]

[29] We only use the domestic interest rate for the United States.

[30] We do not report the results of a bivariate setting between trade balances and real exchange rates to keep the interpretation of our results transparent. However, these are available upon request.

Similar to our analysis based on quarterly data, we start with the results obtained for the individual economies. The findings are provided in Tables VII and VIII.

*** Tables VII and VIII about here ***

For Australia and Canada, no long-run relationship is found according to the results of the trace test. For this reason, we do not consider both economies from this stage on.[31] For France, Hong Kong, and Korea, two long-run relationships are identified. The estimates of the long-run coefficients are in line with the results for our quarterly dataset. In most cases, a real domestic depreciation coincides with an improvement of the domestic trade balance in real terms. Only the findings for Germany and France suggest an inverse relationship. Naturally, the adjustment effects become increasingly complex, since we have added a fourth variable to our system. As a result, two long-run relationships are observed more frequently. For Mexico, Korea, Hong Kong, and the U.S., the trade balance adjusts to long-run deviations from the equilibrium condition including exchange rates and trade balances in one regime. For these economies, we can confirm the importance of allowing for regime-switching dynamics. Considering the finding for the remaining economies, the trade channel can only partly explain the adjustment of current accounts observed for quarterly data in the previous subsections. Exchange rate adjustment is only correctly signed and significant in case of Germany. While income drives the adjustment pattern for the UK, relative interest rates show a theory-conform adjustment for most economies.

Finally, we turn to our global model derived from our common factor approach. Interestingly, exchange rate shocks adjust to deviations in the first regime while no adjustment of the trade balance is observed in one of the two regimes. The fact that industrial production adjusts in both regimes suggests that our previous finding that the current account relative to GDP adjusts might not be driven by the trade channel introduced in Section 2.1. Instead, valuation effects and production adjustment might be responsible for this finding.

As outlined previously, a great deal of research has focused on the source of shocks to effective exchange rates and current accounts/trade balances. As a final step, we therefore turn to the analysis of the driving forces of system. In doing so, the cointegrated VAR model is rearranged into its moving

[31] Since the currencies of both economies are often labeled as “commodity currencies”, movements in global commodity prices might be a main driver of exchange rate and current account dynamics.

average (MA) representation given below

$$Y_t = C \sum_{i=1}^{t} \varepsilon_i + \tau_0 + A_t, \tag{14}$$

where

$$C = \beta_\perp \left[\alpha'_\perp \left(-I + \sum_{i=1}^{p-1} \Gamma_i \right) \beta_\perp \right]^{-1} \alpha'_\perp, \tag{15}$$

$\tau_0 = C(Y_0 + \Phi D_t)$ depends on the initial values Y_0, and $A_t = \alpha(\beta'\alpha)^{-1} \sum_{i=0}^{\infty} (I - \beta'\alpha)^i \beta'(\varepsilon_i + \Phi D_i)$ is a stationary moving average process. Also note that $\alpha_\perp$ and $\beta_\perp$ denote the orthogonal complements of α and β. The MA representation of the cointegrated VAR model allows us to study the long-run impact of shocks or to conduct an analysis of the pushing forces of the system. Considering the several potential causalities that have been discussed in Section 2, such an analysis appears to be well suited to deliver further insights. Therefore, the long-run impact matrix C indicates how each variable is affected by accumulated shocks to other variables. Altogether, this part of the analysis provides a different representation of long-run causalities. Up to this point, we have adopted the switching adjustment coefficients as a measure of causality. At this stage, we draw conclusion based on the dissection of our linear benchmark model. The estimated coefficients of C are reported in Table IX.

*** Table IX about here ***

A first look at the main diagonal shows that the main underlying dynamics seem to be adequately specified, since nearly all elements turn out to be positive and significant. In the following we start with direct linkages between trade balances and exchange rates. For Korea, Mexico, and the U.S. a domestic appreciation worsens the trade balance in real terms. This pattern mirrors the overall findings in the previous subsections, where a theory-conform adjustment has been observed for those three economies. In most other cases, the exchange rate impact has the correct sign but turns out to be insignificant. For France, Italy, and the global model a reverse causality is observed: an improvement of the trade balance leads to a depreciation of the domestic economy. In contrast, the findings for Germany and Korea suggest that an improvement of the trade balance leads to an appreciation of the economy. Those ambiguous findings reflect theoretical considerations which are not necessarily clear-cut regarding the impact of changes in the net foreign asset position on (nominal) exchange rates.

As a next step, it seems useful to consider the role of interest rates and productivity shocks for the path of exchange rates and trade balances. A first key finding is that trade balances are more frequently driven by production shocks while exchange rates are more often influenced by interest rates shocks. For Japan, Korea, UK, the U.S., and Germany, an increase in the interest rate differential appreciates the domestic currency. At first sight, this finding violates the implications of uncovered interest rate parity (UIP). However, UIP can hardly be directly considered for real effective exchange rates and many studies have found that an increase in the relative interest rate appreciates the domestic exchange rate on a country level (Taylor and Sarno, 2004). On the opposite, an impact of industrial production on exchange rates is only found for Korea and France. Interestingly, exchange rate shocks for their turn influence industrial production for all countries except for France, Italy, Japan and also for the global factor model.

Trade balances are influenced by industrial production in case of Italy, the U.S., Korea, Mexico, and Japan with both negative and positive impacts being observed. An impact of interest rates on the trade balance is found for the U.S., Mexico, and Korea. Hence, we can confirm previous findings that a theory-conform relationship is mostly observed in cases where both exchange rates and trade balances are influenced by the same shocks. The finding that interest rate shocks negatively influence industrial production for most economies is also in line with theoretical considerations, since an expansionary monetary policy should increase production. The impact of industrial production on interest rates is less clear-cut. Unsurprisingly, monetary policy plays a major role for both the path of exchange rate and trade balances. While the exchange rate is more often affected directly, effects on the trade balance might occur through second round effects via exchange rates and industrial production.

At the global level, in terms of common shocks, interest rates and production shocks both also increase exchange rate shocks while they execute no influence on the trade balances. This pattern is in line with the results of Lee and Chinn (2006) that exchange rates are more influenced by long-run shocks compared to the current account. It is important to keep in mind that the common factors correspond to all economies provided by the BIS rather than the eleven economies we have analyzed on a country perspective. This might explain the different findings regarding the adjustment of industrial production.

4 Conclusion

Having analyzed the relationship between exchange rates and current accounts from a broad perspective, we find that the relationship between current accounts and effective exchange rates significantly varies between countries. In most cases, the long-run relationship between effective exchange rates and current accounts is in line with theory: a real appreciation coincides with a worsening of the current account relative to GDP. The causality mainly runs from the exchange rate to the current account. However, we have also shown that short-run dynamics are characterized by regime-switching with the current account only adjusting during specific periods, which we label the "adjustment regime". By contrast, there are also periods where no adjustment takes place. We also identify Germany and Canada as cases where the long-run relationship is either not in line with theory or no long-run relationship can be observed at all. Including interest rates into our analysis does not change the key results regarding the relationship between exchange rates and the current account. The findings for our monthly dataset still suggest a theory-conform link between exchange rates and trade balances. However, the underlying causalities in terms of adjustment seem to be more complex with a response of trade balances to long-run deviations observed less frequently compared to the current account. Taking the distinction between trade and valuation channel outlined in Section 2 into account, this suggests that valuation effects are an important driver of current account responses to exchange rate changes. We keep in mind that our conclusions regarding this issue are not based on a direct measure of the net foreign asset position, since the corresponding data is neither available for all economies nor the entire sample period under investigation.

Turning to a global perspective, we find that the cross-country trends, which drive exchange rates and current accounts, also share similar dynamics over the long-run. This finding also holds for the relationship between real trade balances and exchange rates. At first sight, this is an interesting finding in terms of global policy recommendations: if policymakers were able to influence the common trends which drive exchange rates, an impact on restoring global imbalances could be triggered. However, the simple and obvious problem is that policymakers are unable to systematically influence the trends that drive exchange rates. This is the lesson from various attempts at a coordinated exchange rate policy over recent decades. A reasonable aim of policymakers might be a volatility reduction of the underlying shocks, for example, through monetary policy shocks, which are an important driver of exchange rates according to our results. However, such a reduction would not guarantee any transmission to a reduction in global imbalances. On the whole, the exchange rate should obviously not be considered as a direct instrument but the current account in many cases seems likely to follow a depreciation or appreciation path. The general finding of country-specific regime switching patterns and the fact that trade balances do not respond to long-run deviations at a monthly frequency for many economies also underlines the complexity of the linkages between current accounts and exchange rates.

With regard to the correction of global imbalances, one should bear in mind that our approach provides an in-sample investigation without predicting future exchange rate movements. Several questions remain on the agenda for further research: disentangling nominal exchange rates and price dynamics, country-specific case studies and the simulation of potential policy shocks are obvious examples. A promising framework is a global vector error correction model in the spirit of Pesaran and Smith (2006), which combines country-specific long-run and short-run dynamics into a global model. A detailed normative approach based on panel estimation techniques has been provided by the external balance assessment methodology of the IMF.

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Tables

Table I: Unit root tests (Quarterly data)

Country	**e**				**ca**			
	ADF[Lags]	ADF-GLS[Lags]	NP[Lags]	KPSS[Lags]	ADF[Lags]	ADF-GLS[Lags]	NP[Lags]	KPSS[Lags]
Australia	-1.50[0]	-1.38[0]	-3.91[0]	30.96**[0]	-4.16**[0]	-1.83[0]	-6.38[0]	0.97**[0]
Canada	-1.59[1]	-1.62[1]	-5.63[1]	24.37**[1]	-2.29[0]	-2.18*[0]	-9.03*[0]	16.68**[0]
France	-3.50**[1]	-0.74[1]	-1.42[1]	26.50**[1]	-1.57[1]	-1.40[1]	-3.61[1]	25.70**[1]
Germany	-2.87[0]	-1.22[0]	-2.75[0]	3.46**[0]	1.12[0]	0.25[1]	0.53[1]	264.91**[1]
Hong Kong	-1.31[0]	-1.27[0]	-3.24[0]	46.75**[0]	-2.29[8]	-0.98[8]	-2.14[8]	30.14**[8]
Italy	-2.24[1]	-1.73[1]	-6.46[1]	6.76**[1]	-1.87[7]	-0.93[7]	-1.77[7]	17.07**[7]
Japan	-2.30[0]	-1.22[0]	-2.84[0]	16.69**[0]	-3.99**[0]	-0.80[0]	-1.25[0]	12.01**[0]
Korea	-2.53[0]	-1.16[0]	-2.96[0]	39.30**[0]	-3.51**[0]	-0.98[0]	-2.56[0]	9.48**[0]
Mexico	-2.99*[0]	-2.34*[0]	-10.20*[0]	7.69**[0]	-2.95*[2]	-2.76**[2]	-14.07**[2]	2.48**[2]
UK	-2.10[0]	-1.08[0]	-3.08[0]	29.88**[0]	-2.89*[1]	-2.55*[1]	-13.06*[1]	11.50**[1]
USA	-1.67[0]	-1.67[0]	-5.55[0]	35.00**[0]	-1.61[0]	-0.92[1]	-1.72[0]	131.35**[0]

Note: The table reports test statistics for three tests, viz. the augmented Dickey and Fuller (1979) test, the GLS-detrended version of the latter proposed by Elliott *et al.* (1996), the Ng and Perron (2001) MZ_α test and the Kwiatkowski *et al.* (1992) stationarity test. In each case we have used a test regression with an intercept, but without trend. The 5% and the 1% critical values are as follows: 5% (ADF) -2.88, (ADF-GLS) -1.94, (NP) -8.10, and (KPSS) 0.46; 1% (ADF) -3.48, (ADF-GLS) -2.58, (NP) -13.80, and (KPSS) 0.74. The lag length has been chosen according to the Schwarz criterion and the maximum lag length has been set to 12. In case of NP and KPSS, the spectral density has been estimated using a GLS-detrended AR process. KPSS tests the null of stationarity while all other test the null of a unit root. e denotes the effective exchange rate and ca denominates the current account. * denotes a rejection at a 5% level and ** at a 1% level.

Table II: Trace tests and autocorrelation tests (Quarterly data)

Country	Tests		p-values bivariate	p-values trivariate
Australia	Trace test	$H_0 : r = 0$	0.031	0.055
Australia	Trace test	$H_0 : r \leq 1$	0.675	0.811
Australia	AC test	LM(1)	0.325	0.203
Australia	AC test	LM(2)	0.202	0.011
Australia	AC test	LM(3)	0.168	0.468
Australia	AC test	LM(4)	0.942	0.050
Canada	Trace test	$H_0 : r = 0$	0.499	0.751
Canada	Trace test	$H_0 : r \leq 1$	0.666	0.776
France	Trace test	$H_0 : r = 0$	0.063	0.787
France	Trace test	$H_0 : r \leq 1$	0.065	0.973
France	AC test	LM(1)	0.000	
France	AC test	LM(2)	0.013	
France	AC test	LM(3)	0.036	
France	AC test	LM(4)	0.000	
Germany	Trace test	$H_0 : r = 0$	0.008	0.007
Germany	Trace test	$H_0 : r \leq 1$	0.111	0.103
Germany	AC test	LM(1)	0.663	0.218
Germany	AC test	LM(2)	0.779	0.560
Germany	AC test	LM(3)	0.050	0.071
Germany	AC test	LM(4)	0.352	0.482
Hong Kong	Trace test	$H_0 : r = 0$	0.001	0.002
Hong Kong	Trace test	$H_0 : r \leq 1$	0.050	0.276
Hong Kong	AC test	LM(1)	0.106	0.083
Hong Kong	AC test	LM(2)	0.000	0.164
Hong Kong	AC test	LM(3)	0.026	0.024
Hong Kong	AC test	LM(4)	0.000	0.000
Italy	Trace test	$H_0 : r = 0$	0.000	0.000
Italy	Trace test	$H_0 : r \leq 1$	0.047	0.310
Italy	AC test	LM(1)	0.200	0.000
Italy	AC test	LM(2)	0.017	0.003
Italy	AC test	LM(3)	0.513	0.161
Italy	AC test	LM(4)	0.000	0.000
Japan	Trace test	$H_0 : r = 0$	0.014	0.031
Japan	Trace test	$H_0 : r \leq 1$	0.183	0.367
Japan	AC test	LM(1)	0.722	0.109
Japan	AC test	LM(2)	0.994	0.793
Japan	AC test	LM(3)	0.843	0.235
Japan	AC test	LM(4)	0.249	0.030
Korea	Trace test	$H_0 : r = 0$	0.000	0.000
Korea	Trace test	$H_0 : r \leq 1$	0.052	0.028
Korea	AC test	LM(1)	0.0891	0.905
Korea	AC test	LM(2)	0.372	0.484
Korea	AC test	LM(3)	0.447	0.276
Korea	AC test	LM(4)	0.786	0.729
Mexico	Trace test	$H_0 : r = 0$	0.007	0.097
Mexico	Trace test	$H_0 : r \leq 1$	0.152	0.221
Mexico	AC test	LM(1)	0.296	0.075
Mexico	AC test	LM(2)	0.000	0.000
Mexico	AC test	LM(3)	0.577	0.466
Mexico	AC test	LM(4)	0.114	0.016
UK	Trace test	$H_0 : r = 0$	0.013	0.009
UK	Trace test	$H_0 : r \leq 1$	0.189	0.220
UK	AC test	LM(1)	0.034	0.001
UK	AC test	LM(2)	0.711	0.869
UK	AC test	LM(3)	0.426	0.631
UK	AC test	LM(4)	0.188	0.410
USA	Trace test	$H_0 : r = 0$	0.071	0.000
USA	Trace test	$H_0 : r \leq 1$	0.247	0.500
USA	AC test	LM(1)	0.664	0.225
USA	AC test	LM(2)	0.796	0.422
USA	AC test	LM(3)	0.427	0.757
USA	AC test	LM(4)	0.850	0.680
Global	Trace test	$H_0 : r = 0$	0.001	0.002
Global	Trace test	$H_0 : r \leq 1$	0.050	0.043
Global	AC test	LM(1)	0.106	0.100
Global	AC test	LM(2)	0.000	0.000
Global	AC test	LM(3)	0.026	0.028
Global	AC test	LM(4)	0.000	0.000

Note: The table reports the p-values for testing the null hypotheses of no cointegration, of at most one cointegrating relationship, and of no serial correlation up to order four. The former two hypotheses are tested by the trace test proposed by Johansen (1988) and the latter by a Lagrange multiplier (LM) test. r denotes the cointegration rank.

Table III: Logarithms of marginal likelihoods of trivariate models

Country	**M = 2**	**M = 3**
Australia	-37.47	-39.38
Germany	170.45	165.18
Hong Kong	-268.56	-271.27
Italy	-6.73	-5.08
Japan	165.63	175.91
Korea	-124.18	-126.30
Mexico	-282.46	-286.46
UK	-32.61	-36.64
USA	166.32	164.02
Global	-830.27	-817.75

Note: The table reports the logarithms of marginal likelihoods of trivariate models (including the effective exchange rate, the current account, and the interest rate for quarterly data) with Markovian shifts by allowing for 2 and 3 regimes, respectively. The marginal likelihood has been computed through an BIC approximation.

Table IV: Coefficient estimates (Quarterly data)

Country			**bivariate**			**trivariate**			
			e	ca	constant	e	ca	ir	constant
Australia	Long-run coefficients		0.948 (14.145)	1	0	1	1	0.44 (0.805)	−0.527 (−1)
	Adjustment coefficients	Regime 1	−0.001 (−0.179)	−0.183 (−1.410)		0.004 (1.406)	−0.280 (27.794)	−0.196 (−1.367)	
		Regime 2	0.020 (1.040)	−0.579 (−1.835)		0 (0.065)	−0.253 (−2.525)	−0.021 (−0.247)	
France	Long-run coefficients		−0.214 (−11.495)	−0.214 −11.495	1				
	Adjustment coefficients	Regime 1	0.005 (1.314)	0.994 (0.170)					
		Regime 2	−0.040 (−2.577)	4.170 (3.854)					
Germany	Long-run coefficients		1	−0.002 (−0.766)	−4.604 (−433.942)	1	−0.008 (−2.142)	−0.008 (−2.142)	−4.541 (−176.002)
	Adjustment coefficients	Regime 1	−0.208 (−0.852)	−2.902 (−0.293)		−0.124 (2.156)	−0.778 (9.383)	−1.076 (−0.970)	
		Regime 2	−0.266 (−0.913)	−1.985 (−0.146)		−0.317 (−2.954)	−6.823 (−1.678)	4.584 (1.189)	
Hong Kong	Long-run coefficients		1	0.075 (5.587)	−5.046 (−46.743)	21.947 (4.649)	1	1.343 (6.538)	−114.805 (−5.111)
	Adjustment coefficients	Regime 1	−0.017 (−2.233)	−4.801 (−4.612)		0 (−0.787)	0.006 (3.537)	−0.220 (−0.626)	
		Regime 2	0.010 (0.625)	−0.574 (−1.648)		5.321 (−0.001)	−0.215 (15.080)	−0.013 (−0.765)	
Italy	Long-run coefficients		11.313 (2.711)	1	−51.269 (−2.682)	14.551 (3.973)	1	0 (0)	−65.870 (−3.928)
	Adjustment coefficients	Regime 1	0.001 (1.373)	−0.476 (−5.710)		−0.001 (1.892)	−0.444 (7.226)	−0.073 (−0.884)	
		Regime 2	−0.000 (−0.036)	−0.244 (−1.056)		−0.003 (−0.344)	−0.362 (−1.100)	0.175 (0.615)	
Japan	Long-run coefficients		1	0.458 (4.046)	−5.844 (−18.374)	1	1.137 (5.075)	0.195 (2.254)	−8.279 (−10.887)
	Adjustment coefficients	Regime 1	0.018 (0.760)	0.098 (0.289)		0.010 (1.454)	−0.089 (3.183)	0.009 (−0.099)	
		Regime 2	0.008 (0.570)	−0.442 (−1.635)		0.011 (2.095)	−0.125 (22.888)	0.060 (1.463)	
Korea	Long-run coefficients		7.039 (−1.611)	1.000	−34.106 (−1.654)	1	1	0	−5.871 (−6.443)
						1	0	−0.039 (−7.537)	−4.401 (−71.946)
	Adjustment coefficients	Regime 1	0.004 (2.599)	−0.201 (−3.840)		0.002 (−1.344)	−0.138 (−0.007)	4.868 (−0.061)	
						−0.038 (−0.073)	0.649 (0.749)	3.743 (−0.657)	
		Regime 2	0.001 (0.653)	−0.142 (−1.465)		0.006 (2.507)	−0.120 (−1.384)	−0 (−0)	
						−0.037 (4.743)	−0.112 (0.073)	0.909 (0.748)	
Mexico	Long-run coefficients		1	1	−2.990 (−5.974)	1	0.004 (3.529)	0.004 (3.529)	−4.697 (−123.442)
	Adjustment coefficients	Regime 1	0.039 (1.839)	−0.171 (−0.853)		−0.207 (0.754)	−0.292 (−0.631)	−2.255 (−0.923)	
		Regime 2	0.001 (0.567)	−0.142 (−2.906)		−0.248 (3.259)	−0.403 (−0.301)	3.227 (0.786)	
UK	Long-run coefficients		1	1	−3.129 (−7.288)	0.535 (4.033)	1	−0.087 (−1.174)	0
	Adjustment coefficients	Regime 1	−0.003 (−1.202)	−0.364 (−2.655)		−0.002 (4.910)	−0.172 (−0.063)	−0.153 (15.788)	
		Regime 2	−0.001 (−0.292)	−0.122 (−1.212)		−0.001 (1.986)	−0.336 (−3.257)	0.021 (0.551)	
USA	Long-run coefficients		1	0.038 (1.755)	−4.522 (−69.223)	1	0.045 (4.427)	−0.024 (−5.192)	−4.399 (−93.499)
	Adjustment coefficients	Regime 1	−0.695 (−1.792)	0.609 (0.088)		−0.015 (1.647)	−1.318 (10.862)	−1.976 (−1.292)	
		Regime 2	−0.013 (−0.368)	−0.751 (−3.361)		−0.048 (−0.654)	−1.576 (−3.027)	−0.174 (−0.225)	
Global	Long-run coefficients		0.171 (2.403)	1	−0.849 (−1.588)	0.215 (3.516)	1	0	−0.855 (−1.820)
						1	0	−0.914 (−3.898)	0
	Adjustment coefficients	Regime 1	−0.057 (−0.111)	−0.330 (−0.829)		0.015 (0.991)	−0.410 (−0.083)	−0.258 (11.513)	
						−0.034 (0.151)	0.071 (1.004)	0.071 (0.117)	
		Regime 2	−0.085 (−0.173)	−0.303 (−0.729)		0.144 (2.097)	0.134 (0.086)	3.161 (1.541)	
						−0.142 (4.524)	−0.049 (−0.062)	−1.310 (−0.841)	

Note: The table reports long-run and adjustment coefficients with t-statistics in parentheses. e denotes the effective exchange rate, ca denominates the current account, and ir gives the short-term interest rate. The term global corresponds to the common factors of the variables.

Table V: Transition probabilities

Country		**bivariate**		**trivariate**	
		Regime 1	Regime 2	Regime 1	Regime 2
Australia	Regime 1	0.912 (11.465)	0.267 (2.042)	0.938 (24.668)	0.048 (1.412)
	Regime 2	0.088 (1.112)	0.733 (5.612)	0.062 (1.622)	0.952 (28.000)
France	Regime 1	0.957 (37.711)	0.123 (1.871)	-	-
	Regime 2	0.043 (1.700)	0.877 (13.372)	-	-
Germany	Regime 1	0.859 (7.423)	0.180 (1.449)	0.919 (15.015)	0.187 (2.149)
	Regime 2	0.141 (1.216)	0.820 (6.591)	0.081 (1.321)	0.813 (9.345)
Hong Kong	Regime 1	0.825 (12.313)	0.067 (2.094)	0.877 (15.121)	0.058 (1.381)
	Regime 2	0.175 (2.612)	0.933 (29.156)	0.123 (2.121)	0.942 (22.429)
Italy	Regime 1	0.966 (47.386)	0.145 (1.859)	0.864 (8.858)	0.208 (1.891)
	Regime 2	0.034 (1.685)	0.855 (10.962)	0.136 (1.391)	0.792 (7.200)
Japan	Regime 1	0.766 (5.458)	0.203 (1.809)	0.907 (22.675)	0.097 (2.310)
	Regime 2	0.234 (1.663)	0.797 (7.102)	0.093 (2.325)	0.903 (21.500)
Korea	Regime 1	0.835 (10.705)	0.207 (2.226)	0.863 (16.922)	0.131 (2.729)
	Regime 2	0.165 (2.115)	0.793 (8.527)	0.137 (2.686)	0.869 (18.104)
Mexico	Regime 1	0.739 (7.229)	0.054 (2.145)	0.842 (9.791)	0.167 (1.856)
	Regime 2	0.261 (2.556)	0.946 (37.246)	0.158 (1.837)	0.833 (9.256)
UK	Regime 1	0.830 (11.033)	0.205 (2.397)	0.877 (15.661)	0.089 (1.978)
	Regime 2	0.170 (2.257)	0.795 (9.312)	0.123 (2.196)	0.911 (20.244)
USA	Regime 1	0.718 (6.119)	0.047 (1.109)	0.806 (7.858)	0.132 (1.650)
	Regime 2	0.282 (2.406)	0.953 (22.684)	0.194 (1.897)	0.868 (10.850)
Global	Regime 1	0.806 (5.373)	0.193 (1.222)	0.941 (11.476)	0.317 (2.099)
	Regime 2	0.194 (1.293)	0.807 (5.108)	0.059 (0.720)	0.683 (4.523)

Note: The table reports the estimated transition probabilities with t-statistics in parentheses for both types of quarterly data model.

Table VI: Regime classification measure

Country	bivariate	trivariate
Australia	12.18	65.32
France	22.61	-
Germany	95.59	55.22
Hong Kong	17.90	26.08
Italy	12.12	46.34
Japan	67.00	16.06
Korea	6.44	56.67
Mexico	16.99	38.55
UK	53.40	31.50
USA	22.74	47.42
Global	1.41	76.82

Note: The table reports the regime classification measure (RCM) proposed by Ang and Bekaert (2002) computed for both types of quarterly data model. This measure is defined as $RCM(M) = 100M^2 \frac{1}{T_j} \sum_{t=1}^{T_j} \prod_{j=1}^{M} \tilde{p}_{j,t}$, where $\tilde{p}_{j,t}$ stands for the smoothed probability for regime j. The RCM provides a degree of accuracy with which a model identifies regime switching behavior over the sample period under observation. $RCM \in [0, 100]$, with 0 representing a perfect regime classification performance and 100 denoting that the model fails to exhibit any information about the regime-dependence.

Table VII: Trace tests and autocorrelation tests (Monthly data)

Country	Tests		p-values
France	Trace test	$H_0 : r = 0$	0.000
		$H_0 : r \leq 1$	0.010
		$H_0 : r \leq 2$	0.045
	AC test	LM(1)	0.000
		LM(2)	0.128
		LM(3)	0.110
		LM(4)	0.217
Germany	Trace test	$H_0 : r = 0$	0.000
		$H_0 : r \leq 1$	0.093
	AC test	LM(1)	0.003
		LM(2)	0.357
		LM(3)	0.156
		LM(4)	0.116
Hong Kong	Trace test	$H_0 : r = 0$	0.000
		$H_0 : r \leq 1$	0.000
		$H_0 : r \leq 2$	0.352
	AC test	LM(1)	0.116
		LM(2)	0.000
		LM(3)	0.116
		LM(4)	0.104
Italy	Trace test	$H_0 : r = 0$	0.014
		$H_0 : r \leq 1$	0.349
	AC test	LM(1)	0.001
		LM(2)	0.017
		LM(3)	0.017
		LM(4)	0.210
Japan	Trace test	$H_0 : r = 0$	0.031
		$H_0 : r \leq 1$	0.233
	AC test	LM(1)	0.000
		LM(2)	0.009
		LM(3)	0.071
		LM(4)	0.228
Korea	Trace test	$H_0 : r = 0$	0.000
		$H_0 : r \leq 1$	0.000
		$H_0 : r \leq 2$	0.009
	AC test	LM(1)	0.000
		LM(2)	0.153
		LM(3)	0.082
		LM(4)	0.408
Mexico	Trace test	$H_0 : r = 0$	0.000
		$H_0 : r \leq 1$	0.067
	AC test	LM(1)	0.000
		LM(2)	0.000
		LM(3)	0.029
		LM(4)	0.050
UK	Trace test	$H_0 : r = 0$	0.001
		$H_0 : r \leq 1$	0.059
	AC test	LM(1)	0.000
		LM(2)	0.074
		LM(3)	0.057
		LM(4)	0.958
USA	Trace test	$H_0 : r = 0$	0.009
		$H_0 : r \leq 1$	0.256
	AC test	LM(1)	0.000
		LM(2)	0.063
		LM(3)	0.137
		LM(4)	0.164
Global	Trace test	$H_0 : r = 0$	0.000
		$H_0 : r \leq 1$	0.735
	AC test	LM(1)	0.150
		LM(2)	0.021
		LM(3)	0.039
		LM(4)	0.164

Note: See Table II for details.

TABLE VIII: COEFFICIENT ESTIMATES (MONTHLY DATA)

Country			e_t	tb_t	$\tilde{y}_t$	$\tilde{ir}_t$	constant
France	Long-run coefficients		1		−0.486	0.018	−4.233
					(−5.849)	(5.379)	(−57.259)
			1	−0.039			−3.659
				(−4.231)			(−16.633)
	Adjustment coefficients	Regime 1	0.002	−1.031	−0.024	−3.045	
			(0.147)	(−0.472)	(−1.795)	(−2.421)	
			−0.001	0.379	0.006	1.031	
			(−0.170)	(0.547)	(1.391)	(2.400)	
		Regime 2	0.011	−3.424	0.006	−1.541	
			(0.691)	(−1.182)	(0.279)	(−1.438)	
			−0.004	1.095	−0.004	0.521	
			(−0.717)	(1.155)	(−0.595)	(1.436)	
Germany	Long-run coefficients		1	−0.003		0.007	−4.562
				(−2.971)		(2.251)	(−353.995)
	Adjustment coefficients	Regime 1	−0.015	−0.521	−0.075	−3.177	
			(−1.054)	(−0.322)	(−3.680)	(−3.622)	
		Regime 2	−0.024	1.125	−0.016	−0.509	
			(−2.149)	(0.407)	(−0.955)	(−1.368)	
Hong Kong	Long-run coefficients		1	0.013		0.041	−5.013
				(2.351)		(8.022)	(−130.267)
			22.167		1		−116.458
			(27.323)				(−29.943)
	Adjustment coefficients	Regime 1	−0.001	−8.453	−0.000	−0.017	
			(−1.548)	(−1.599)	(−1.582)	(−1.117)	
			0.000	10.655	−0.001	0.007	
			(0.615)	(0.926)	(−3.797)	(0.599)	
		Regime 2	−0.000	−3.399	0.000	−0.001	
			(−0.153)	(−0.853)	(0.186)	(−0.255)	
			−0.000	41.387	−0.001	0.001	
			(−0.463)	(2.022)	(−2.400)	(0.141)	
Italy	Long-run coefficients		−0.043	−0.043	1	−0.048	0.172
			(−3.116)	(−3.116)		(−7.344)	(2.551)
	Adjustment coefficients	Regime 1	0.003	−0.504	0.003	1.424	
			(0.604)	(−1.222)	(0.343)	(1.902)	
		Regime 2	0.001	−0.406	0.014	0.599	
			(0.157)	(−0.950)	(2.347)	(0.931)	
Japan	Long-run coefficients		1	1		3.723	
						(7.289)	
	Adjustment coefficients	Regime 1	0.000	−0.006	0.000	−0.023	
			(1.542)	(−0.807)	(−2.322)	(−2.375)	
		Regime 2	0.000	0.006	0.000	−0.003	
			(1.963)	(0.637)	(0.295)	(−0.755)	
Korea	Long-run coefficients		1		0.152		−4.709
					(5.413)		(−158.448)
			1	0.234		0.073	−5.319
				(6.261)		5.259	(−64.016)
	Adjustment coefficients	Regime 1	−0.027	0.040	−0.057	0.224	
			(−2.237)	(0.065)	(−4.807)	(0.487)	
			0.009	−0.528	0.000	−0.103	
			(1.742)	(−1.856)	(−0.005)	(−0.376)	
		Regime 2	−0.015	−0.268	−0.031	1.745	
			(−1.612)	(−1.061)	(−3.148)	(2.768)	
			0.007	−0.014	−0.013	−1.074	
			(1.736)	(−0.076)	(−2.388)	(−3.122)	
Mexico	Long-run coefficients		1	0.238	−3.725	−0.007	−4.541
				(4.253)	(−5.398)	(−3.339)	(−87.019)
	Adjustment coefficients	Regime 1	0.010	−0.380	0.008	−0.426	
			(0.576)	(−3.256)	(1.864)	(−0.252)	
		Regime 2	0.023	−0.396	0.025	−2.414	
			(0.428)	(−1.650)	(2.418)	(−0.452)	
UK	Long-run coefficients		−0.019	−0.019	1	−0.069	0.047
			(−3.594)	(−3.594)		(−5.982)	(1.398)
	Adjustment coefficients	Regime 1	0.003	0.319	−0.014	0.089	
			(0.277)	(0.338)	(−2.521)	(0.435)	
		Regime 2	−0.017	−0.483	0.003	1.697	
			(−1.764)	(−1.084)	(0.614)	(4.334)	
USA	Long-run coefficients		1	0.010		−0.104	−3.708
				(3.265)		(−5.426)	(−20.810)
	Adjustment coefficients	Regime 1	−0.004	−1.722	0.001	−0.027	
			(−1.163)	(−2.835)	(1.038)	(−0.388)	
		Regime 2	−0.007	−0.687	0.006	0.314	
			(−1.602)	(−1.360)	(3.475)	(2.002)	
Global	Long-run coefficients		−0.206	−2.001	0.198	1	−5.429
			(−1.906)	(−1.687)	(2.011)		(−2.019)
	Adjustment coefficients	Regime 1	0.017	0.000	−0.042	−0.014	
			(1.245)	(0.239)	(−4.878)	(−1.606)	
		Regime 2	0.030	0.002	−0.032	0.007	
			(2.049)	(0.727)	(−3.234)	(1.468)	

Note: See Table IV for details. Moreover, tb gives the trade balance, $\tilde{y}$ denotes industrial production relative to the OECD industrial production, and $\tilde{ir}$ denominates the short-term interest rate relative to the U.S. interest rate.

Table IX: MA Representation (Monthly data)

Country		e_t	tb_t	$\tilde{y}_t$	$\tilde{ir}_t$
France	e_t	−0.311 (−0.528)	−0.006 (−2.817)	0.267 (3.085)	0.000 (0.075)
	tb_t	−46.027 (−1.099)	0.359 (2.232)	3.571 (0.582)	−0.174 (−1.011)
	$\tilde{y}_t$	−2.089 (−1.192)	−0.008 (−1.225)	0.837 (3.260)	−0.004 (−0.570)
	$\tilde{ir}_t$	−11.807 (−0.528)	−0.242 (−2.817)	10.107 (3.058)	0.007 (0.075)
Germany	e_t	0.498 (3.120)	0.001 (4.200)	−0.006 (−0.296)	−0.005 (−4.243)
	tb_t	4.798 (0.145)	0.635 (10.612)	4.203 (1.043)	−0.098 (−0.417)
	$\tilde{y}_t$	−1.024 (−2.599)	0.002 (2.256)	0.631 (13.180)	−0.005 (−1.709)
	$\tilde{ir}_t$	−64.984 (−3.472)	0.123 (3.693)	2.667 (1.172)	0.603 (4.546)
Hong Kong	e_t	0.760 (0.517)	−0.001 (−0.539)	−0.146 (−1.124)	0.002 (0.343)
	tb_t	57.571 ((0.510))	0.346 (2.414)	2.858 (0.288)	0.107 (0.271)
	$\tilde{y}_t$	−16.840 (−0.517)	0.022 (0.539)	3.227 (1.124)	−0.039 (−0.343)
	$\tilde{ir}_t$	−36.476 (−0.520)	−0.082 (−0.925)	2.688 (0.435)	−0.076 (−0.311)
Italy	e_t	1.523 (12.034)	−0.005 (−1.694)	0.074 (0.945)	−0.007 (−1.043)
	tb_t	4.256 (0.779)	0.699 (5.549)	−7.240 (−2.183)	0.383 (1.366)
	$\tilde{y}_t$	−0.031 (−0.279)	−0.004 (−1.664)	0.694 (10.335)	−0.009 (−1.564)
	$\tilde{ir}_t$	−5.854 (−0.901)	−0.712 (−4.763)	20.868 (5.301)	−0.522 (−1.571)
Japan	e_t	1.251 (9.787)	0.005 (2.049)	0.096 (0.572)	0.062 (2.326)
	tb_t	4.888 (1.759)	0.631 (11.045)	7.853 (2.155)	0.618 (1.072)
	$\tilde{y}_t$	−0.033 (−1.101)	0.000 (−0.030)	0.751 (19.313)	−0.003 (−0.508)
	$\tilde{ir}_t$	−0.895 (−2.153)	−0.093 (−10.856)	−1.159 (−2.126)	−0.099 (−1.149)
Korea	e_t	0.321 (2.474)	0.007 (2.013)	−0.098 (−3.715)	0.003 (1.892)
	tb_t	−9.334 (−2.198)	0.245 (2.292)	−2.309 (−2.662)	−0.132 (−2.282)
	$\tilde{y}_t$	−2.111 (−2.474)	−0.043 (−2.013)	0.648 (3.715)	−0.022 (−1.892)
	$\tilde{ir}_t$	25.719 (1.896)	−0.883 (−2.581)	8.810 (3.179)	0.381 (2.057)
Mexico	e_t	1.026 (15.618)	0.017 (1.421)	−0.256 (−1.073)	0.000 (0.010)
	tb_t	−3.073 (−7.974)	0.195 (2.711)	9.084 (6.262)	0.026 (4.528)
	$\tilde{y}_t$	0.105 (4.425)	0.020 (4.475)	0.437 (4.869)	−0.002 (−4.389)
	$\tilde{ir}_t$	−13.437 (−2.239)	−1.415 (−1.261)	37.137 (1.643)	1.636 (18.442)
UK	e_t	1.262 (12.779)	0.002 (1.223)	0.142 (1.021)	0.015 (1.910)
	tb_t	−4.654 (−1.375)	0.547 (12.727)	3.576 (0.750)	0.134 (0.500)
	$\tilde{y}_t$	−0.116 (−2.987)	0.000 (−0.324)	0.738 (14.354)	0.001 (0.370)
	$\tilde{ir}_t$	−0.739 (−0.732)	−0.125 (−11.863)	10.257 (7.214)	−0.024 (−0.305)
USA	e_t	0.774 (3.923)	−0.002 (−1.517)	0.077 (0.317)	0.035 (2.361)
	tb_t	−44.694 (−2.512)	0.461 (5.087)	91.534 (4.170)	3.378 (2.547)
	$\tilde{y}_t$	0.184 (2.380)	0.001 (2.042)	0.956 (10.033)	−0.011 (−1.908)
	$\tilde{ir}_t$	3.352 (1.228)	0.027 (1.969)	9.078 (2.696)	0.641 (3.149)

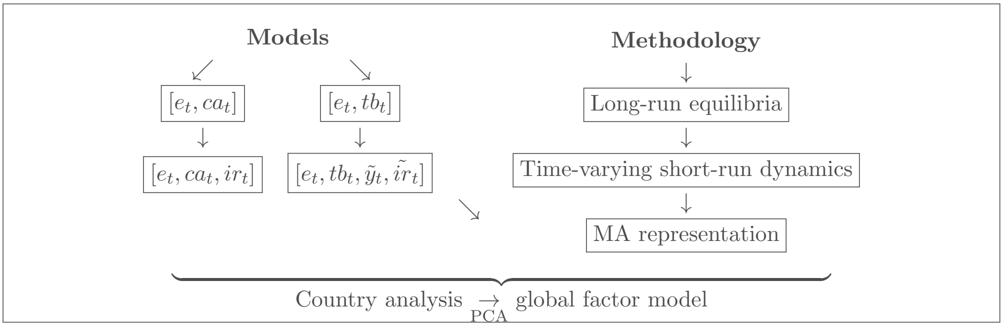
Note: The table reports the coefficient estimates for the long-run matrix C of the moving average (MA) representation of the cointegrated VAR model. e denotes the effective exchange rate, tb denominates the trade balance, $\tilde{y}$ gives industrial production relative to the OECD industrial production, and $\tilde{ir}$ represents the short-term interest rate relative to the U.S. interest rate.

Table IX *continued*

Country		e_t	tb_t	$\tilde{y}_t$	$\tilde{ir}_t$
Global	e_t	1.011 (5.578)	−1.428 (−1.960)	0.738 (3.023)	1.143 (2.527)
	tb_t	−0.009 (−0.760)	0.558 (11.949)	−0.013 (−0.848)	−0.009 (−0.295)
	$\tilde{y}_t$	0.552 (4.154)	1.838 (3.442)	0.345 (1.930)	−0.982 (−2.961)
	$\tilde{ir}_t$	0.057 (0.406)	0.133 (0.237)	0.064 (0.337)	1.677 (4.800)

Figures

Figure I: Model cycle



Note: e denotes the effective exchange rate, ca denominates the current account, and ir gives the short-term interest rate. Moreover, tb represents the trade balance, $\tilde{y}$ denotes industrial production relative to the OECD industrial production, and $\tilde{ir}$ denotes the short-term interest rate relative to the U.S. interest rate. PCA stands for principal component analysis.

Figure II: Effective exchange rates and current accounts

Data Appendix

Table X: Time series

Series	Frequency	Data source
Trade weighted real effective exchange rate	Q & M	BIS
Current-account-to-GDP ratio	Q	World Bank
Nominal short-term money market interest rate with maturity of 3 months	Q & M	OECD
Industrial production index	M	OECD
Consumer price index	M	OECD
Trade balance	M	OECD

Note: The table provides information about all time series used in our study. We rely on both quarterly (Q) and monthly data (M) for a sample period running from January 1980 to March 2013 and our study includes the following economies: Australia, Canada, France, Germany, Hong Kong, Italy, Japan, Korea, Mexico, UK, and the USA. See Section 3.1 for details.