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Flexible exchange rates in emerging markets: shock absorbers or drivers of endogenous cycles?

Karsten Kohler and Engelbert Stockhammer

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Flexible exchange rates in emerging markets: shock absorbers or drivers of endogenous cycles?*

Karsten Kohler[†] Engelbert Stockhammer[‡]

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Abstract

While flexible exchange rates are commonly regarded as shock absorbers, heterodox views suggest that they can play a pro-cyclical role in emerging markets. This article provides theoretical and empirical support for this view. Drawing on post-Keynesian and structuralist theory, we propose a simple model in which flexible exchange rates in conjunction with external shocks become endogenous drivers of boom-bust cycles, once financial effects from foreign-currency debt are accounted for. We present empirical evidence for regular cycles in nominal US-dollar exchange rates in several emerging markets that are closely aligned with cycles in economic activity. An econometric analysis suggests the presence of a cyclical interaction mechanism between exchange rates and output, in line with the theoretical model, in Chile, South Africa, and partly the Philippines. Further evidence indicates that such exchange rate cycles cannot exclusively be attributed to external factors, such as commodity prices, US monetary policy or the global financial cycle. We therefore argue that exchange rate cycles in emerging markets are driven by the interplay of external shocks and endogenous cycle mechanisms. Our argument implies that exchange rate management may be beneficial for macroeconomic stability.

Keywords: Exchange rates, emerging markets, boom-bust cycles, structuralism, global financial cycle, commodity prices

JEL Codes: E12, E32, F31, C32

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[†]Economics Department, Leeds University Business School. E-Mail: k.kohler@leeds.ac.uk.

[‡]Department of European and International Studies. King's College London. E-Mail: engelbert.stockhammer@kcl.ac.uk

1 Introduction

Whether flexible exchange rates are beneficial for macroeconomic stability is a long-standing economic debate. After some spectacular breakdowns of fixed and semi-fixed exchange rate regimes in the 1990s (e.g. 1992-93 in the EEC, 1994 in Mexico, 1997-98 in East Asia), a view became prominent whereby countries should either completely give up their monetary sovereignty or adopt freely floating exchange rates (Eichengreen, 1994; Fischer, 2001). The putative benefit of fully flexible exchange rates is considered to be their role as shock absorbers that facilitate macroeconomic adjustment after adverse external shocks. This view also informed policy recommendations by the International Monetary Fund and the World Bank to developing and emerging market economies (DEEs) (Gabor, 2010; Rodrik, 2006).

By contrast, post-Keynesian and structuralist economists have long doubted the stabilising features of flexible exchange rates for DEEs. It has been argued that flexible exchange rates often play a pro-cyclical role in boom-bust cycles in DEEs, which is related to foreign-currency debt in the private sector (Harvey, 2010; Kohler, 2019; Ocampo, 2002, 2016; Stiglitz et al., 2006). Exchange rate appreciation during booms reduces the value of foreign-currency debt, which improves balance sheets and stimulates spending, whereas depreciation induces contractionary deleveraging. This 'financial channel' of exchange rates in DEEs has recently gained further support from researchers at the Bank for International Settlements (BIS) (Avdjiev et al., 2019; Banerjee et al., 2020; Kearns and Patel, 2016). In addition, the literature on currency hierarchies and subordinated financialisation of DEEs argues that flexible exchange rates invite speculative behaviour by international investors that amplifies exchange rate volatility (Andrade and Prates, 2013; Bortz and Kaltenbrunner, 2017; de Paula et al., 2017; Kaltenbrunner and Painceira, 2015). Consequently, heterodox economists endorse a strategy of smoothing exchange rate fluctuations through foreign exchange intervention, so-called managed floating, to improve macroeconomic stability; ideally supported by controls on short-term capital movements (Frenkel, 2006; Ferrari-Filho and De Paula, 2008; Guzman et al., 2018; Ocampo, 2002).[1]

This article contributes to the debate on the pro-cyclical role of flexible exchange rates in DEEs through a theoretical and econometric analysis. Drawing on post-Keynesian and structuralist theory, we first develop a simple model in which flexible exchange rates can become endogenous drivers of macroeconomic fluctuations. Our key innovation is to com-

[1] The idea that (sterilised) foreign exchange intervention can enhance macroeconomic stability has recently also gained traction in mainstream macroeconomic theory (see Benes et al. 2015; Ghosh et al. 2016).

bine the financial channel of exchange rates, according to which currency depreciations are contractionary, with an external adjustment channel through which output contractions feed back into exchange rate revaluation, so that endogenous cyclical fluctuations between exchange rates and output may emerge. Procyclicality is thereby hypothesised to stem from an endogenous cycle mechanism. We then present empirical results from a spectral analysis documenting *periodic* exchange rate cycles for South Africa, Chile, Mexico, South Korea and the Philippines, and show that these cycles are often closely aligned with cycles in real gross domestic product (GDP), in line with the theoretical model.

Our main contribution is an econometric analysis that formally tests for the presence of an endogenous interaction mechanism between the exchange rate and GDP that can give rise to boom-bust cycles. This allows us to assess whether the pro-cyclical role of flexible exchange rates discussed in the post-Keynesian and structuralist literature and the recent work at the BIS indeed stems from our hypothesised mechanism. Our focus on the presence of endogenous cycle mechanisms sets our approach apart from previous econometric studies that compared macroeconomic volatility of fixed and flexible exchange rate regimes (Broda, 2004; Hoffmann, 2007; Obstfeld et al., 2019). Estimation results from vector-autoregressions yield robust evidence for the presence of such a cyclical interaction mechanism in South Africa and Chile, and to a lesser extent for the Philippines – the countries with the longest spells of (semi-)flexible exchange rate regimes in our sample period (1972-2017). For Mexico and South Korea there is some, but less robust, evidence for a cycle mechanism; possibly because these countries underwent major crises and changing exchange rate regimes during the first part of the sample period. For Brazil and Thailand, there is no evidence for a cycle mechanism; in Brazil, this is arguably because of numerous chaotic exchange rate episodes, whereas Thailand had a fixed exchange rate until 1997. Our results thus confirm that endogenous exchange rate cycle mechanisms are more likely to occur in flexible exchange rate regimes. Given that many large DEEs switched to floating exchange rates in the last two decades, we expect this mechanism to become increasingly relevant.

Our analysis offers a new perspective on the relationship between external and internal drivers of boom-bust cycles in DEEs. Recently, both mainstream and heterodox scholars have highlighted external sources of domestic macrofinancial instability, such as commodity price swings, US monetary policy, and changing risk perceptions of financial investors (Bonizzi and Kaltenbrunner, 2018; Cunha et al., 2020; IMF, 2012; Kalemli-Özcan, 2019; Ocampo, 2016; Rey, 2015). In seminal contributions, Rey (2015, 2016) has questioned the insulating properties of flexible exchange rates based on the presence of a 'global financial cycle' in risky asset prices that is strongly correlated with US monetary policy. Our stylized model accounts for these external factors and indeed requires exogenous shocks to set cyclical dynamics in motion. In our estimations, we thus control for various external drivers. However, we emphasise the critical role of *endogenous* interaction mechanisms that transform exogenous shocks into cyclical swings between exchange rates and domestic economic activity. To assess the relative importance of external factors, we also examine co-movements in exchange rates across our sample of DEEs and find that they are much less synchronised than those of major advanced economies (AEs). External factors, such as the global financial cycle, thus do not seem to tell the full story, and we suggest combining them with endogenous cycle mechanisms to explain exchange-rate driven boom-bust cycles in DEEs.[2]

Our argument casts doubt on the benefits of fully flexible exchange rates as shock absorbers. On the contrary, the presence of a financial channel of exchange rates in conjunction with an external adjustment channel can turn it into a variable that amplifies macroeconomic instability. With respect to policy, our analysis thus lends support to managed floating that reduces some of the volatility in exchange rates but allows for more flexibility than pegged exchange rates (Frenkel, 2006; Ferrari-Filho and De Paula, 2008; Guzman et al., 2018; Ocampo, 2002).

The remainder of the article is structured as follows. Section 2 discusses the notion of a pro-cyclical exchange rate and provides a stylized model of endogenous cycles. Section 3 presents evidence from spectral analysis for the presence of periodic exchange rate cycles in DEEs and their relationship to cycles in economic activity. Sections 4 and 5 discuss our econometric approach and present evidence for the presence of a cycle mechanism between exchange rates and output in line with the theoretical model. Section 6 discusses some additional results on the relative importance of external drivers of cyclical fluctuations: commodity prices, US monetary policy, and global risk aversion. The last section summarises and spells out theoretical and political implications of our analysis.

2 Pro-cyclical exchange rates: existing literature and a stylized model

Large parts of the post-Keynesian and structuralist literature on boom-bust cycle episodes in DEEs were developed under the impression of the collapsing fixed exchange rate regimes in

[2]At the same time, our argument does not imply that flexible exchange rates are the only relevant driver of macroeconomic instability in DEEs and acknowledges the complex and country-specific sources of specific boom-bust cycle episodes.

the 1980s and 1990s, especially in East Asia and Latin America (e.g. Arestis and Glickman 2002; Cruz et al. 2006; Frenkel and Rapetti 2009; Palma 1998; Taylor 1998). Correspondingly, those accounts focussed on the *real* exchange rate whose dynamics were largely driven by domestic price inflation. The shift of many large DEEs towards more flexible exchange rates after those crises raises the question of the role of *nominal* exchange rates in boom-bust cycles episodes.

According to a mainstream argument that goes back at least to Milton Friedman (1953), flexible nominal exchange rates act as shock absorbers and should thus curb cyclical dynamics. In this view, a change in foreign demand or interest rates will require a change in domestic macroeconomic variables to re-establish equilibrium. Flexible exchange rates facilitate this process (e.g. by swiftly improving price competitiveness through depreciation) and thereby reduce the adjustment pressure on other critical macroeconomic variables such as consumer prices and interest rates, which is argued to reduce overall macroeconomic volatility. Modern incarnations of this argument can be found in benchmark New Keynesian general equilibrium models (Galí and Monacelli, 2005) which predict higher macroeconomic volatility of fixed compared to floating exchange rate regimes. Several econometric studies compare the macroeconomic volatility of fixed and flexible exchange rates, and seem to provide empirical support for lower macroeconomic volatility of the latter (Broda, 2004; Edwards and Yeyati, 2005; Hoffmann, 2007; Obstfeld et al., 2019).

By contrast, structuralist development economists such as José Antonio Ocampo (2002: 8) have long argued that flexible exchange rates can assume a pro-cyclical role in DEEs:[3]

> Exchange rate fluctuations have significant wealth effects in countries with large net external liabilities. The capital gains generated by appreciation during booms further fuels spending booms, whereas wealth losses generated by depreciation have the opposite effect ... Thus, the wealth effects of exchange rate variations are pro-cyclical in debtor countries.

The main factor behind these ‘wealth’ or ‘balance sheet effects’ is the foreign-currency denomination of most of the external debt held in DEEs’ private sectors, which generates currency mismatches on balance sheets (Adler et al., 2020; Chui et al., 2018; Eichengreen et al., 2007). A nominal depreciation then raises the domestic value of foreign debt, which reduces firms’ net worth and tightens financial constraints. As a result, firms cut back on their expenditures in order to deleverage. Opposite dynamics take place during periods of appreciation. This sets apart DEEs from AEs, whose private sectors typically borrow in

[3] A similar argument has been made in Stiglitz et al. (2006: Ch. 7), as well as Harvey (2010) and de Paula et al. (2017). Kohler (2019) provides a formal model.

home currency (or have access to foreign currency even in times of financial distress). As a result, fluctuations in exchange rates tend to have different effects on economic activity in DEEs compared to AEs.

Recently, economists at the BIS have started to investigate this phenomenon in more detail (Avdjiev et al., 2019; Banerjee et al., 2020; Kearns and Patel, 2016). They distinguish the more conventional 'trade channel of exchange rates', whereby currency depreciations have expansionary effects on economic activity through their effects on export performance, from the 'financial channel of exchange rates' that operates through wealth effects as described by Ocampo (2002, 2016). Empirical studies show that the financial channel tends to offset the trade channel in DEEs, whereas it is much weaker in AEs (Kearns and Patel, 2016). The contractionary effects of currency depreciation in DEEs appears to stem mostly from business investment (Banerjee et al., 2020) and the effect is stronger for firms in countries with flexible exchange rates (Avdjiev et al., 2019). In addition, the trade channel in DEEs is typically weak due to the invoicing of exports in US dollars, which mutes the standard price-competitiveness effect of currency depreciation (Adler et al., 2020).

A possible implication of this financial channel of exchange rates that has been less explored, albeit being very much in the spirit of the post-Keynesian and structuralist argument, is its potential contribution to endogenous fluctuations in economic activity. Indeed, post-Keynesian and structuralist approaches argue that macroeconomic cycles are at least partially endogenous (Taylor, 2004: Ch. 9). At their heart, endogenous cycles stem from dynamic interaction mechanisms between key macroeconomic variables, such as private debt and investment, or wage shares and employment. Applying this perspective to the interaction between exchange rates and economic production, we argue that an endogenous cycle mechanism may arise if exchange rate depreciations not only drag down output, but if the reduction in output also leads to exchange rate revaluation through an external adjustment channel.

In order to provide intuition for such a cyclical interaction mechanism, we combine the financial channel of exchange rates with an external adjustment channel in a simple macroeconomic model. We stress that this is a highly stylized model whose sole purpose is to emphasise the endogenous cycle mechanism between exchange rates and output. The model abstracts from many other factors that may be relevant for DEEs, such as inflationary and distributional effects of exchange rate dynamics (see Ribeiro et al. 2017), external debt service (see Bortz et al. 2018), and long-run effects of exchange rate undervaluation on technological change (see Frenkel 2008; Guzman et al. 2018; Razmi et al. 2012).[4]

[4]See Kohler (2021) for a survey of post-Keynesian and structuralist models of boom-bust cycles in DEE.

The core components of the model are similar to the classic Mundell-Fleming model. While this model has lost the prominence it enjoyed in the 1960s and 1970s, it continues to be a foundation for applied analyses in both heterodox (e.g. Asada et al. 2003; Guschanski and Stockhammer 2020) and mainstream macroeconomics (e.g. Blanchard et al. 2010; Gourinchas 2017). By augmenting this model with the financial channel of exchange rates and introducing a dynamic extension that allows for cyclical dynamics, we give it a more post-Keynesian twist. An important feature of our model is that it encompasses both external and endogenous sources of cyclical dynamics, as we will illustrate below.

We start out with a static version, whose equations are given by:[5]

$$Y^d = A + NX \tag{1}$$

$$A = A(Y, i, s); \qquad A_Y \in (0,1), A_i < 0, A_s < 0 \tag{2}$$

$$NX = NX(Y, Y^{d*}, s); \qquad NX_Y < 0, NX_{Y^{d*}} > 0, NX_s > 0 \tag{3}$$

$$F = F(i - i^*, \rho); \qquad F_{i-i^*} > 0, F_\rho < 0 \tag{4}$$

$$\Delta R = F + NX. \tag{5}$$

where Y^d is domestic aggregate demand, composed of domestic absorption (A) and net exports (NX). Domestic absorption depends positively on output (Y), negatively on the domestic interest rate (i), and negatively on the spot exchange rate (s) (defined as the price of foreign currency in domestic currency, so that an increase in s implies a depreciation of the domestic currency). The last assumption is unconventional and captures the financial channel of exchange rates, whereby depreciation against the US dollar tightens borrowing constraints and discourages private spending. Depreciation might also be contractionary due to adverse distributional effects (Ribeiro et al., 2017). Net exports are a negative function of domestic output (through import demand) and positively related to foreign demand (Y^{d*}) and other exogenous factors that improve export performance, such as commodity prices. The last term captures the conventional trade channel of exchange rates whereby depreciations improve export performance through their effect on price competitiveness, provided that the Marshall-Lerner condition holds.[6] Net capital inflows F are assumed to be a function of the interest rate differential between the domestic rate (i), which is determined exogenously by domestic monetary policy (Serrano and Summa, 2015), and the foreign rate i^* (e.g.

Kohler (2019) presents a more complex model of endogenous boom-bust cycles that are driven by the financial channel of exchange rates that also considers external debt dynamics.

[5]All variables are expressed in deviations from their steady state values. The notation A_Y represents the partial derivative of A with respect to Y.

[6]A violation of the Marshall-Lerner condition would reinforce the mechanism highlighted in this paper.

the US policy rate). In addition, they are negatively related to risk perceptions of global investors ρ. The last equation is the balance-of-payments, with ΔR representing changes in foreign reserves.[7]

In a dynamic extension, actual output adjusts sluggishly to changes in aggregate demand: $Y_{t+1} - Y_t = \alpha(Y_t^d - Y_t)$ (with $\alpha > 0$), e.g. due to time lags in economic production and in the procurement of intermediate inputs.[8] Likewise, balance-of-payments disequilibria ($\Delta R \neq 0$) lead to gradual adjustments in the exchange rate: $s_{t+1} - s_t = -\beta(F_t + NX_t)$ (with $\beta > 0$), such that losses in reserves are accompanied by currency depreciation and vice versa.[9] The sluggish adjustment of the exchange rate may reflect the widespread practice of managed floating in DEEs, where pressures on the balance-of-payments lead to adjustments in both the exchange rate and in the stock of foreign reserves through foreign exchange intervention (Frankel, 2019). We then have:

$$Y_{t+1} - Y_t = \alpha[A_t(Y_t, i_t, s_t) + NX_t(Y_t, Y_t^{d*}, s_t) - Y_t] \tag{6}$$

$$s_{t+1} - s_t = -\beta[F_t(i_t - i_t^*, \rho_t) + NX_t(Y_t, Y_t^{d*}, s_t)]. \tag{7}$$

The Jacobian matrix, which collects the first-order partial derivatives of the system in (6)-(7), is given by:

$$J(Y_t, s_t) = \begin{bmatrix} \frac{\partial Y_t}{\partial Y_{t-1}} & \frac{\partial Y_t}{\partial s_{t-1}} \\ \frac{\partial s_t}{\partial Y_{t-1}} & \frac{\partial s_t}{\partial s_{t-1}} \end{bmatrix} = \begin{bmatrix} 1 + \alpha(A_Y + NX_Y - 1) & \alpha(A_s + NX_s) \\ -\beta NX_Y & -\beta NX_s + 1 \end{bmatrix}. \tag{8}$$

Importantly, the sign of the term $\alpha(A_s + NX_s)$ will depend on the relative importance of the financial and the trade channel of exchange rates. If the financial channel dominates the trade channel, as seems to be the case in many DEEs (Kearns and Patel, 2016), we have $\alpha(A_s + NX_s) < 0$. By contrast, the term $-\beta NX_Y$ will always be positive, as an increase in domestic demand reduces net exports via an increase in import demand. Taken together, we thus have opposite algebraic signs on the off-diagonal elements of the Jacobian matrix: $\frac{\partial Y_t}{\partial s_{t-1}} \frac{\partial s_t}{\partial Y_{t-1}} < 0$. Economically, this constitutes a cyclical interaction mechanism, whereby two dynamic variables act upon each other in opposite directions. Stockhammer et al. (2019) show formally that this is a necessary condition for the emergence of cyclical dynamics in simple bivariate models.[10]

[7] For simplicity, the model abstracts from interest payments on foreign debt.

[8] See, e.g., Asada (1995) for this type of specification in a Kaldorian open economy model.

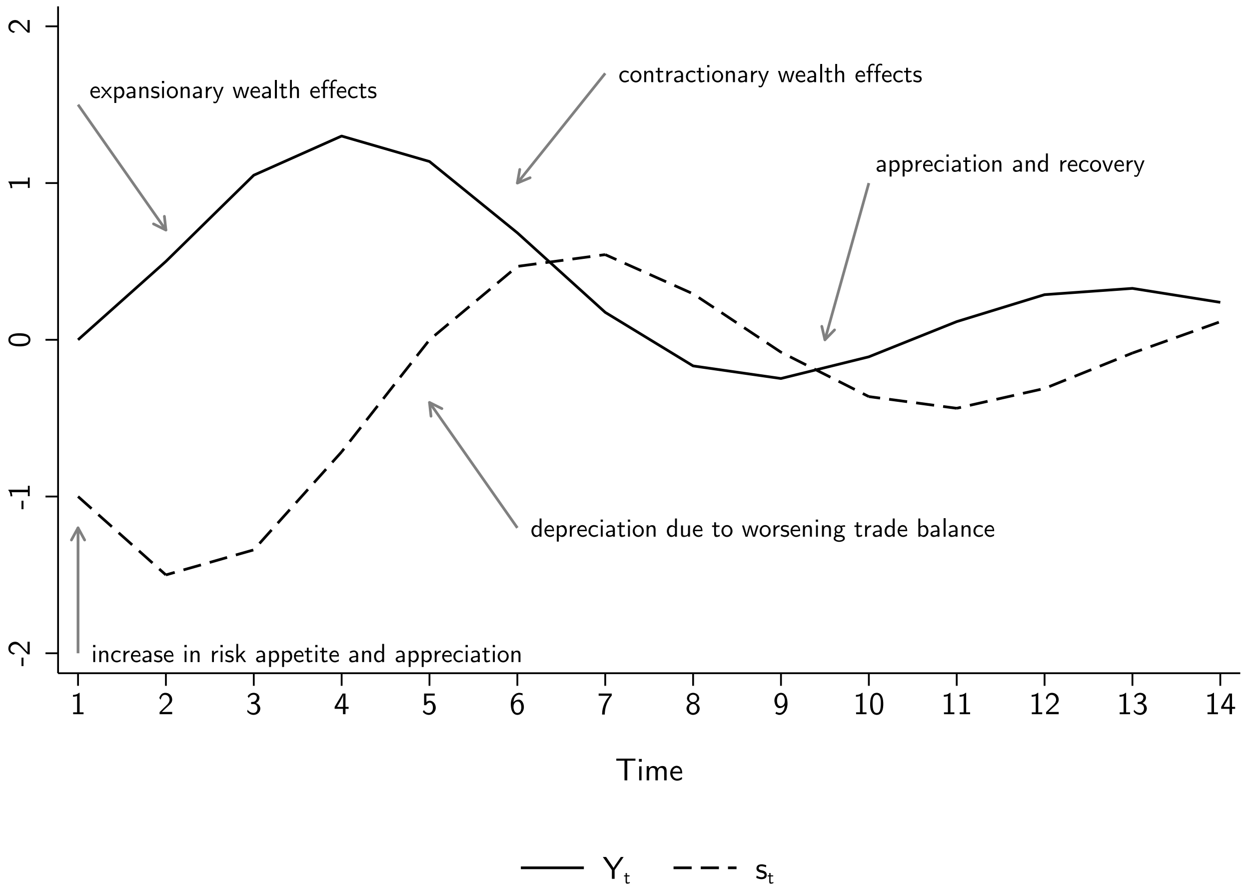
[9] See, e.g., Bhaduri (2003); Botta (2017); Chiarella et al. (2006) and Kohler (2019) for similar specifications.

[10] To see this, observe that the eigenvalues of the Jacobian matrix in (8) are the roots of the characteristic equation $\lambda^2 - \lambda Tr + Det = 0$, where Tr and Det are the trace and determinant of the Jacobian matrix,

Figure 1 displays a numerical simulation of a linear version of the model (6)-(7). To appreciate the interplay between exogenous shocks and endogenous cycle mechanisms, consider a scenario where the risk aversion of foreign investors (ρ) suddenly decreases and remains reduced for several periods until it gradually returns to normality. Such an increase in foreign risk appetite comes with a sustained appreciation of the exchange rate and economic expansion due to the financial channel of exchange rates. However, as the economy booms and the trade balance worsens, the exchange rate eventually begins to depreciate. This triggers contractionary wealth effects and turns the boom into a bust.

respectively. The roots of the characteristic equation are given by $\lambda_{1,2} = \frac{Tr \pm \sqrt{Tr^2 - 4Det}}{2}$. Complex roots, which give rise to oscillations, emerge when the discriminant of this expression becomes negative. This requires $(J_{11} + J_{22})^2 - 4(J_{11}J_{22} - J_{12}J_{21}) < 0$, (where J_{ij} are the elements of the Jacobian matrix) which simplifies to $(J_{11} - J_{22})^2 + 4(J_{12}J_{21}) < 0$. From this, it is immediate that $J_{12}J_{21} < 0$ is a necessary condition for complex eigenvalues (Stockhammer et al., 2019).

Figure 1: Model simulation: a shock to risk appetite triggers pro-cyclical exchange rate dynamics



Notes: Numerical simulation of a linear version of (6)-(7) with Y_t: output; s_t: nominal exchange rate (price of foreign currency in domestic currency). Parameterisation of (8): $J_{11} = 0.6$, $J_{12} = -0.5$, $J_{21} = 0.7$, $J_{22} = 0.7$. For this parameterisation, the roots of (8) are a pair of complex conjugates. Intercepts were set to zero, so that the equilibrium is at $(0, 0)$. A shock-sequence ($\rho_t = 0.8\rho_{t-1}$) was added to (7) to emulate a sustained increase in risk appetite, where ρ_t was initialised at minus unity in $t = 1$. For simplicity, it was assumed that $\frac{\partial s_t}{\partial \rho_{t-1}} = 1$.

In this way, our simple framework combines external shocks, which are frequently highlighted as sources of macroeconomic instability in DEEs (Cunha et al., 2020; Ocampo, 2016; Kalemli-Özcan, 2019; Rey, 2015), with an endogenous interaction mechanism that generates periodic cyclical fluctuations.

3 Empirical evidence of exchange rate cycles

From the theoretical argument of an endogenous cycle mechanism, one would expect to find regular 'exchange rate cycles' (Ocampo, 2002: 13) that are of similar length to fluctuations in economic output. In this section, we use spectral techniques to assess the periodicities in exchange rates and output for a group of seven DEEs over the (maximum) period 1972Q1 to 2019Q3: South Africa, Chile, the Philippines, Mexico, South Korea, Brazil, and Thailand.[11] The exchange rate regime is expected to be an important source of cross-country differences, as the endogenous cycle mechanism we are interested in requires a sufficient degree of exchange rate flexibility. We thus consider a heterogeneous group of countries to check whether the exchange rate regime indeed makes a difference. Our sample selection is thus governed by a combination of data availability constraints[12] and the need for heterogeneity.

We identify a first group of countries in our sample with relatively long spells of semi-flexible or flexible exchange rate regimes[13] (South Africa, Chile, and with exceptions the Philippines)[14]; a second, intermediate, group with episodes of both fixed and flexible exchange rate regimes that were interrupted by currency crashes (Mexico and South Korea); and a third group with fixed or chaotic exchange rates throughout most of the sample period (Thailand and Brazil).[15] We hypothesise a decreasing chance of finding regular cycles in exchange rates across these three groups.

To study cyclical properties of nominal exchange rates, we focus on the (logged) bilateral nominal exchange rate with the US dollar (henceforth XR), defined as domestic currency unit per foreign currency unit. The US dollar is the dominant currency for trade invoicing and external borrowing by firms in DEEs (Adler et al., 2020; Avdjiev et al., 2019). In line with the literature on the financial channel of exchange rates (Avdjiev et al., 2019; Banerjee et al., 2020; Kearns and Patel, 2016), we focus on *nominal* exchange rates as a key determinant of the real exchange rate.[16] This allows us to isolate nominal exchange rate dynamics from changes in relative prices (which are driven by different economic mechanisms). Our preferred detrending method is the regression filter proposed in Hamilton (2018), but we also compare

[11] Detailed information on the dataset can be found in Appendix A.

[12] In particular, we are constrained by the small number of DEEs that had floating exchange rate regimes for more than two decades.

[13] We use the exchange rate regime classification in Ilzetzki et al. (2019). See Appendix B for details.

[14] South Africa had a float since 1973 (with a parallel market between 1985 and 1995), Chile since 1983 (with a managed regime in the 1990s), and the Philippines had (semi-)flexible exchange rates most of the time, interrupted only by a short-lived peg in the mid-1990s and a few currency crashes (e.g. in 1998).

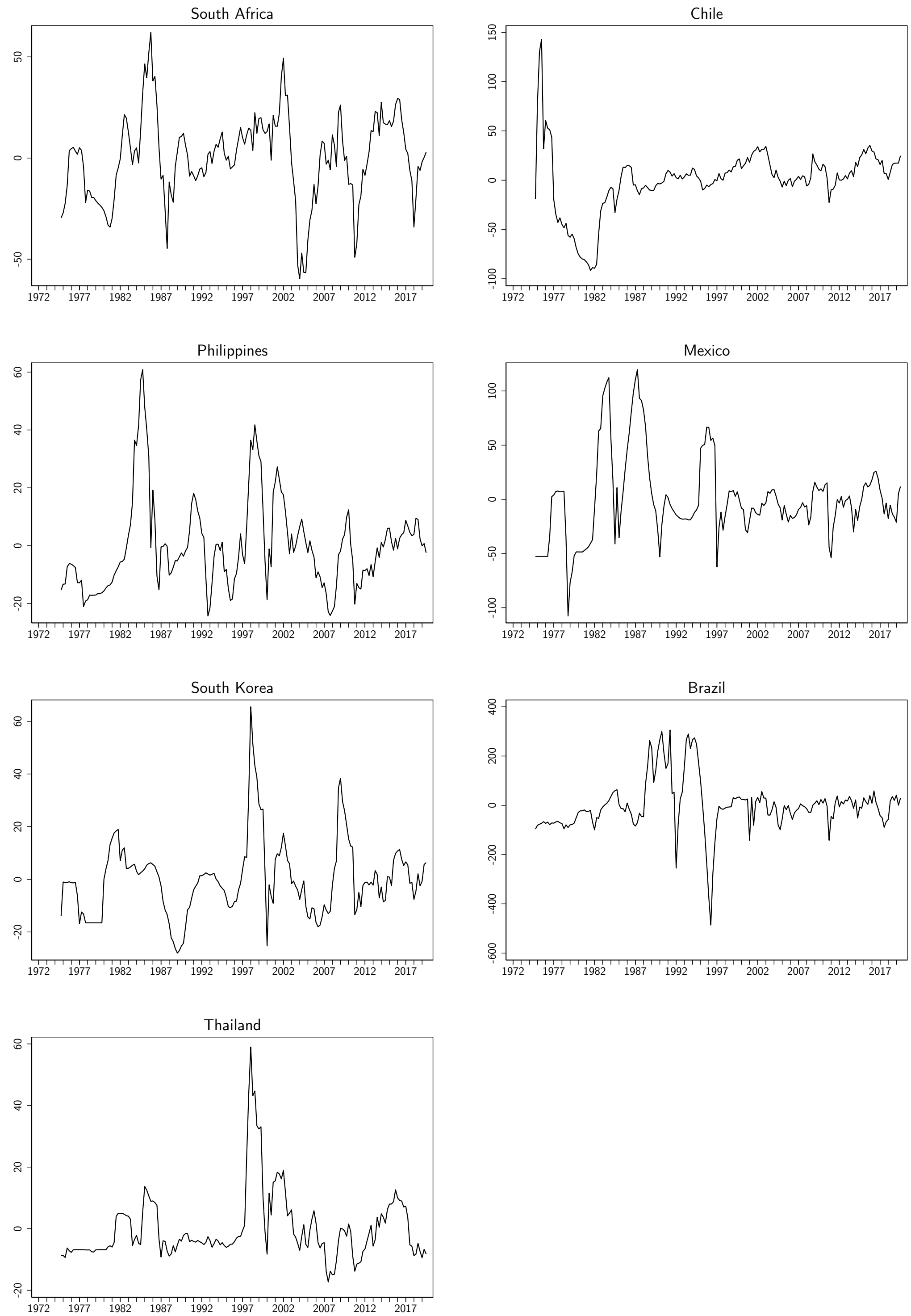
[15] Thailand had a peg up until the East Asian crisis in 1998, and Brazil had several chaotic episodes of hyperinflation and currency crises up until the end of the 20th century.

[16] Nominal exchange rates are strongly correlated with real exchange rates (Cordella and Gupta, 2015).

it with a growth rate filter to assess its robustness.[17]

[17] We construct the Hamilton filter as the residual from the regression $x_{t+8} = \beta_0 + \beta_1 x_t + \beta_2 x_{t-1} + \beta_3 x_{t-2} + \beta_4 x_{t-3} + \nu_{t+8}$. Hamilton (2018) argues that unlike the frequently used Hodrick-Prescott filter, his regression filter does not generate spurious dynamics and prevents filtered values at the end of the sample from behaving differently from those in the middle. An alternative approach is to take (annualised) growth rates, which, however, are known to amplify higher frequencies and may remove lower frequencies in the data (Hamilton, 1994: :171). This is especially problematic for exchange rates series which typically exhibit substantial high-frequency fluctuations that are unrelated to the boom-bust cycles we are interested in. For this reason, the Hamilton filter is our preferred one.

Figure 2: Detrended nominal US-dollar exchange rates, 1972Q1 – 2019Q3



Notes Cyclical components were extracted through Hamilton's regression filter (see Appendix **??**) and are expressed in per cent deviation from trend.

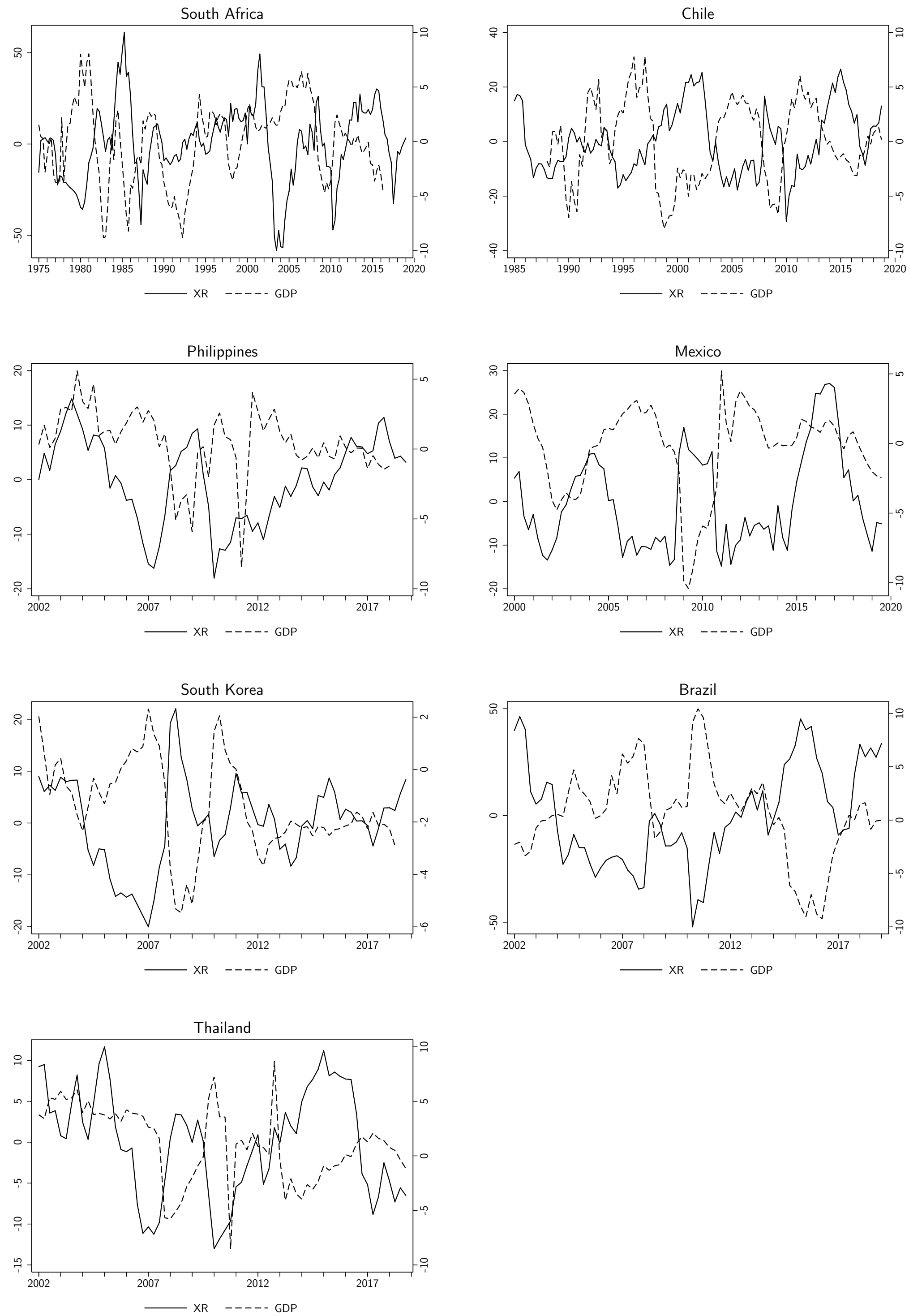
Figure 2 reports the cyclical components in US-dollar exchange rates for the full sample (1972Q1 – 2019Q3). It is noteworthy that for most countries after the end of chaotic episodes, a period of floating exchange rates began that is characterised by much more regular exchange rate fluctuations. A case in point is Chile, where the 1982 crisis appears to have introduced a new regime of fairly periodic exchange rate cycles. Similar patterns can be observed for Mexico (after 1994) and South Korea (after 1998). Brazil (after 2001) and Thailand (after 1998) also display signs of this pattern, but the fluctuations are less regular. South Africa and the Philippines display fluctuations over the whole sample period, those in South Africa being more regular.

Given our interest in the cyclical role of flexible exchange rates, we zero in on those post-crisis periods to assess how the more regular fluctuations in exchange rates relate to cycles in economic activity.[18] Figure 3 therefore reports cyclical components in *XR* along with (logged) real *GDP*, where the sample start was set so as to exclude major currency crises episodes and fixed exchange rate regimes.[19] Cycles are most pronounced in South Africa and Chile, which seem to have a frequency in the range of 8 to 10 years. Exchange rate cycles are also visible in the Philippines, Mexico, and South Korea, albeit a bit more erratic and with a shorter frequency. By contrast, it is more difficult to identify regular cycles in Brazil and Thailand, which display largely idiosyncratic fluctuations. There is also evidence of joint cyclical behaviour in *XR* and *GDP*, often in form of a negative co-movement, which is especially strong in Chile, South Korea, and Brazil.

[18] Figure A2 in Appendix C compares the Hamilton-filtered *XR* series with the growth rate filter. It can be seen that the results are qualitatively very similar.

[19] To exclude fixed exchange rate regimes and crises episodes, we relied on the coarse classification in Ilzetzki et al. (2019) (scores 1 and 5; see notes to Figure A1). In some cases, a few additional data points in the vicinity of crises were excluded if the series still exhibited extreme values. The restricted sample starts are as follows: South Africa: 1972Q4, Brazil: 1999Q4, Chile: 1983Q1, Mexico: 1997Q2, South Korea: 2000Q1, Philippines: 2000Q1, Thailand: 2000Q1.

Figure 3: Cycles in US-dollar exchange rates (left scale) and GDP (right scale)

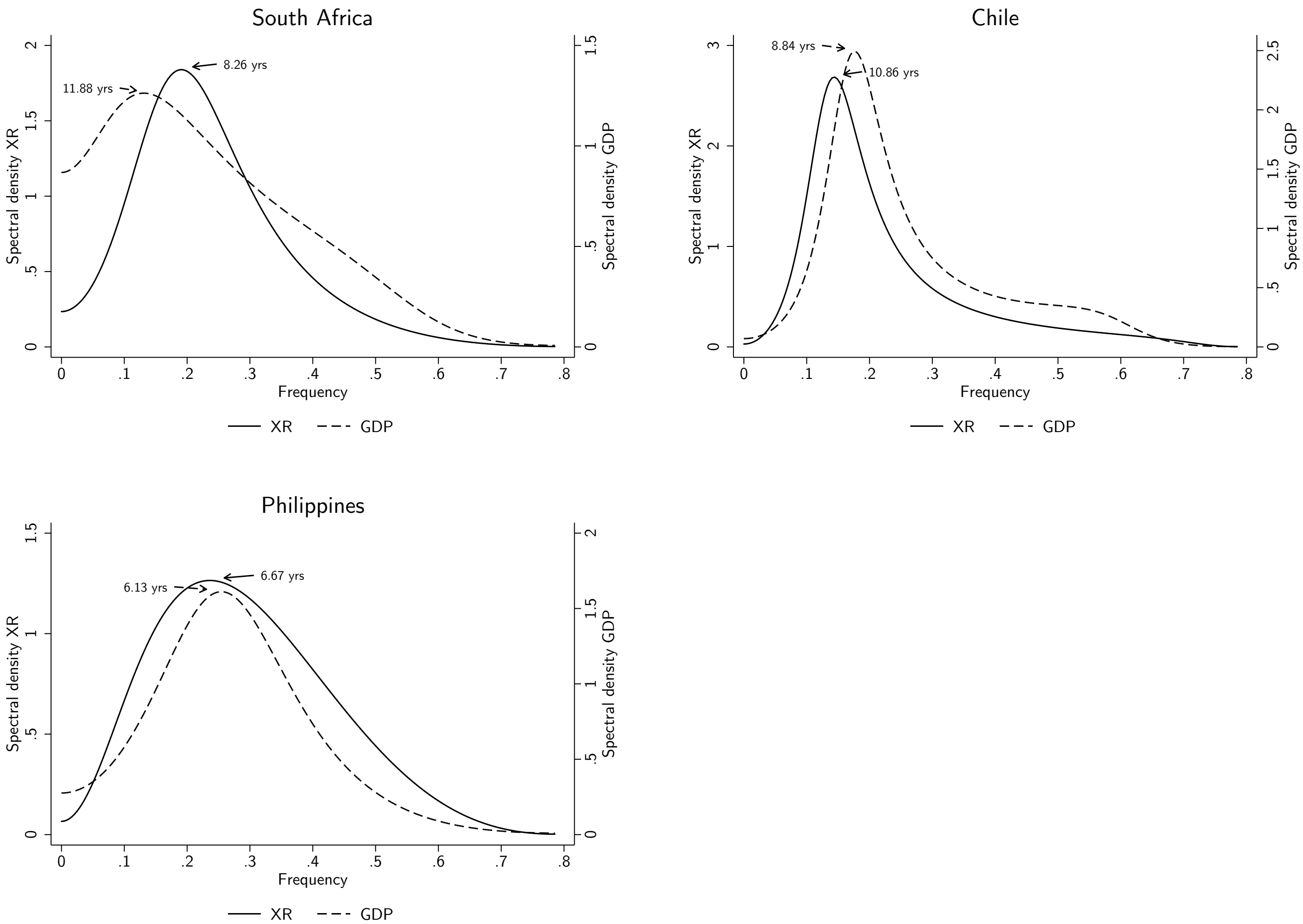


Notes: XR: logged nominal US-dollar exchange rate (cyclical component); *GDP*: (seasonally adjusted) logged real GDP (cyclical component). Cyclical components were extracted through Hamilton's regression filter (see Appendix **??**) and are expressed in per cent deviation from trend.

To examine the periodicities in XR and GDP more rigorously, we estimate spectral density functions (Hamilton, 1994: Ch. 6). A spectral density function describes how much of the total variance of the series is due to different frequencies. Isolated peaks in a spectral density function indicate dominant cycles that are periodic and pinpoint their length. The more the spectral density function is concentrated around a modal value, the more regular the cycle length indicated by that peak. Importantly, if a spectral density function does not exhibit distinct peaks, the series is mostly driven by irregular components. Spectral density functions thereby allow to asses whether fluctuations in a time series have a periodic character, which points to endogenous cycles mechanism of the kind discussed in the previous section. Parametric spectral density estimation has been used to study financial cycles in credit and house prices of AEs (Strohsal et al., 2019), but has not been applied to exchange rates in DEEs.[20] The methodology for the parametric estimation of spectral density functions is explained in Appendix D. The underlying estimated ARMA models are reported in Appendix F.

[20]The main advantage of parametric estimation over non-parametric approaches is its efficiency as it requires fewer degrees of freedom.

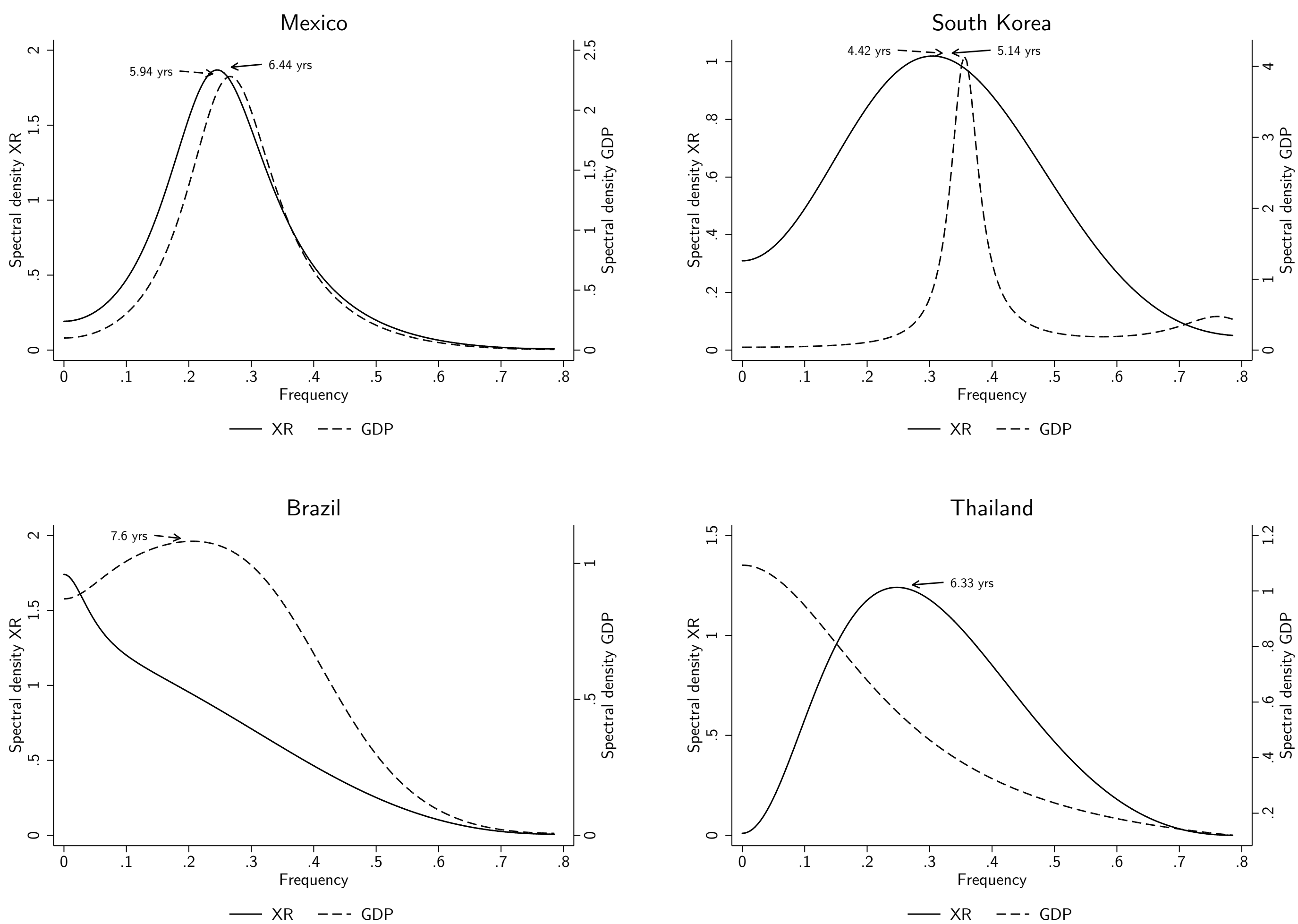
Figure 4: Spectral densities of US-dollar exchange rates and GDP: South Africa, Chile, the Philippines



Notes: *XR*: logged nominal US-dollar exchange rate (cyclical components); *GDP*: logged real GDP (cyclical components). Cyclical components were extracted through Hamilton's regression filter. Spectral densities were estimated parametrically from ARMA models (see Appendix F). Arrows indicate the cycle length (in years) associated with the peak in the spectral densities.

Figures 4-5 displays univariate spectral densities for *XR* and *GDP*. All countries, except for Brazil, exhibit a dominant cycle frequency in *XR*. Estimated cycle lengths range from 4 1/2 years (South Korea) to almost 11 years (Chile). Estimated cycle frequencies in *GDP* are in a similar range; from around 5 years in South Korea up to almost 12 years in South Africa. In several countries, the dominant frequency in *XR* closely corresponds to that of *GDP*, notably in Chile, the Philippines, Mexico, and South Korea. Only Thailand does not exhibit a dominant periodicity in *GDP*. Exchange rate cycles are particularly pronounced in South Africa, Chile, and Mexico, whose spectral density functions are strongly centred on a dominant peak. Periodicities appear to be less pronounced in the Asian countries, where spectral densities are more dispersed around the peak. A potential reason for this is the shorter sample period. Appendix E also reports estimated spectral densities for *XR* and *GDP* with the growth rate rather than Hamilton's filter. The results are similar, but the estimated cycle lengths tend to be shorter (by around one to two years). Estimated cycle lengths thus come with some uncertainty related to the detrending method. Importantly, the qualitative presence of regular periodicities is not sensitive to the detrending method.[21]

Figure 5: Spectral densities of US-dollar exchange rates and GDP: Mexico, South Korea, Brazil, Thailand



Notes: XR: logged nominal US-dollar exchange rate (cyclical components); *GDP*: logged real GDP (cyclical components). Cyclical components were extracted through Hamilton's regression filter. Spectral densities were estimated parametrically from ARMA models (see Appendix F). Arrows indicate the cycle length (in years) associated with the peak in the spectral densities.

[21] The only exception is Thailand's GDP, for which the Hamilton-filtered series does not display a dominant cycle length, whereas the growth rate-filtered series displays a short cycle of 2.5 years.

Overall, the descriptive evidence suggests that several major DEEs exhibit periodic exchange rate cycles since they embarked on floating exchange rate regimes. Estimated cycle lengths in exchange rates range between 4 1/2 to 11 years. These cycles are often closely aligned with cycles in GDP, especially in South Africa, Chile, the Philippines Mexico, and South Korea. Prima facie, this is consistent with the theoretical notion of an interaction mechanism between exchange rates and economic activity that endogenously drives joint cycles.

4 Estimating exchange rate cycle mechanisms: econometric method and data

To asses whether a cyclical interaction mechanism as postulated by the stylized model in section 2 is present in our sample of DEEs, we build on the econometric approach developed in Stockhammer et al. (2019). They use bivariate vector-autoregressions (VARs) with a financial variable (e.g. corporate debt) and GDP in order to detect the presence of cyclical interactions for a dataset of AEs. We extend their approach to the interaction of exchange rates and economic activity, while also controlling for various external factors that may impact macroeconomic dynamics in DEEs.

First, consider a linear version of the model in (6)-(7),

$$\mathbf{y_t} = \alpha + A\mathbf{y_{t-1}} + B\mathbf{z_{t-1}}, \tag{9}$$

where the endogenous variables are collected in $\mathbf{y_t} = [Y_t, s_t]'$ and the exogenous variables in $\mathbf{z_t} = [Y_t^{d*}, i_t^*, \rho_t]'$. The coefficient matrix A corresponds to the (linearised) Jacobian matrix in (8) and B contains the coefficients on the exogenous variables. These variables can be a source of shocks or shifts and thereby drive some of the dynamics in the endogenous variables. However, cyclical dynamics are the outcome of the interaction mechanism between exchange rates and output, given by the product on the off-diagonal elements in A: $a_{12}a_{21} < 0$. It is then this interaction mechanism that transforms shocks into regular fluctuations.[22]

For the empirical analysis, the linear model in (9) can be readily transformed into a bivariate reduced-form VAR with lag length p, augmented by a vector of exogenous variables $\mathbf{z_t}$ with

[22] An alternative source of cycles is that one or more of the exogenous variables in $\mathbf{z_{t-1}}$ are cyclical themselves. We explore this possibility in section 6 below.

lag length h. This yields the following VARX(p, h):[23]

$$\mathbf{y_t} = \alpha + \sum_{i=1}^{p} A_i \mathbf{y_{t-i}} + \sum_{j=0}^{h} B_j \mathbf{z_{t-h}} + \epsilon_\mathbf{t}. \tag{10}$$

While the model in (9) only has first-order lags, a VAR estimated on empirical time series will likely require higher-order lags to render the error terms well-behaved. Thus, the lag length p in (10) is chosen so as to remove serial correlation. Stockhammer et al. (2019) show that in the case of a VAR(p), the coefficients that make up the interaction mechanism are still identified and can be obtained from the off-diagonal of the reduced-form coefficient matrix A_1. Note that no further identification assumptions are required as the focus is only on those two coefficients of the reduced-form coefficient matrix (as opposed to, for example, contemporaneous effects that would be needed to identify impulse response functions).

To estimate the VARX in (10) with real GDP and nominal US-dollar exchange rates, we use annual data. Stockhammer et al. (2019) argue that annual data are more suitable for estimating the interaction mechanism on the first-order lags of the system, as VARs with quarterly data typically require a larger number of lags, which exacerbate multicollinearity problems, may introduce irrelevant high-frequency fluctuations, and overall make it difficult to attribute cyclical dynamics to the coefficients on the first-order lags. The results in section 3 have shown that cycles are at frequencies of 4 to 12 years, suggesting that annual data are suitable to pick them up. To determine the appropriate lag length p for the endogenous variables, we start with a minimum lag length of two. We then check for serial correlation in the residuals and successively increase the number of lags up to six until all serial correlation is removed. Mindful of the relatively small sample size and the fact that data are at annual frequency, we set the lag structure of the exogenous variables to $j = h = 1$, aiming for parsimony.[24]

In order to maximise degrees of freedom, we use the full time span and set the sample start to 1972. As a result, crises episodes reported in Figure A1 will be included in the estimations. We deal with this problem in two ways: first, by augmenting the bivariate VARs with external variables which are treated as exogenous, thus obtaining the VARX in (10). Insofar as crises were triggered by shocks to those global variables, the VARX will control for those episodes. A further advantage of adding external variables is that it enables us to assess whether the cyclical relationship between exchange rates and economic activity is still significant when

[23]See Lütkepohl (2005: Ch. 10) for a general treatment of VARXs.
[24]We check the robustness of our baseline specification to the case where $j = h = 0$.

controlling for external drivers. Second, we also report a bivariate VAR specification with step indicator saturation (SIS) to capture crises events (Castle et al., 2015). Step indicators capture outliers *and* mean shifts and thus mitigate both heteroskedasticity and structural breaks related to crises. Finally, note that while the descriptive analysis in section 3 was based on detrended data, the VARX is estimated on unfiltered data in (log-)levels.[25]

Three external factors will be considered in line with the theoretical model in section 2. First, to capture movements in international commodity prices that may affect export performance and exchange rates (IMF, 2012), we use a country-specific (logged) commodity export price index (CMP^W) provided by Gruss and Kebjah (2019) that weights global commodity prices by the share of each commodity in the total commodity exports of the respective country.[26] Second, we consider the real US monetary policy rate, defined as the Federal Funds rate minus the (annualised) US CPI inflation rate ($FFUND$).[27] Third, the (logged) VXO, a precursor to the VIX, which measures implied volatility in the S&P100 and serves as a measure for risk aversion by global investors.[28]

Controlling for external variables as well as step indicators may not fully address the problem of structural breaks created by crises episodes. We therefore expect point estimates of those countries that underwent major crisis episodes during the sample period (e.g. Brazil) to be less reliable compared to those that had fewer or no crises (e.g. South Africa, Chile). Similarly, countries that had fixed or semi-fixed exchange rate regimes throughout most of the sample period, such as Thailand before the East Asian crisis, are less likely to exhibit a cycle mechanism over the full sample period.

[25] This is common when it is unclear whether the relevant variables contain a unit root and/or are cointegrated. As Kilian and Lütkepohl (2017: Ch. 3) point out, there is an asymmetry between incorrectly imposing a unit root (and then overdifferencing the data) and failing to impose a unit root when there is one. While the former renders the VAR-estimator inconsistent under standard assumptions, the latter approach preserves consistency and may only come with a loss in efficiency. The VAR(X) in levels can be consistently estimated with asymptotically normal standard errors even if some variables are I(1) because the presence of lags would allow the I(1) variables to be re-written as coefficients on differenced and thus I(0) variables (Sims et al., 1990).

[26] The index covers international prices of 45 commodities that are deflated by a unit value index for manufactured exports. See Appendix A for further information.

[27] The US policy rate is a common measure for spillover effects from US monetary policy (Bruno and Shin, 2015; Kalemli-Özcan, 2019).

[28] The VXO is similar to the VIX but uses a smaller set of stock prices. It starts in 1986, whereas the VIX starts in 1990. The VXO and VIX are highly correlated (0.99). The VXO/VIX have become standard proxies of the global financial cycle (Avdjiev et al., 2019; Bruno and Shin, 2015; Cunha et al., 2020; Kalemli-Özcan, 2019; Obstfeld et al., 2019; Rey, 2015).

5 Estimation results

Tables 1-3 report the regression results for the three country groups. In each Table, we start with a baseline specification with CMP^W as the main external control variable (we show in section 6 below that CMP^W is the most relevant of the three external driver for DEEs' exchange rates). The second and third specifications use instead $FFUND$ and VXO, respectively. Note that with VXO, the sample start is reduced to 1987 and the results have to be taken with some caution due to lower degrees of freedom. In Table 1, it can be seen that in the baseline specification with CMP^W, the condition for a cycle mechanism ($a_{12}a_{21} < 0$) is satisfied for South Africa and Chile (with both coefficients statistically significant),[29] as well as for the Philippines (with only one of the two coefficients statistically significant). The signs correspond to the financial channel of exchange rates where currency depreciations are contractionary ($a_{12} < 0$) and the external adjustment channel where output expansions lead to downward pressure on currencies ($a_{21} > 0$). The results are generally robust to the use of different control variables,[30] except for a loss in statistical significance of the effect of GDP on XR for South Africa in the specification with the VXO, which is likely due to the small sample size. Overall, there is robust evidence for a cycle mechanism for these three countries, which are also those for which visual evidence from the detrended exchange rate series in Figures 2 and 3 was most suggestive of relatively stable cycles.[31]

[29] In the VARX with CMP^W for Chile, the null hypothesis of no serial correlation was rejected at the 10% level on the first lag of the error. This did not vanish with the inclusion of up to 6 lags. When adding $FFUND$ to the model (see Table A8 in Appendix G) serial correlation vanishes and the results are very similar.

[30] Note that the VAR for the Philippines with $FFUND$ exhibits serial correlation on the first lag of the error. This vanishes when using the contemporaneous value of $FFUND$, while the results are highly similar.

[31] Recall that for Chile, this only begins at around 1983 after chaotic episodes at the sample start.

Table 1: VARX with GDP, XR, and external factors: South Africa, Chile, and Philippines

	ZAF	ZAF	ZAF	CHL	CHL	CHL	PHL	PHL	PHL
GDP									
L.GDP	1.195***	1.208***	1.377***	1.018***	0.885***	0.610***	1.198***	1.160***	1.166***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.007)	(0.000)	(0.000)	(0.000)
L.XR	-0.073***	-0.074***	-0.044**	-0.065**	-0.059**	-0.145**	-0.142***	-0.127***	-0.034
	(0.002)	(0.001)	(0.031)	(0.023)	(0.026)	(0.037)	(0.006)	(0.005)	(0.478)
L.CMP^W	0.005			0.010			0.017		
	(0.685)			(0.674)			(0.226)		
L.FFUND		-0.001			-0.010***			-0.006***	
		(0.657)			(0.006)			(0.000)	
L.VXO			-0.014*			-0.020			-0.007
			(0.071)			(0.108)			(0.561)
XR									
L.GDP	1.940**	1.362*	1.406	0.931*	1.486***	1.701***	0.494	0.808	1.642**
	(0.017)	(0.084)	(0.195)	(0.091)	(0.006)	(0.004)	(0.309)	(0.177)	(0.024)
L.XR	1.075***	1.138***	1.197***	1.941***	1.925***	1.630***	1.221***	1.255***	1.116***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
L.CMP^W	-0.211**			0.022			-0.125***		
	(0.010)			(0.797)			(0.007)		
L.FFUND		0.015**			0.035**			0.020***	
		(0.048)			(0.010)			(0.002)	
L.VXO			-0.047			0.035			0.095**
			(0.478)			(0.296)			(0.029)
Lags	2	2	2	3	3	2	2	6	2
$a_{12}a_{21} < 0$	YES	YES	YES	YES	YES	YES	YES	YES	YES
CL 1	6.689	6.913	8.271	8.744	10.587	19.821	9.620	9.603	6.540
CL 2				3.939	4.043	2		3.266	
CL 3								2.574	

Notes: Sample period: 1972-2017 (with VXO: 1987-2017). p-values in parentheses. GDP: logged real GDP; XR: logged nominal US-dollar exchange rate; CMP^W: logged commodity export price index; $FFUND$: real federal funds rate; VXO: logged volatility index. A constant term was included in each equation (not reported). Only the coefficients on the first lags are reported. CL: cycle lengths (in years) implied by the complex eigenvalues computed as $\frac{\pi}{2\arccos(\frac{re}{mod})}$, where re is the real part of the eigenvalue and mod is the modulus.

Table 2 presents results for Mexico and South Korea, which had more mixed exchange rate regimes over the sample period compared to the first group. The condition for a cycle mechanism is mostly satisfied and in line with the theoretical framework (except for Mexico with the VXO and Korea with $FFUND$), but the signs are only partially statistically significant.

Table 2: VARX with GDP, XR, and external factors: Mexico and South Korea

	MEX	MEX	MEX	KOR	KOR	KOR
GDP						
L.GDP	0.921***	0.839***	0.806***	0.837***	0.894***	0.549**
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.031)
L.XR	-0.028	-0.050	0.074**	-0.002	0.019	-0.007
	(0.353)	(0.162)	(0.048)	(0.975)	(0.792)	(0.930)
L.CMP^{W}	0.011			-0.016		
	(0.472)			(0.159)		
L.FFUND		-0.005**			0.001	
		(0.019)			(0.529)	
L.VXO			-0.029**			-0.031*
			(0.039)			(0.091)
XR						
L.GDP	1.684	2.129*	1.268	1.097*	0.687	2.327***
	(0.138)	(0.061)	(0.150)	(0.087)	(0.263)	(0.006)
L.XR	1.538***	1.679***	1.021***	1.217***	1.055***	1.199***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
L.CMP^{W}	-0.179*			0.006		
	(0.050)			(0.860)		
L.FFUND		0.043***			0.012*	
		(0.001)			(0.061)	
L.VXO			0.012			0.087
			(0.858)			(0.154)
Lags	2	3	2	3	3	2
$a_{12}a_{21} < 0$	YES	YES	NO	YES	NO	YES
CL 1	15.535	7.980	12.955	8.744	10.587	19.821
CL 2		3.352	2	3.939	4.043	2

Notes: Sample period: 1972-2017 (with VXO: 1987-2017). p-values in parentheses. GDP: logged real GDP; XR: logged nominal US-dollar exchange rate; CMP^{W}: logged commodity export price index; $FFUND$: real federal funds rate; VXO: logged volatility index. A constant term was included in each equation (not reported). Only the coefficients on the first lags are reported. CL: cycle lengths (in years) implied by the complex eigenvalues computed as $\frac{\pi}{2\arccos(\frac{re}{mod})}$, where re is the real part of the eigenvalue and mod is the modulus.

Table 3 contains the results for Brazil and Thailand, which exhibited mostly chaotic or fixed exchange rates during the sample, rendering the presence of a cycle mechanism less likely. Indeed, the signs on the relevant coefficients tend to be statistically insignificant and mostly do not meet the condition for a cycle mechanism. Only the specification with the VXO for Thailand yields estimates in line with the hypothesised cycle mechanism, but the fact that the signs are reversed in the specification with $FFUND$ points to a lack of robustness.

Table 3: VARX with GDP, XR, and external factors: Brazil and Thailand

	BRA	BRA	BRA	THA	THA	THA
GDP						
L.GDP	1.010***	1.091***	1.058***	1.270***	1.439***	1.174***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
L.XR	0.002	-0.004	0.023***	-0.074	0.033	-0.057
	(0.739)	(0.507)	(0.007)	(0.485)	(0.728)	(0.614)
L.CMP^{W}	0.041*			-0.028		
	(0.069)			(0.139)		
L.FFUND		-0.002			-0.001	
		(0.295)			(0.706)	
L.VXO			-0.012			-0.008
			(0.433)			(0.709)
XR						
L.GDP	1.989	-0.320	3.033	-0.427	-0.227	0.055
	(0.400)	(0.893)	(0.350)	(0.287)	(0.522)	(0.922)
L.XR	1.635***	1.777***	1.410***	0.898***	0.988***	1.140***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
L.CMP^{W}	-0.921***			-0.060		
	(0.005)			(0.129)		
L.FFUND		0.028			0.005	
		(0.375)			(0.190)	
L.VXO			0.009			0.043
			(0.976)			(0.349)
Lags	2	2	6	2	2	2
$a_{12}a_{21} < 0$	NO	NO	NO	NO	YES	YES
CL 1	25.763	33.964	22.359	24.686	29.456	25.942
CL 2			8.820	2		
CL 3			4.073			
CL 4			2.956			

Notes: Sample period: 1972-2017 (with VXO: 1987-2017). p-values in parentheses. GDP: logged real GDP; XR: logged nominal US-dollar exchange rate; CMP^{W}: logged commodity export price index; $FFUND$: real federal funds rate; VXO: logged volatility index. A constant term was included in each equation (not reported). Only the coefficients on the first lags are reported. CL: cycle lengths (in years) implied by the complex eigenvalues computed as $\frac{\pi}{2\arccos(\frac{re}{mod})}$, where re is the real part of the eigenvalue and mod is the modulus.

Two further insights can be gained from Tables 1-3. First, the external factors tend to exhibit the expected signs, insofar they are statistically significant. An improvement in CMP^{W} is associated with an appreciation of the domestic currency for most countries; an increase in $FFUND$ is associated with a depreciation of domestic currencies against the US dollar and economic contraction; and VXO shocks have contractionary effects and lead to currency depreciation for most countries. Second, the Tables also report estimated cycle lengths (CL) that are implied by the coefficient matrices of the VARX.[32] Focusing on the baseline with CMP^W, we find cycle frequencies of around 6 1/2 and 8 1/2 years for South Africa and Chile that are similar to the estimated frequency in exchange rates in section 3, but slightly shorter. For the Philippines, the estimated cycle length is around 9 1/2, which is a bit longer compared to section 3; probably because of the longer sample span. For the other countries, estimated frequencies are often substantially longer, which is likely to be due to the presence of currency crises episodes that are not captured by the control variables.

As a final exercise, we obtain additional results from bivariate VARs with endogenously selected step indicators, which are dummy variables that are equal to unity from a specific break year onwards and zero otherwise (see Table A5 in Appendix G).[33] Step indicators absorb unexplained mean shifts in the exchange rate that are due to currency crises and hyperinflation episodes that are unrelated to the cycle mechanism, but may affect the results. The selected step indicators indeed capture many of the crises and changes in exchange rate regimes documented in Ilzetzki et al. (2019) and reported in Figure A1, such as the 1998 East Asian crisis in Korea and Thailand.[34] They also pick up other external events, e.g. strong depreciations in the Philippines and Thailand during the Great Recession of 2009.

The cycle condition is again satisfied in South Africa, Chile, and the Philippines, and is statistically significant for all three countries. By contrast, Mexico, South Korea, Brazil, and Thailand, which either underwent numerous crises episodes or substantial exchange regime shifts throughout the sample period, display no evidence for a stable cycle mechanism even when controlling for these shifts through step indicators. Estimated cycle lengths range from 3 years (South Korea) to almost 8 years (Chile), and are generally are close (up to around 1 year) to the estimated frequencies displayed in Figures 4-5 (except for Mexico whose length of around 10 years is substantially longer).

Further robustness tests are reported in Appendix G. The main results hold up. In a VAR

[32] From the polar representation of the complex eigenvalues $\lambda = |\lambda|(\cos\theta \pm i.\sin\theta)$ of the companion matrix, the implied cycle length is given by $CL = \frac{2\pi}{\theta} = \frac{2\pi}{\arccos(Re/|\lambda|)}$. Note that these complex eigenvalues cannot be directly mapped to the interaction mechanism in A_1 as they may also stem from the coefficients on the higher order terms A_{i+1}, $i = 1, ...p-1$.

[33] The selection of step indicators is based on the split half approach (Castle et al., 2015): first, create step indicators for the entire sample period; then estimate the model on the full sample, first with only the first half of step indicators, and then with the second half. Retain those step indicators from both estimations whose p-value is equal or below $1/T$ and re-estimate the model with only those step indicators. Lastly, exclude step indicators whose p-value exceeds $1/T$. As we are interested in controlling for exogenous shifts in the XR series, we select those step indicators that are statistically significant in the XR-equation and insert them in both equations.

[34] It is noteworthy that the Philippines is the only one of the three East Asian countries for which no step indicator is retained that would capture the 1998 crisis. This confirms visual evidence in Figures A1 and 2 that 1998 did not involve a structural break for the Philippines.

specification without exogenous variables, South Africa, Chile, the Philippines, and Mexico, meet the condition for a cycle mechanism between exchange rates and GDP. Results from a VARX with CMP^W entering contemporaneously rather than lagged are qualitatively identical with respect to the cycle mechanism. Finally, in a VARX with both CMP^W and $FFUND$, the cycle condition holds for all countries except Brazil and Thailand, with both coefficients statistically significant in South Africa and Chile, and partially significant coefficients in the Philippines and Mexico.

In summary, we find robust evidence for the presence of a cyclical interaction mechanism between exchange rates and output in South Africa, Chile, and the Philippines (the latter only partly significant). Results for Mexico and South Korea are mixed and not always robust. For Brazil and Thailand, the cycle condition is never satisfied or significant. These results are consistent with our hypothesis that the endogenous cycle mechanism is more likely to operate in floating exchange rate regimes. Indeed, countries with relatively stable regimes of flexible or semi-flexible exchange rate during the post-Bretton Woods period (South Africa, Chile, the Philippines) exhibit the strongest evidence for a cycle mechanism. By contrast, countries that underwent multiple crises episodes and/or shifts in the exchange rate regime at a relatively late stage of the sample period (Mexico, South Korea) do not exhibit robust evidence of a stable interaction mechanism. Brazil was particularly heavily affected by numerous chaotic episodes, which may explain the complete absence of a stable interaction mechanism. Thailand had a pegged exchange rate throughout most of the sample period, which may have prevented the emergence of a cycle mechanism.

6 How important are external cycle drivers?

Our estimations show that cyclical interaction mechanisms between exchange rates and output are present in several DEEs even when controlling for potential external cycle drivers, such as commodity prices, US monetary policy, and risk perceptions of international investors. However, what is the relative importance of these external factors for exchange rate cycles in DEEs? This section summarises results from a number of additional empirical exercise that are documented in Appendix H.

First, we assess whether there is strong co-movement in exchange rates and GDP across countries, which would indicate an important role for external drivers. To this end, we examine the co-movement of (detrended) nominal US-dollar exchange rates and real GDP across our seven DEEs over the period 2002Q4–2019Q3 through correlation as well as principal component analysis. For comparison, we do the same exercise for a group of seven small open AEs (see Appendix H.1).[35] We find a moderate average correlation coefficient in exchange rates across DEEs of 0.4. For AEs, the correlation is higher (0.6). Similarly, the first principal component only explains around 53% of the variation in exchange rates in DEEs, but about 72% in AEs. Very similar results are found for GDP. Thus, there is a notable co-movement in exchange rates and economic activity across DEEs in our sample, but there is also a substantial amount of independence. This supports the theoretical notion of endogenous cycle mechanisms that are country-specific and may lead to uneven domestic responses to common external shocks.

Second, we assess which of the three external factors (commodity prices, US monetary policy, global risk aversion) is most closely correlated with the co-movements in exchange rates. To do so, we estimate a dynamic factor model that allows us to extract a common factor in nominal US-dollar exchange rates in our sample of DEEs that could account for any joint co-movement. We then assess potential external determinants of these joint fluctuations in exchange rates by estimating auto-regressive distributed lag models (ARDLs) of the common dynamic factor as a function of external variables (see Appendix H.2 for details on method and estimation results). We find that commodity prices are the closest correlate of the common factor in exchange rates across DEEs.[36] There is also some, but weaker, evidence for an effect of global risk aversion, and no evidence for a major role of US monetary policy.[37] This suggests that our baseline VAR specification with commodity prices successfully controls for one of the major external drivers of exchange rates in DEEs.

Finally, we investigate whether external factors may in fact fully account for the cycle periodicities documented in section 3. If that was the case, the cyclical interaction mechanism that we emphasise would not add anything to the explanation of exchange rate cycles. To assess this possibility, we examine the cyclical properties of the three external factors by estimating spectral density functions of their cyclical components in the same way we did for exchange rates and GDP (see Appendix H.3). The results reveal that the only variable

[35] United Kingdom, France, Norway, Sweden, Canada, Japan, Australia.

[36] In contrast to the VAR estimations, where we used a commodity price index with country-specific weights (CMP^W), here we employ a uniform global (logged) primary commodity price index (denominated in US dollars and deflated by the US consumer price index) (CMP), which is a weighted average based on global import shares and contains 68 commodities covering energy, agricultural products, fertilizers and metals.

[37] More specifically, the US policy rate is never jointly statistically significant, while the VXO is jointly statistically significant only in the specification where it enters as the sole explanatory variable. By contrast, the commodity price index is jointly significant both in the bivariate and multivariate specification. The adjusted R^2 of 90% in the specification with all variables is barely higher than the adjusted R^2 of a specification with the commodity price index only. This suggests that the global commodity price index explains the largest share of the variance of the dynamic factor among the external variables under consideration.

that exhibits a dominant cycle frequency is the *VXO* with an estimated cycle length of 13 years. This is above the estimated frequencies for exchange rates in DEEs, which range from 4 to 11 years.[38] For *FFUND* and CMP^W, no dominant periodicity can be found, which most likely reflects the fact that these series exhibit erratic dynamics or time-varying periodicities. Overall, spectral analysis of global factors does not suggest dominant periodicities that would match the periodicities found in exchange rates and GDP. Global factors alone thus do not seem to account for regular cycles in exchange rates and economic activity in our sample of DEEs, and need to be combined with cyclical interaction mechanisms to explain boom-bust cycles.

7 Conclusion

This article has investigated a pro-cyclical role of flexible exchange rates in developing and emerging market economies. Drawing on post-Keynesian and structuralist frameworks as well as recent work at the Bank for International Settlements, we proposed a simple model in which the procyclicality of exchange rates stems from an endogenous cycle mechanism. Central to this is an interaction mechanism, whereby depreciations are contractionary due to the financial channel of exchange rates and output contractions feed into exchange rate revaluation via an external adjustment channel. In this way, exchange rates become an endogenous driver of cyclical fluctuations along with external shocks. We provided descriptive and econometric evidence for the presence of such a cyclical interaction mechanism, which was especially strong for South Africa and Chile, and to a lesser extent for the Philippines. Mexico and South Korea also display evidence for an interaction mechanism in some specifications, but this is less robust. There is no evidence for Brazil and Thailand. However, our results also show that Mexico and South Korea display regular cycles since they switched to flexible exchange rate regimes in the late 1990s, so that econometric evidence for a cycle mechanism may become more clear-cut in the future.

Three important implications follow from our analysis. First, we provide theoretical and empirical support for the notion of pro-cyclical flexible exchange rates put forward in some of the post-Keynesian and structuralist literature on emerging markets (Harvey, 2010; Kohler, 2019; Ocampo, 2002, 2016; Stiglitz et al., 2006). This literature has mostly emphasised destabilising balance sheet effects from fluctuating exchange rates. We show that when combined with a standard feedback mechanism from economic activity to exchange rates, an

[38]Furthermore, the *VXO*'s spectral density is widely dispersed around the peak, suggesting that the 13-year periodicity in *VXO* is not very pronounced.

endogenous cycle mechanism emerges. Correspondingly, flexible exchange rates in emerging markets may rather be described as 'drivers of endogenous cycles' than as 'shock absorbers', as much of the mainstream discussion has it.

Second, our analysis opens up a new perspective on the debate on macroeconomic instability in emerging markets. While large parts of the recent mainstream and heterodox literature have strongly emphasised global factors (Bonizzi and Kaltenbrunner, 2018; Cunha et al., 2020; IMF, 2012; Kalemli-Özcan, 2019; Ocampo, 2016; Rey, 2015), we argue that they at best account for a part of the story. Co-movements in exchange rates and economic activity across emerging markets are limited. In addition, external factors do not exhibit the periodicities found in exchange rates and GDP, suggesting that country-specific mechanisms do play an important role. In our view, external shocks, e.g. to commodity prices or 'the global financial cycle' (Rey, 2015, 2016), can be important triggers, but only in combination with cyclical interactions mechanisms between output and exchange rates do they generate periodic boom-bust cycle dynamics. An advantage of our perspective is that it directs attention to the destabilising role of domestic variables such as the nominal exchange rate that are, at least to some degree, under the control of policy makers.

Third, our argument implies that the endorsement of fully flexible exchange rates by major international policy institutions should be questioned. Post-Keynesian and structuralist economists have argued for the practice of managed floating, i.e. occasional interventions in foreign exchange markets to smooth exchange rate fluctuations (Frenkel, 2006; Ferrari-Filho and De Paula, 2008; Guzman et al., 2018; Ocampo, 2002). While this will not fully mute pro-cyclical effects of exchange rates, it is likely to reduce some of their destabilising effects without running the risk of speculative attacks on a peg. At the same time, exchange rate management alone should not be regarded as a panacea. In conjunction with liberalised financial accounts, it can fail to curb speculative capital flows or even attract them as investors believe central banks will prevent losses (Kaltenbrunner and Painceira, 2017). Therefore, managed floating should be combined with capital controls that not only discourage speculative short-term capital flows, but also hamper the occurrence of currency mismatches that are at the heart of the pro-cyclical effects of exchange rates highlighted in this article.

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Appendix

A Dataset

Table A1: Data definitions and sources

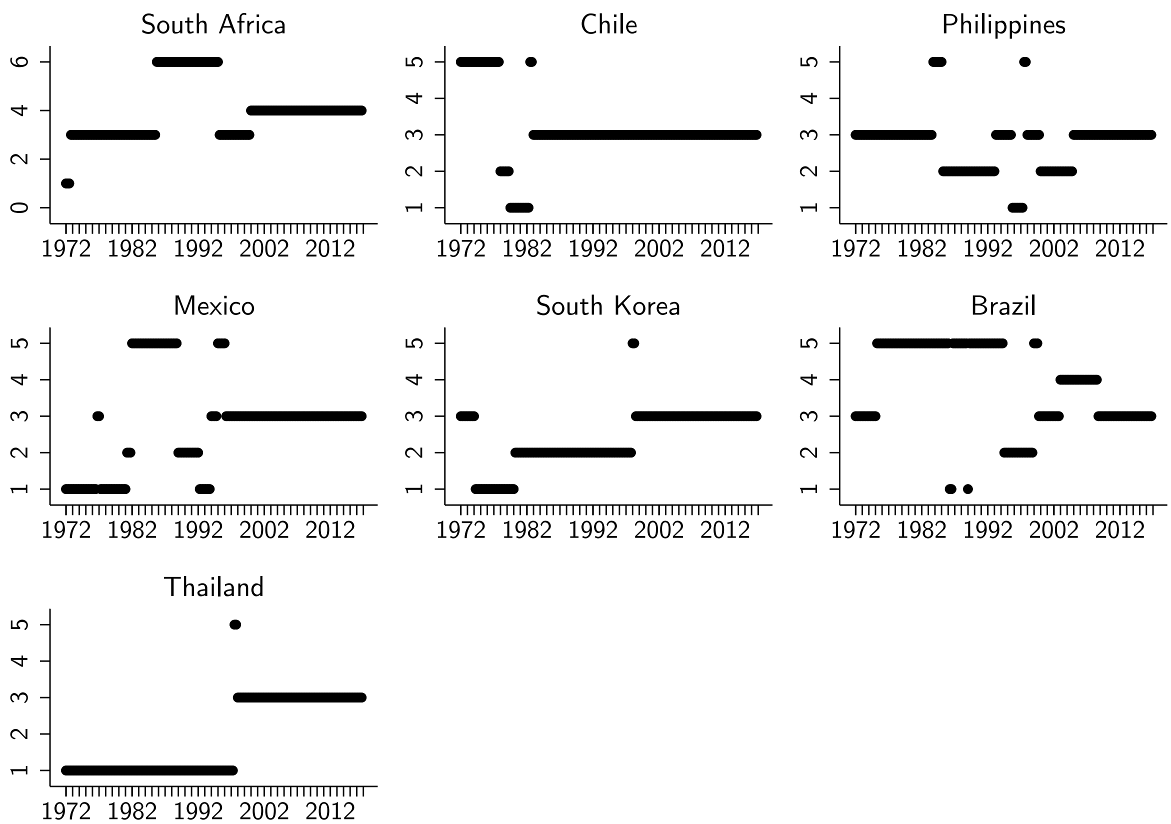
Variable	**Definition**	**Source(s)**	**Sample range & notes**
CMP	Global commodity price index, deflated by US CPI, natural log, average over period	IMF	1991Q1-2019Q3; weighted average over 68 commodities based on global import shares
CMP^{W}	Country-specific global commodity export price index, deflated by manufacturing unit value (MUV), natural log, average over period	Gruss and Kebjah (2019)	1980Q1-2019Q3 (quarterly series); 1972-2017 (annual series); weighted average over 45 commodities based on the share of each commodity in the total commodity exports of the country
$FFUND$	Real effective federal funds rate, constructed as nominal rate minus CPI inflation rate	FRED	1972Q1-2019Q3
GDP	Real gross domestic product, natural log	IMF (IFS), OECD, World Bank (WDI)	Quarterly series are seasonally adjusted. Where adjusted series were not available, seasonal adjustment was performed manually using the X-13 ARIMA SEATS routine of the United States Census Bureau. The routine was accessed through the R-package *seasonal*.
VXO	CBOE S&P 100 Volatility Index (implied volatility of stock options), natural log, average over period	FRED	1986Q1-2019Q3.
XR	Nominal US-dollar exchange rate, natural log, average of period	IMF (IFS)	

Table A2: Country-specific sample range

Country	Quarterly *XR*	Quarterly *GDP*	Quarterly Restricted sample period	Annual *XR*	Annual *GDP*
South Africa	1972Q1-2019Q3	1972Q1-2016Q4	1972Q4-2019Q3	1972-2017	1972-2017
Chile	1972Q1-2019Q3	1986Q1-2019Q3	1983Q1-2019Q3	1972-2017	1972-2017
Philippines	1972Q1-2019Q3	1981Q1-2018Q4	2001Q1-2019Q3	1972-2017	1972-2017
Mexico	1972Q1-2019Q3	1980Q1-2019Q3	1997Q2-2019Q3	1972-2017	1972-2017
South Korea	1972Q1-2019Q3	1972Q1-2019Q1	2001Q1-2019Q3	1972-2017	1972-2017
Brazil	1972Q1-2019Q3	1996Q1-2019Q3	1999Q4-2019Q3	1972-2017	1972-2017
Thailand	1972Q1-2019Q3	1993Q1-2019Q3	2001Q1-2019Q3	1972-2017	1972-2017

B Exchange rate regimes

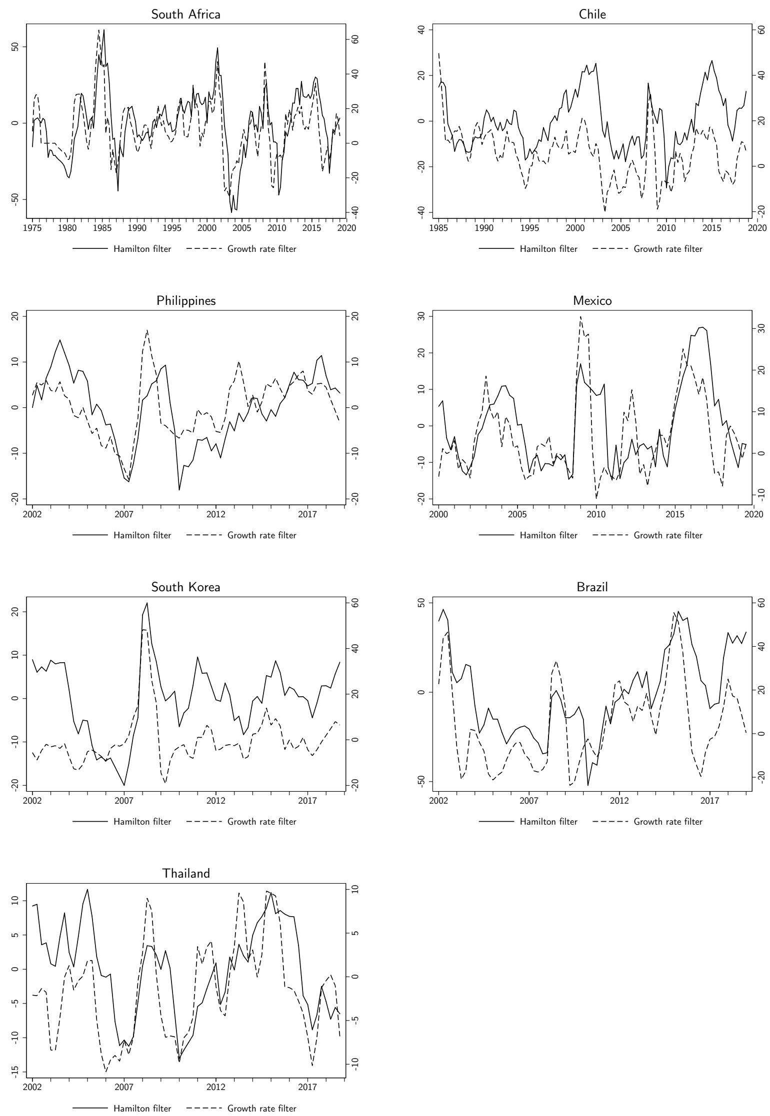
Figure A1: Exchange rate regimes, 1972Q1 – 2016Q4



Data source: Coarse exchange rate regime classification in Ilzetzki et al. (2019).
Notes 1: fixed exchange rate (no separate legal tender; currency board; pre-announced or de facto peg; or pre-announced horizontal band $\leq \pm 2\%$); 2: semi-fixed (pre-announced or de facto crawling band $\leq \pm 2\%$); 3: semi-flexible (pre-announced crawling band $\geq \pm 2\%$; de facto crawling band $\leq \pm 5\%$; moving band $\leq \pm 2\%$; managed floating; 4: flexible; 5: freely falling (inflation $> 40\%$ p.a. and/or currency crash $> 25\%$ p.m. (and 10%-pts greater than that of the previous month)); 6: parallel market with unavailable exchange rate data. Monthly data were converted into quarterly medians.

C Hamilton's filter versus growth-rate filter

Figure A2: Nominal US-dollar exchange rates, cyclical components, Hamilton's filter versus growth-rate filter



Notes Hamilton-filter (left-axis) is the residual from the regression $x_{t+8} = \beta_0 + \beta_1 x_t + \beta_2 x_{t-1} + \beta_3 x_{t-2} + \beta_4 x_{t-3} + \nu_{t+8}$. Growth rate filter (right-axis) is obtained as $\frac{x_t - x_{t-4}}{x_{t-4}}$. Both series are measured in per cent deviation from trend.

D Estimating spectral density functions: methodology

Parametric estimation of the spectral density function of a time series x_t is based on ARMA(p, q) models, which can be written as:

$$\theta(L)x_t = \delta + \phi(L)\epsilon_t, \tag{D.1}$$

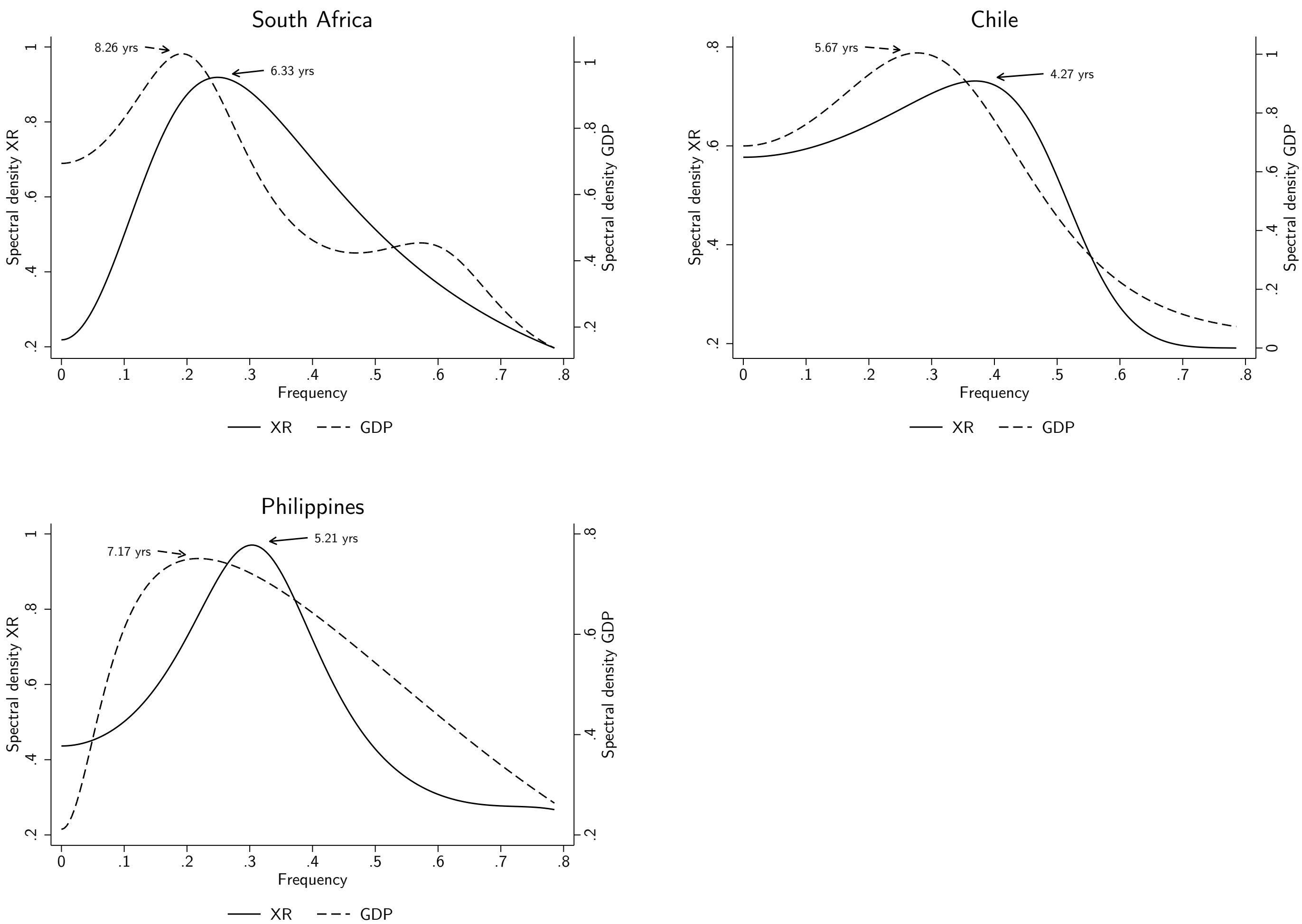
where $\theta(L)$ and $\phi(L)$ are lag polynomials of order p and q, respectively, and ϵ_t is a white noise error term with variance σ_ϵ^2. The spectral density function of x_t can then be obtained as:

$$s_x(\omega) = \frac{\sigma_\epsilon^2}{2\pi}\frac{|\phi(e^{-i\omega})|^2}{|\theta(e^{-i\omega})|^2}, \tag{D.2}$$

where $\omega \in [0, \pi]$ denotes the frequency and i is the imaginary number $i^2 = -1$. We estimate individual ARMA models for each time series, starting from a lag length of 10 and then pare them down by successively, excluding statistically insignificant terms provided that error terms remain serially uncorrelated.

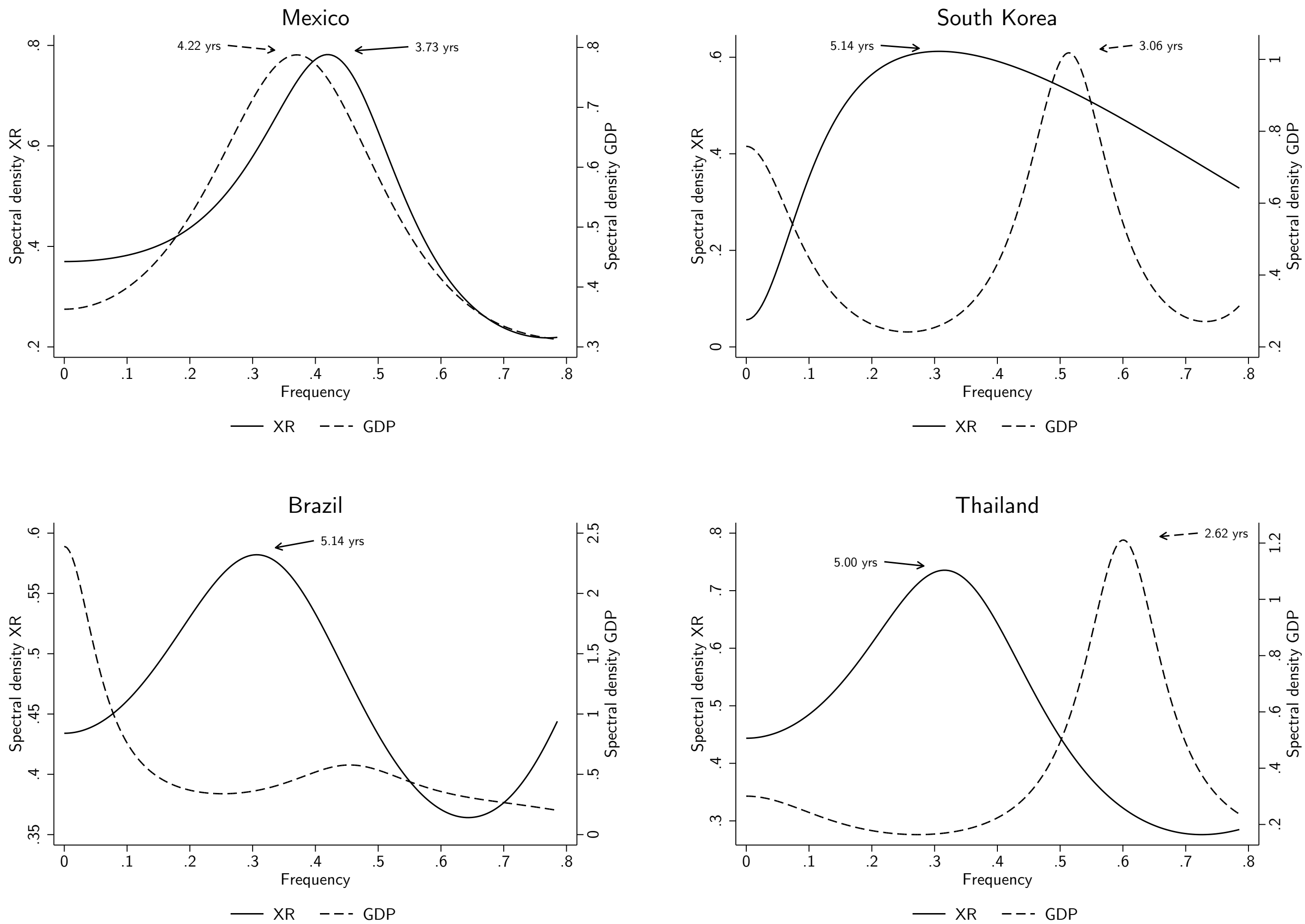
E Spectral densities of growth-rate filtered series

Figure A3: Spectral densities of nominal US-dollar exchange rates and real GDP, growth-rate filtered: South Africa, Chile, the Philippines



Notes: XR: logged nominal US-dollar exchange rate, growth-rate filtered. Spectral densities were estimated parametrically from ARMA models. The sample start was set to the sample start of the Hamilton-filtered series. For Chile, the first four observations (1985Q4-1986Q3) of *XR* were dropped due to extreme values during this period.

Figure A4: Spectral densities of nominal US-dollar exchange rates and real GDP, growth-rate filtered: Mexico, South Korea, Brazil, Thailand



Notes: XR: logged nominal US-dollar exchange rate, growth-rate filtered. Spectral densities were estimated parametrically from ARMA models. The sample start was set to the sample start of the Hamilton-filtered series.

F Estimated ARMA models for spectral density functions

Table A3: ARMA of nominal US-dollar exchanges (*XR*) (cyclical component)

	ZAF	CHL	PHL	MEX	KOR	BRA	THA
AR(1)	0.845	1.129	1.168	0.766	1.117	0.951	1.097
	(15.860)	(16.070)	(11.974)	(9.206)	(6.385)	(28.336)	(7.649)
AR(2)		-0.231	-0.278		-0.324		-0.224
		(-3.182)	(-2.414)		(-1.777)		(-1.422)
AR(3)	0.344			0.216			
	(4.165)			(2.286)			
AR(4)	-0.204						
	(-2.377)						
AR(6)	-0.104			-0.208			
	(-1.508)			(-2.863)			
AR(7)		0.219					
		(2.849)					
AR(8)		-0.238					
		(-1.419)					
AR(9)		0.254					
		(1.353)					
AR(10)		-0.240					
		(-2.095)					
MA(1)	0.190						
	(3.989)						
MA(8)	-0.757	-0.864	-0.802	-0.467	-0.394	-0.607	-0.913
	(-11.611)	(-7.298)	(-3.977)	(-5.165)	(-2.284)	(-4.981)	(-2.626)
Constant	6.944	4.048	2.790	4.862	3.968	9.387	2.420
	(22.219)	(15.035)	(9.886)	(14.015)	(12.357)	(10.482)	(6.233)
Period	1975Q3	1985Q4	2002Q4	2000Q1	2002Q4	2002Q3	2002Q4
	2019Q3	2019Q3	2019Q3	2019Q3	2019Q3	2019Q3	2019Q3
p PMT	0.897	0.691	0.392	0.451	0.468	0.125	0.993

Notes:XR: logged nominal US-dollar exchange rate (cyclical component); t-values in parentheses. p PMT: p-value of portmanteau test for white noise.

Table A4: ARMA of real GDP (cyclical component)

	ZAF	CHL	PHL	MEX	KOR	BRA	THA
AR(1)	1.026	0.970	0.946	0.945	1.195	0.966	0.469
	(22.514)	(18.387)	(12.626)	(14.130)	(9.299)	(15.014)	(5.113)
AR(2)					-0.773		0.239
					(-3.869)		(2.029)
AR(3)		0.220	0.218		0.671		
		(2.219)	(1.764)		(3.823)		
AR(4)	-0.089	-0.354	-0.320		-0.997		
	(-1.760)	(-4.603)	(-3.034)		(-6.421)		
AR(5)					0.738		
					(3.598)		
AR(6)				-0.155	-0.623		
				(-2.748)	(-2.666)		
AR(7)	-0.141				0.712	-0.231	
	(-2.168)				(3.892)	(-2.068)	
AR(8)					-0.979	0.182	
					(-5.965)	(1.682)	
AR(9)	0.334	0.181			0.776		
	(3.658)	(2.606)			(3.836)		
AR(10)	-0.208				-0.477		
	(-2.941)				(-3.101)		
AR(11)		-0.176					
		(-2.243)					
MA(8)	-0.495	-0.681	-0.454	-0.525		-0.520	
	(-5.670)	(-7.112)	(-2.760)	(-6.119)		(-4.182)	
Constant	1.314	1.206	0.638	1.092	0.494	1.544	1.921
	(19.496)	(14.535)	(11.430)	(21.981)	(9.667)	(11.761)	(14.264)
Period	1975Q3	1988Q4	2002Q4	2000Q1	2002Q4	2002Q3	2002Q4
	2016Q4	2019Q3	2018Q4	2019Q3	2019Q1	2019Q3	2019Q3
p PMT	0.804	0.660	0.526	0.377	0.750	0.784	0.799

Notes: *GDP*: logged real GDP (cyclical component); t-values in parentheses. p PMT: p-value of portmanteau test for white noise.

G VAR(X) estimations: additional results and robustness tests

Table A5: VAR with GDP and XR, step indicator saturation

	ZAF	CHL	PHL	MEX	KOR	BRA	THA
GDP							
L.GDP	1.216***	0.491***	1.084***	0.879***	0.709***	1.117***	1.113***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
L.XR	-0.075***	-0.175***	-0.104**	0.112***	0.128***	0.006	-0.018
	(0.003)	(0.000)	(0.012)	(0.000)	(0.003)	(0.605)	(0.784)
XR							
L.GDP	2.101***	0.653***	0.870**	0.984**	0.989**	-1.004*	-0.051
	(0.004)	(0.007)	(0.035)	(0.018)	(0.014)	(0.095)	(0.857)
L.XR	0.932***	1.625***	1.125***	0.630***	0.987***	1.044***	1.123***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Lags	4	3	2	2	3	2	3
$a_{12}a_{21} < 0$	YES	YES	YES	NO	NO	YES	NO
CL 1	6.075	7.748	6.515	10.320	7.331		12.562
CL 2	3.926	2.898		2	3.218		4.075
CL 3	2.873						2
CL 4							
CL 5							
CL 6							
CL 7							
CL 8							
SI	1983, 2002	1972-73, 1976, 1981, 1983-86, 1992	1982, 2009	1976-77, 1981-82, 1984-86, 1994	1974, 1979, 1997-99, 2001	1979, 1982, 1987-92, 1994	1996, 1998-99, 2009

Notes: Sample period: 1972-2017. p-values in parentheses. GDP: logged real GDP; XR: logged nominal US-dollar exchange rate; SI: step indicator. A constant term was included in each equation (not reported). Only the coefficients on the first lags are reported. CL: cycle lengths (in years) implied by the complex eigenvalues computed as $\frac{\pi}{2\arccos(\frac{re}{mod})}$, where re is the real part of the eigenvalue and mod is the modulus.

Table A6: Estimation results for VAR with GDP and XR

	ZAF	CHL	PHL	MEX	KOR	BRA	THA
GDP							
L.GDP	1.213***	1.035***	1.235***	0.903***	0.937***	1.142***	1.392***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
L.XR	-0.077***	-0.064**	-0.162***	-0.039	0.036	-0.005	-0.002
	(0.000)	(0.025)	(0.001)	(0.144)	(0.582)	(0.383)	(0.981)
XR							
L.GDP	1.212	0.967*	0.217	1.973*	1.059*	-1.001	-0.199
	(0.138)	(0.069)	(0.671)	(0.092)	(0.079)	(0.660)	(0.562)
L.XR	1.244***	1.942***	1.376***	1.704***	1.203***	1.791***	1.103***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Lags	2	3	2	2	2	2	4
$a_{12}a_{21} < 0$	YES	YES	YES	YES	NO	NO	NO

Notes: Sample period: 1972-2017. p-values in parentheses. XR: logged nominal US-dollar exchange rate; GDP: logged real GDP. A constant term was included in each equation (not reported). Only the coefficients on the first lags are reported.

Table A7: Estimation results for VARX with GDP, XR, and CMP^W (contemporaneous)

	ZAF	CHL	PHL	MEX	KOR	BRA	THA
GDP							
L.GDP	1.154***	1.041***	1.168***	1.003***	0.878***	0.952***	1.326***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
L.XR	-0.058**	-0.053*	-0.134***	-0.003	0.015	0.011	-0.048
	(0.015)	(0.062)	(0.007)	(0.935)	(0.836)	(0.211)	(0.655)
$\mathrm{CMP^W}$	0.018*	0.041*	0.023*	0.030**	-0.007	0.072***	-0.019
	(0.097)	(0.065)	(0.068)	(0.036)	(0.525)	(0.000)	(0.287)
XR							
L.GDP	2.012***	0.966*	0.639	1.076	0.780	2.804	-0.548
	(0.008)	(0.070)	(0.173)	(0.320)	(0.235)	(0.198)	(0.140)
L.XR	0.977***	1.940***	1.199***	1.381***	1.105***	1.736***	0.786***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$\mathrm{CMP^W}$	-0.247***	-0.009	-0.144***	-0.272***	-0.035	-0.831***	-0.090**
	(0.000)	(0.910)	(0.000)	(0.001)	(0.309)	(0.004)	(0.012)
Lags	2	3	2	2	2	4	2
$a_{12}a_{21} < 0$	YES	YES	YES	YES	NO	NO	NO

Notes: Sample period: 1972-2017. *XR*: logged nominal US-dollar exchange rate; *GDP*: logged real GDP. p-values in parentheses. A constant term was included in each equation (not reported). Only the coefficients on the first lags are reported. The VAR for Chile exhibits serial correlation on the first lag.

Table A8: Estimation results for VARX(p) with GDP, XR, CMP^W and $FFUND$

	ZAF	CHL	PHL	MEX	KOR	BRA	THA
GDP							
L.GDP	1.200***	0.888***	1.160***	0.838***	0.831***	1.010***	1.270***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
L.XR	-0.072***	-0.059**	-0.124***	-0.051	-0.005	0.002	-0.059
	(0.002)	(0.027)	(0.008)	(0.176)	(0.945)	(0.754)	(0.573)
L.CMP$^{\mathrm{W}}$	0.003	-0.003	0.003	-0.001	-0.015	0.040	-0.039*
	(0.863)	(0.899)	(0.851)	(0.950)	(0.206)	(0.141)	(0.061)
L.FFUND	-0.000	-0.010***	-0.006***	-0.005**	0.000	-0.000	-0.003
	(0.804)	(0.006)	(0.001)	(0.025)	(0.883)	(0.947)	(0.236)
XR							
L.GDP	1.866**	1.408***	0.812	2.022*	0.821	1.889	-0.426
	(0.023)	(0.009)	(0.163)	(0.071)	(0.189)	(0.422)	(0.285)
L.XR	1.066***	1.919***	1.180***	1.601***	1.105***	1.621***	0.879***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
L.CMP$^{\mathrm{W}}$	-0.173*	0.067	-0.083	-0.111	0.033	-1.083***	-0.046
	(0.092)	(0.403)	(0.113)	(0.228)	(0.356)	(0.006)	(0.296)
L.FFUND	0.006	0.037***	0.016**	0.037***	0.014**	-0.026	0.003
	(0.549)	(0.007)	(0.026)	(0.006)	(0.036)	(0.465)	(0.471)
Lags	2	3	6	3	2	2	2
$a_{12}a_{21} < 0$	YES	YES	YES	YES	YES	NO	NO

Notes: Sample period: 1972-2017. p-values in parentheses. GDP: logged real GDP; XR: logged nominal US-dollar exchange rate; CMP^W: logged commodity export price index. $FFUND$: real federal funds rate. A constant term was included in each equation (not reported). Only the coefficients on the first lags are reported.

H Additional results on external factors

H.1 Co-movements in exchange rates and output across countries

Table A9: Cross-country correlation in nominal US-dollar exchange rates and real GDP

	XR		***GDP***	
	Corr	PC	Corr	PC
DEEs	0.411	0.529	0.402	0.540
AEs	0.642	0.723	0.634	0.717

Notes: Corr: average of bilateral cross-country correlation coefficient; PC: variance explained by first principal component; *XR*: logged nominal US-dollar exchange rate (cyclical component extracted from Hamilton's filter, 2002Q4 – 2019Q3); *GDP*: logged real GDP (cyclical component extracted from Hamilton's filter, 2002Q4 – 2016Q4); DEEs: South Africa, Brazil, Chile, Mexico, South Korea, the Philippines, Thailand. AEs: United Kingdom, France, Norway, Sweden, Canada, Japan, Australia. For France, the franc exchange rate was used before 1991Q1 and the euro exchange rate after.

H.2 Common dynamic factors in DEEs' exchange rates and their correlates

The dynamic factor model can be written as:

$$x_t = \Lambda F_t + u_t \tag{H.1}$$

$$F_t = \sum_{i=1}^{r} \Phi_i F_{t-i} + \eta_t, \tag{H.2}$$

where x_t is a 7×1 vector of (detrended) nominal US-dollar exchange rates, Λ is a matrix of factor loadings, and F_t is a vector of common factors, which are assumed to follow a vector-autoregressive process. Any residual autocorrelation is captured by u_t. The model is written in state-space form and estimated by maximum likelihood using the Kalman filter. We assume that the factor follows an AR(2) process and set $r = 2$.

The ARDL of the dynamic factor F_t can be written as

$$\Omega(L)F_t = \beta + \sum_{j=1}^{k} \Gamma(L)z_{tj} + \epsilon_t, \tag{H.3}$$

where $\Omega(L)$ is a lag polynomial of autoregressive terms of order m and $\Gamma(L)$ is a lag polynomial of order n_j representing the distributed lag of the $j = 1, ..., k$ regressors. A combination of m and n_j was chosen that minimises the Akaike information criterion, allowing for a maximum of four lags.

Table A10: ARDL of common dynamic factor in nominal US-dollar exchanges on external variables

	(1)	(2)	(3)	(4)
L.F	0.854***	0.908***	0.449***	0.479***
	(0.000)	(0.000)	(0.000)	(0.000)
FFUND	0.003			-0.001
	(0.186)			(0.129)
L.FFUND	-0.000			
	(0.985)			
L2.FFUND	-0.007**			
	(0.037)			
L3.FFUND	0.004**			
	(0.025)			
VXO		0.006		-0.000
		(0.355)		(0.943)
L.VXO		0.020***		0.004
		(0.010)		(0.458)
L2.VXO		-0.011		0.002
		(0.149)		(0.693)
L3.VXO		-0.011*		-0.009**
		(0.070)		(0.030)
CMP			-0.033***	-0.034***
			(0.001)	(0.001)
L.CMP			-0.067***	-0.061***
			(0.000)	(0.000)
L2.CMP			0.052***	0.047***
			(0.000)	(0.000)
Constant	-0.053	-0.013	0.042	-0.030
	(0.773)	(0.930)	(0.678)	(0.804)
p Wald FFUND	0.112			0.129
p Wald VXO		0.001		0.226
p Wald CMP			0.000	0.000
Period	2003Q4	2003Q4	2003Q4	2003Q4
	2019Q3	2019Q3	2019Q3	2019Q3
Adj. R-squared	0.726	0.761	0.897	0.900

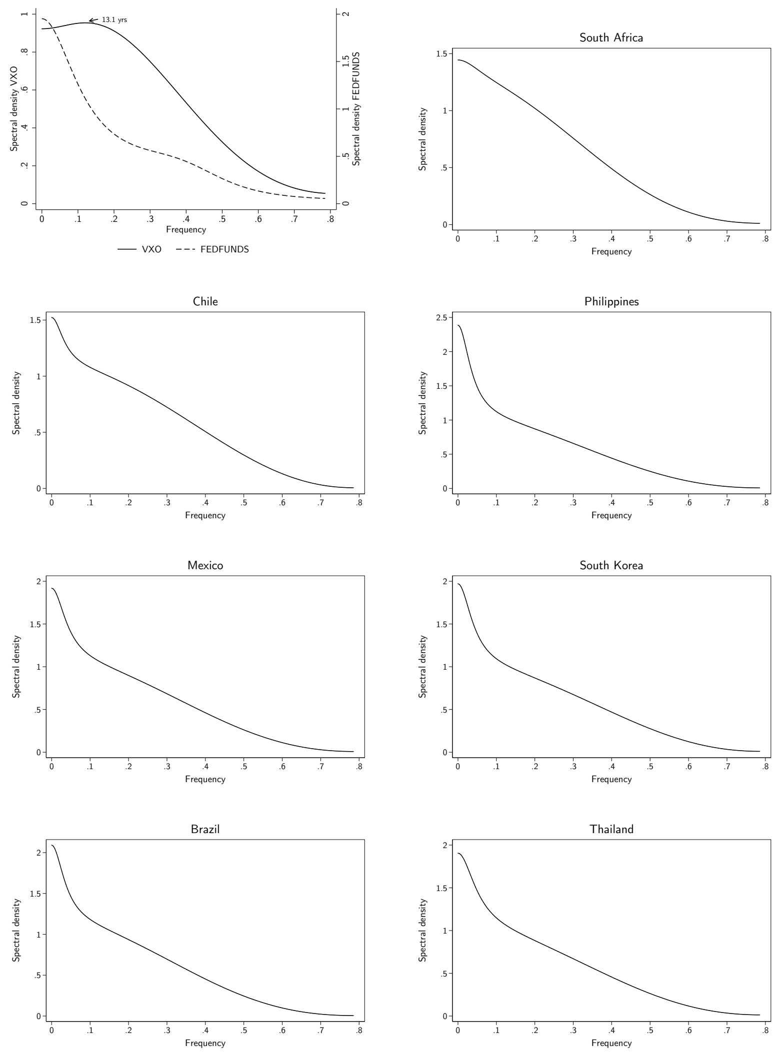
Notes: Dependent variable: dynamic factor extracted from a dynamic factor model of the logged nominal US-dollar exchange rate for South Africa, Brazil, Chile, Mexico, South Korea, the Philippines, Thailand. The dynamic factor was specified as an AR(2) process. *FFUND*: real federal funds rate, *VXO*: logged implied volatility index, *CMP*: logged global commodity price index. All variables are cyclical component extracted from Hamilton's filter. p-values in parentheses.

H.3 Cyclical properties of external variables

Figure A5: Real federal funds rate and VXO (upper left panel) and country-specific commodity export price index (remaining panels)

Notes: VXO: logged volatility index; *FEDFUNDS*: real federal funds rate. All variables are cyclical component extracted from Hamilton's filter.

Figure A6: Spectral densities of real federal funds rate and VXO (upper-left panel), and country-specific commodity export price index (remaining panels)



Notes: VXO: logged volatility index, *FEDFUNDS*: real federal funds rate. Parametrically estimated spectral densities from ARMA models of cyclical components (see Appendix F).