Industrial and Corporate Change, 2022, **00**, 1–22
DOI: https://doi.org/10.1093/icc/dtac036
Original Article

OXFORD

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

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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 theories, 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.

JEL classification: E12, E32, F31, C32

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. 1994 in Mexico and 1997–1998 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 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 stabilizing features of flexible exchange rates for DEEs with liberalized financial accounts. 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; Ocampo, 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 financialization of DEEs argues that flexible exchange rates invite speculative behavior by international investors that amplifies exchange rate volatility (Andrade and Prates, 2013; Bortz and Kaltenbrunner, 2017; de Paula *et al.*, 2017). 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 combine the financial channel of exchange rates, according to which currency depreciations are contractionary due to foreign-currency debt, 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 emerge. Procyclicality is thus hypothesized to stem from an endogenous cycle mechanism as opposed to exogenous shocks only. 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, in line with the theoretical model, these cycles are often closely aligned with cycles in real gross domestic product (GDP).

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 indeed stems from our hypothesized mechanism. Estimation results from vector-autoregressions (VARs) 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 (1972–2019). For Mexico and South Korea there is some, but less robust, evidence for a cycle mechanism; presumably 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 cycle mechanisms with nominal exchange rates 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 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). This literature typically uses panel data to estimate the response of domestic macroeconomic variables to external shocks. Through interaction terms, it is assessed whether countries with flexible exchange rates absorb these shocks better. By contrast, we employ time-series analysis to test for the presence of an interaction mechanism between GDP and the nominal exchange rate that can give rise to endogenous fluctuations. While we also compare results across different exchange rate regimes, our main interest lies in testing the hypothesis that flexible nominal exchange rates can be endogenous drivers of cycles. In this way, our approach is closer in spirit to the recent research on financial

1 The idea that (sterilized) 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).

cycles in credit and house prices that explores endogenous sources of fluctuations as opposed to shocks, but has not investigated flexible exchange rates (Aldasoro *et al.*, 2020; Borio, 2014; Stockhammer *et al.*, 2019; Strohsal et al., 2019).

Our analysis also 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 associated with a global financial cycle (Bonizzi and Kaltenbrunner, 2018; Cunha *et al.*, 2020; IMF, 2012; Kalemli-Özcan, 2019; Ocampo, 2016; Rey, 2015). Our stylized model accounts for these external factors and indeed requires exogenous shocks to set cyclical dynamics in motion. In our estimations, we control for various external drivers accordingly. However, we emphasize 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 synchronized than those of major advanced economies (AEs). Adding to the findings in Aldasoro *et al.* (2020) on the relative independence of domestic financial cycles from the global one, we thus conclude that external factors do not tell the full story. We suggest combining external factors with endogenous cycle mechanisms to explain exchange-rate driven boom-bust cycles in DEEs.[2]

Our argument casts doubt on the unambiguous benefits of flexible exchange rates. While the rigidity of fixed exchange rate regimes and their exposure to speculative attacks have rendered them unattractive for many DEEs, flexible exchange rates are not a panacea either. Indeed, the presence of a financial channel in conjunction with an external adjustment channel can turn the nominal exchange rate into a variable that amplifies macroeconomic instability. Managed floating may reduce some of the resulting volatility but does not mute endogenous mechanisms altogether and should thus be combined with capital controls (Botta *et al.*, 2021; Frenkel, 2006; Ferrari-Filho and De Paula, 2008; Guzman *et al.*, 2018; Kaltenbrunner and Painceira, 2017; 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. The last section summarizes and spells out theoretical and political implications.

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 collapsing fixed exchange rate regimes in 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). Under fixed exchange rates, cycles were often driven by an overaccumulation of external debt during the boom that became contractionary during the bust. Exchange rate dynamics were confined to changes in the *real* exchange rate via domestic inflation. However, the shift of many large DEEs toward 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) and can be found in modern New Keynesian general equilibrium models (Galí and Monacelli, 2005),

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

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. Several econometric studies compare the macroeconomic volatility of different exchange rate regimes, and provide empirical support for lower macroeconomic volatility of floating and managed floating exchange rates (Broda, 2004; Edwards and Yeyati, 2005; Ghosh *et al.*, 2015; Hoffmann, 2007; Obstfeld *et al.*, 2019).

However, structuralist development economists such as José Antonio Ocampo (2002:) 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 to deleverage. Opposite dynamics take place during periods of appreciation. This sets apart DEEs from AEs, whose private sectors typically borrow in 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). 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 appear 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; Gopinath *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

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.

model. We stress that this is a highly stylized model whose sole purpose is to emphasize the endogenous cycle mechanism between exchange rates and output. Our short-run 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), financial Dutch disease (see Botta 2021; Botta *et al.* 2021), and long-run effects of exchange rate undervaluation on technological change (see Bresser-Pereira 2012; Frenkel 2008; Guzman *et al.* 2018; Razmi *et al.* 2012).

Many applied analyses in short-run open-economy macroeconomics continue to take the classic Mundell-Fleming model (MFM) as a starting point (e.g. Blanchard *et al.* 2010; Gourinchas 2017). However, the MFM has been criticized from a post-Keynesian perspective for various unrealistic assumptions such as exogeneity of the money supply and of exchange rate expectations, and the absence of sterilized FX interventions (Serrano and Summa, 2015). Our model overcomes some of these weaknesses and could be regarded as a post-Keynesian alternative to the MFM that retains its simplicity but with added realism.[4] First and foremost, we introduce the financial channel of exchange rates and combine it with sluggish exchange rate adjustment to allow for cyclical dynamics. As will be shown below, a key implication of this feature is that the model encompasses both external and endogenous sources of cyclical dynamics. Second, instead of the two extreme cases of fully fixed or fully flexible exchange rates, our model depicts the case of managed floating nowadays pursued by many large DEEs (Frankel, 2019). In a managed float, both the exchange rate and foreign reserves adjust. The relative flexibility of the two will depend on how strongly the central bank smooths exchange rate fluctuations (or regulates the financial account).

Third, instead of an LM-curve, we treat the short-term interest rates as exogenously determined by domestic monetary policy (Serrano and Summa, 2015). Fourth, changes in foreign reserves are sterilized or "compensated" and do not impact the short-term interest rate (or other variables in the model) (Frenkel, 2008; Ghosh *et al.*, 2016). Fifth, we assume that there are Chartist traders in financial markets that form extrapolative expectations about future exchange rates. The usage of such heuristic behavioral rules is well-documented[5] and constitutes a practical reason for smoothing exchange rates through managed floating. In addition, we introduce a risk premium on the cost of foreign-currency debt that is decreasing in the value of the domestic currency (Blanchard *et al.*, 2010; Botta, 2021) capturing risks from currency mismatches as well as exogenous changes in risk appetite (e.g. the global financial cycle). Sixth, in line with the recent empirical work on "dominant currency pricing" (Adler *et al.*, 2020; Gopinath *et al.*, 2020), we assume that all trade (imports and exports) and financial flows are denominated in a dominant foreign currency—the US dollar. In conjunction with price stickiness, this means that the relevant exchange rate is the bilateral nominal US dollar rate.

We start out with a static version whose equations are given by:[6]

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

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

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

4 While we take the MFM as a starting point, our model could also be derived from a Minskyan angle. In Minskyan approaches, a key financial variable typically is a debt-to-income ratio. For DEEs, this could be an external debt ratio. However, most Minskyan approaches to DEEs were developed in the context of fixed exchange rate regimes. Our model highlights that the Minskyan effects of external debt are mediated by flexible nominal exchange rates. See Kohler (2019) for a Minskyan model with flexible exchange rates that models external debt dynamics explicitly. Yilmaz and Godin (2020) present a fully-specified SFC model. Kohler (2021) provides a survey of post-Keynesian and structuralist models of boom-bust cycles in DEEs.

5 See Dosi *et al.* (2020) for an agent-based macro-model in which a complex and uncertain environment renders the use of such simple forecasting rules rational.

6 The notation A_Y represents the partial derivative of A with respect to Y.

$$F = F[i - i^f - s^e - \rho(s)]; \qquad F_{i-i^f-s^e-\rho} = F_\delta > 0, \rho_s > 0 \tag{4}$$

$$s(dR) = sF + sNX, \tag{5}$$

where Y^d is domestic aggregate demand, composed of domestic absorption (A) and net exports priced in foreign currency (NX) and converted into domestic currency through the exchange rate (s) (defined as the price of foreign currency in domestic currency). Domestic absorption depends positively on output (Y), negatively on the domestic interest rate (i), and negatively on the exchange rate. The last assumption captures the financial channel of exchange rates, whereby depreciation against the US dollar tightens borrowing constraints and discourages private spending.[7] Net exports are a negative function of domestic output (through import demand) and positively related to foreign output (Y^f) as well as 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.[8] Net capital inflows[9] (F) are assumed to be a function of the differential between the exogenous domestic (i) and foreign (i^f) rates of interest as well as the expected exchange rate (s^e) and a risk premium ρ. The premium captures exogenous risk perceptions of global investors and is increasing in the exchange rate due to risk from currency mismatches. The last equation is the balance-of-payments, with dR representing changes in foreign reserves.[10] This equation can also be interpreted as the net supply of foreign currency to the domestic economy.

Next, we consider a dynamic extension of the model. Actual output adjusts sluggishly to changes in aggregate demand: $Y_{t+1} - Y_t = \alpha(Y_t^d - Y_t)$, with $\alpha \in (0,1)$, e.g. due to time lags in economic production and in the procurement of intermediate inputs.[11] Likewise, balance-of-payments disequilibria ($sF + sNX \neq 0$) lead to gradual adjustments in the exchange rate: $s_{t+1} - s_t = -\beta(sF_t + sNX_t)$, such that losses in reserves are accompanied by currency depreciation and vice versa. This specification captures the practice of managed floating where both the exchange rate and the stock of foreign reserves (through sterilized foreign exchange intervention) respond to exchange market pressure (Frankel, 2019). The parameter β specifies the degree of exchange rate management. In the limit, the exchange rate is either fixed and all the adjustment is carried out by changes in reserves ($\beta \to 0$) or the exchange rate is fully flexible and reserves are constant ($\beta \to \infty$). We assume that $\beta \in (0, \infty)$, so that the adjustment of reserves satisfying (5) is accompanied by some adjustment in the exchange rate.[12] Finally, we assume naive backward-looking expectations to capture extrapolative expectation formation in an environment of uncertainty: $s^e = s_t$.[13]

We then have:

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

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

7 Depreciation might also be contractionary due to adverse distributional effects (Ribeiro *et al.*, 2017).

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

9 Net capital flows in this model are ex ante or "notional" flows (Botta, 2021) as opposed to the ex post flows recorded in balance-of-payments data.

10 For simplicity, the model abstracts from interest payments on foreign debt.

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

12 See, e.g., Bhaduri (2003: pp.171-172) and Gandolfo (2016: pp. 208-210, 218-219) for similar specifications that use the balance-of-payments minus the change in reserves to pin down exchange rate dynamics. An alternative specification can be found in Botta (2021: p. 187) who uses the change in reserves as an argument in a generic function for exchange rate dynamics. Such a specification would lead to similar results.

13 See Dosi *et al.* (2020) for more sophisticated heuristics. In their agent-based model, the naive forecasting rule performs well and is frequently used even if agents are allowed to switch to more sophisticated rules. Switching between different rules could introduce another mechanism for endogenous exchange rate cycles, see, e.g., Bauer *et al.* (2009).

The Jacobian matrix evaluated at the steady state is given by:

$$J(Y^*, s^*) = \begin{bmatrix} J_{11} & J_{12} \\ J_{21} & J_{22} \end{bmatrix} \tag{8}$$
$$= \begin{bmatrix} 1+\alpha(A_Y + s^* NX_Y - 1) & \alpha(A_s + NX^* + s^* NX_s) \\ -\beta s^* NX_Y & 1-\beta[F^* + NX^* + s^*(NX_s - F_\delta(1+\rho_s))] \end{bmatrix}.$$

Importantly, the sign of J_{12} will depend on the importance of the financial relative to the trade channel. If the financial channel dominates the trade channel, i.e. $|A_s| > NX_s$, then J_{12} is likely to be negative (even more so if the trade balance is in deficit). Given the strong evidence in favor of the financial channel (Avdjiev *et al.*, 2019; Banerjee *et al.*, 2020; Kearns and Patel, 2016) and the weakness of the trade channel due to dominant currency pricing (Adler *et al.*, 2020; Gopinath *et al.*, 2020), we assume that $J_{12} < 0$. By contrast, the term J_{21} will be positive as an increase in domestic demand reduces net exports. Taken together, we thus have opposite algebraic signs on the off-diagonal elements of the Jacobian matrix: $J_{12}J_{21} < 0$. 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 bivariate models.[14] In section 5, we test empirically whether this condition is satisfied.

Figure 1 displays a numerical simulation of a linearized version of the model in (8).[15] 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. 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. As balance sheets improve, the risk premium declines further, attracting more capital inflows. 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.

Thus, unlike in the MFM and other conventional models in which the exchange rate is represented as a shock absorber, the presence of a strong financial channel turns the exchange rate into a driver of endogenous cycles. An increase in exchange rate flexibility (a higher β) would *increase* the amplitude of cycles in our model.[16] Taken together, our simple framework thus 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 exchange rate interaction mechanism that generates periodic cyclical fluctuations.

3. Empirical evidence of exchange rate cycles

From the theoretical perspective 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:

14 To see this, observe that the eigenvalues of the Jacobian matrix are the roots of the characteristic equation $\lambda^2 - \lambda Tr(J) + Det(J) = 0$, where $Tr(J)$ and $Det(J)$ are the trace and determinant of the Jacobian matrix, respectively. The roots of the characteristic equation are given by $\lambda_{1,2} = \frac{Tr(J) \pm \sqrt{Tr(J)^2 - 4Det(J)}}{2}$. Complex roots, which give rise to oscillations, emerge when the discriminant of this expression becomes negative. This requires $(J_{11}+J_{22})^2 - 4(J_{11}J_{22} - J_{12}J_{21}) < 0$, 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).

15 In line with our discussion of the empirical evidence, we consider a strong financial and a weak trade channel ($|A_s| > NX_s$). No attempt was made to calibrate the model to a specific country.

16 To see this, note that in a model with complex eigenvalues, the amplitude of cycles is governed by the modulus $|\lambda|$. In a two-dimensional discrete-time model, the modulus is given by $|\lambda| = \sqrt{Det(J)}$, where $Det(J)$ is the determinant of the Jacobian. In our case, $\frac{\partial Det(J)}{\partial \beta} = \frac{-J_{11}(1-J_{22})}{\beta} - \frac{J_{21}J_{12}}{\beta}$, which will be positive if $-J_{11}(1-J_{22}) > J_{21}J_{12}$, i.e. if the financial channel is strong relative to the trade channel. By contrast, in a MFM where the trade channel dominates, eigenvalues will be real and the amplitude of fluctuations will be *decreasing* in β.

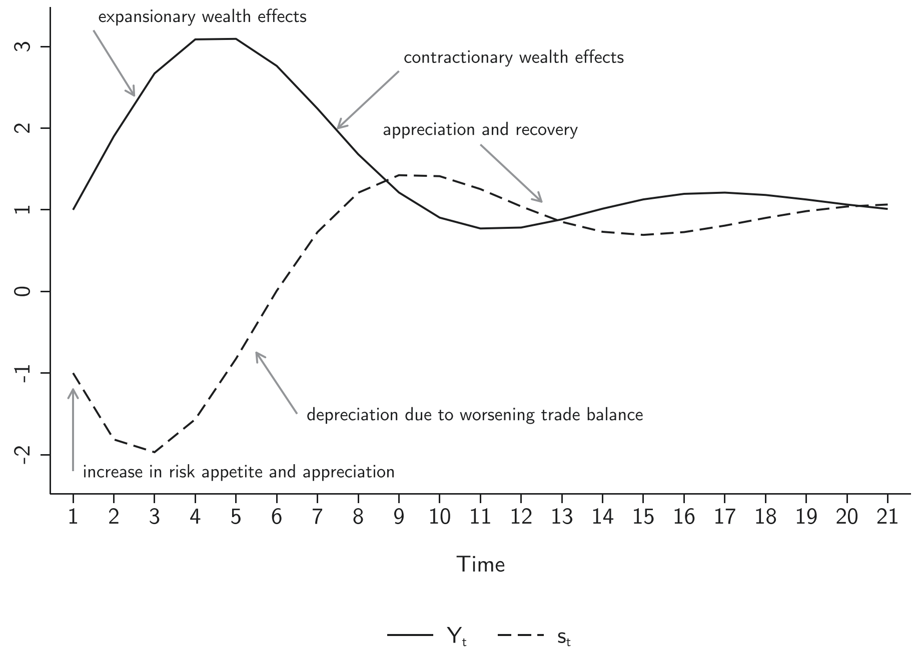

Figure 1. Model simulation: a shock to risk appetite triggers pro-cyclical exchange rate dynamics
Notes: Numerical simulation of a linearized version of (8) with Y_t: output; s_t: nominal exchange rate (price of foreign currency in domestic currency). Parameterization of Jacobian: $\alpha = 0.5$, $\beta = 0.7$, $A_y = 0.7$, $A_s = -1$, $NX_y = -0.8$, $NX_s = 0.1$, $F_\delta = 0.1$, $\rho_s = 0.1$, $s^* = 1$, $Y^* = 1$, $F^* = 0$, $NX^* = 0$, which yields: $J_{11} = 0.45$, $J_{12} = -0.45$, $J_{21} = 0.56$, $J_{22} = 1.007$. For this parameterization, the roots of the Jacobian are a pair of complex conjugates with modulus < 1. A shock-sequence ($\rho_t = 0.8\rho_{t-1}$) was added to (7) to emulate a sustained increase in risk appetite.

South Africa, Chile, the Philippines, Mexico, South Korea, Brazil, and Thailand.[17] We expect the endogenous cycle mechanism with nominal exchange rates to appear only in economies with a sufficient degree of exchange rate flexibility. However, to confirm this theoretical assumption and to assess the susceptibility of our empirical method to false-positives, we also consider some countries with chaotic or fixed exchange rate regimes. Our sample selection is thus governed by a combination of the need for heterogeneity and data availability constraints.[18]

We identify a first group of countries in our sample with relatively long spells of semi-flexible or flexible exchange rate regimes[19] (South Africa, Chile, and with exceptions the Philippines)[20]; 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).[21] We hypothesize a decreasing chance of finding regular cycles in nominal exchange rates across these three groups.

In line with the literature on the financial channel of exchange rates and dominant currency pricing (Avdjiev *et al.*, 2019; Gopinath *et al.*, 2020), we focus on the (logged) bilateral *nominal* exchange rate with the US dollar (henceforth *XR*), defined as domestic currency units per foreign currency unit. This allows us to isolate nominal exchange rate dynamics from changes in relative

17 Detailed information on the dataset can be found in Supplementary Material A.
18 In particular, we are constrained by the small number of DEEs that had floating exchange rate regimes for more than two decades.
19 We use the exchange rate regime classification in Ilzetzki *et al.* (2019). See Supplementary Material B for details.
20 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).
21 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.

prices (which are driven by different economic mechanisms). Our preferred detrending method is the regression filter proposed in Hamilton (2018), but we also compare it with a growth rate filter to assess its robustness.[22]

Figure A2 in Supplementary Material C reports the cyclical components in US-dollar exchange rates for the full sample (1972Q1–2019Q3). For most countries after the end of chaotic episodes, a period of floating exchange rates began that is characterized by much more regular exchange rate fluctuations.[23] 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. Figure 2 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.[24] 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 behavior in *XR* and *GDP*, often in form of a negative co-movement, which is especially strong in Chile, South Korea, and Brazil.[25]

To examine the periodicities in *XR* and *GDP* more rigorously, we estimate spectral density functions (Hamilton, 1994: Ch. 6). 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.[26] 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 (CL) 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.

Figures 3 and 4 display univariate spectral densities for *XR* and *GDP*. All countries, except for Brazil, exhibit a dominant cycle frequency in *XR*. Estimated CLs 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 centered on a dominant peak. Periodicities appear to be less pronounced in the Asian countries, where spectral densities are more dispersed around the peak, presumably due to the shorter sample period. Supplementary Material C 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 CLs tend to be shorter (by around one to two

22 We construct the Hamilton filter as the residual ν 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 (annualized) 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.

23 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.

24 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.

25 Figure A3 in Supplementary Material C compares the Hamilton-filtered *XR* series with the growth rate filter. It can be seen that the results are qualitatively very similar.

26 The main advantage of parametric estimation over non-parametric approaches is its efficiency as it requires fewer degrees of freedom. The methodology for the parametric estimation of spectral density functions as well as the underlying estimated ARMA models are reported in Supplementary Material C.

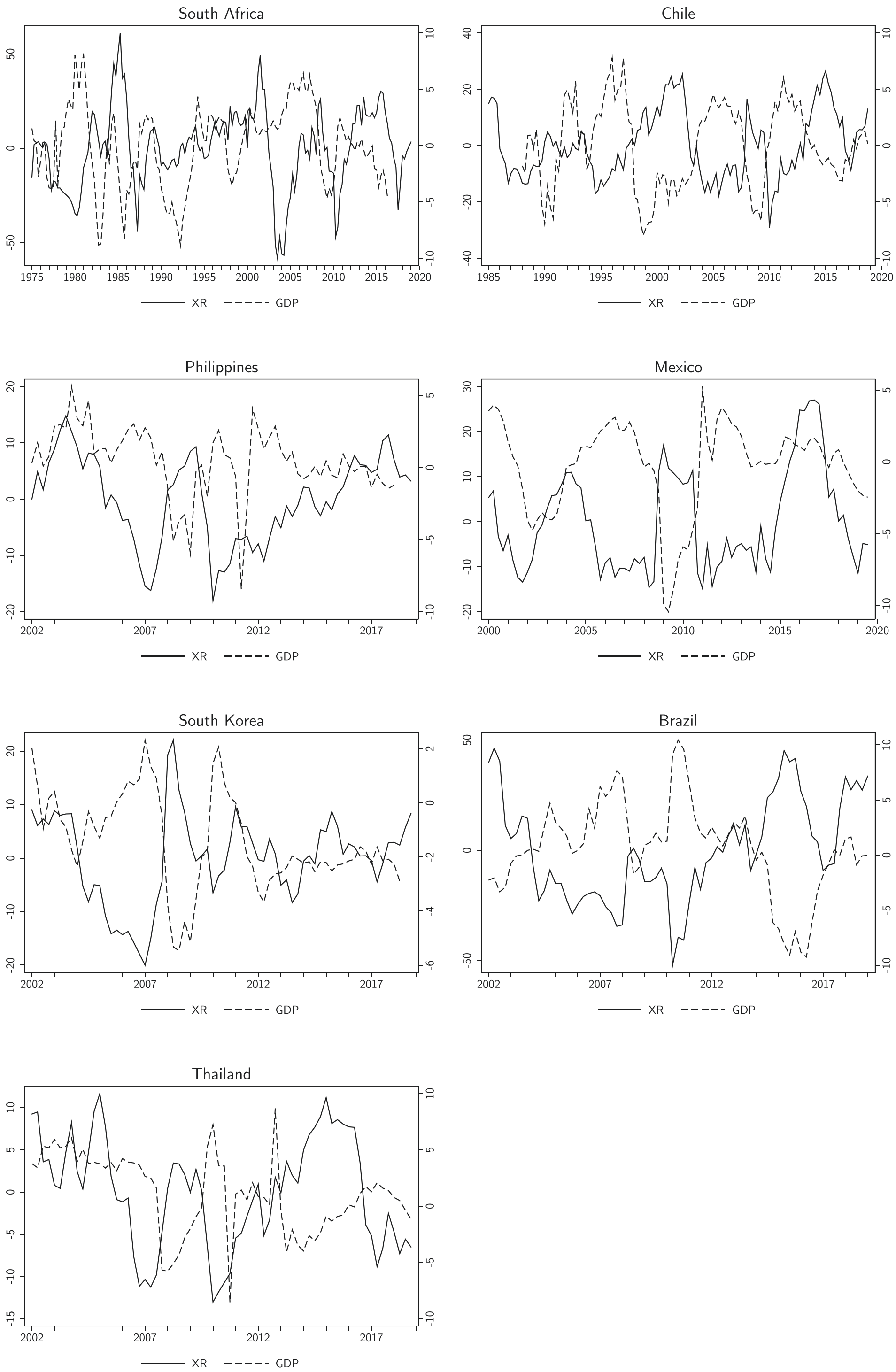

Figure 2. Cycles in US-dollar exchange rates (left scale) and GDP (right scale)
Notes: XR: logged nominal US-dollar exchange rate (cyclical component); *GDP*: (seasonally adjusted) logged real GDP (cyclical component). Cyclical components were extracted through Hamilton's regression filter and are expressed in per cent deviation from trend.

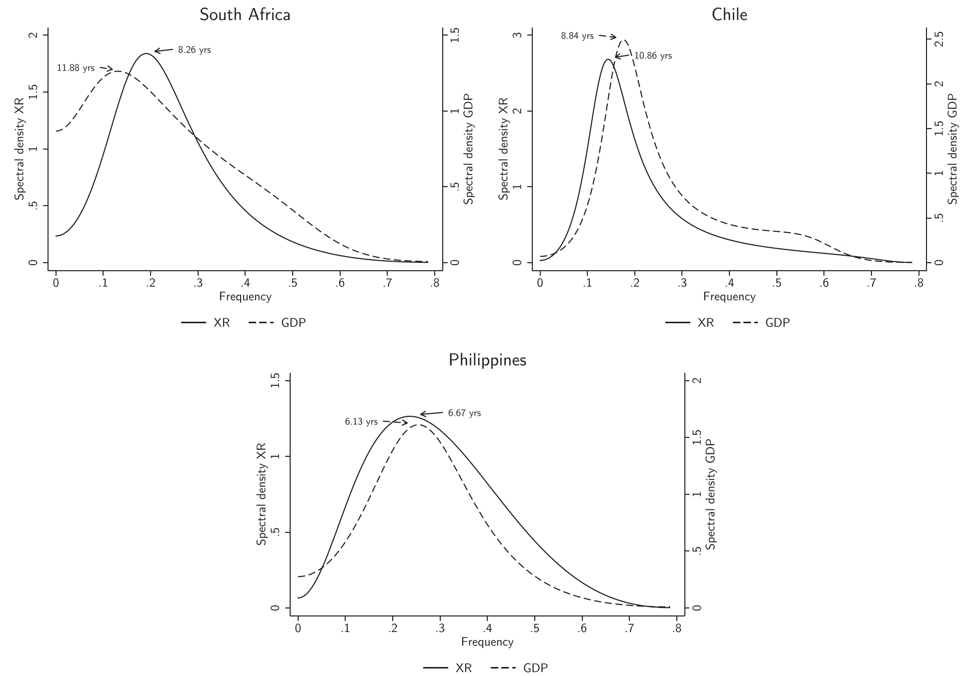

Figure 3. Spectral densities of US-dollar exchange rates and GDP: South Africa, Chile, the Philippines
Notes: XR: logged nominal US-dollar exchange rate (cyclical components); *GDP*: logged real GDP (cyclical components). Cyclical components were extracted through Hamilton's regression filter. Spectral densities were estimated parametrically from ARMA models (see Supplementary Material C). Arrows indicate the CL (in years) associated with the peak in the spectral densities.

years). Estimated CLs thus come with some uncertainty, but the qualitative presence of regular periodicities does not appear to be sensitive to the detrending method.[27]

Overall, the descriptive evidence suggests that several major DEEs exhibit periodic exchange rate cycles ever since they embarked on floating exchange rate regimes. Estimated CLs in exchange rates range between 4 1/2 and 11 years. These cycles are often closely aligned with cycles in GDP, especially in South Africa, Chile, the Philippines Mexico, and South Korea. 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 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) and (7),

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

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

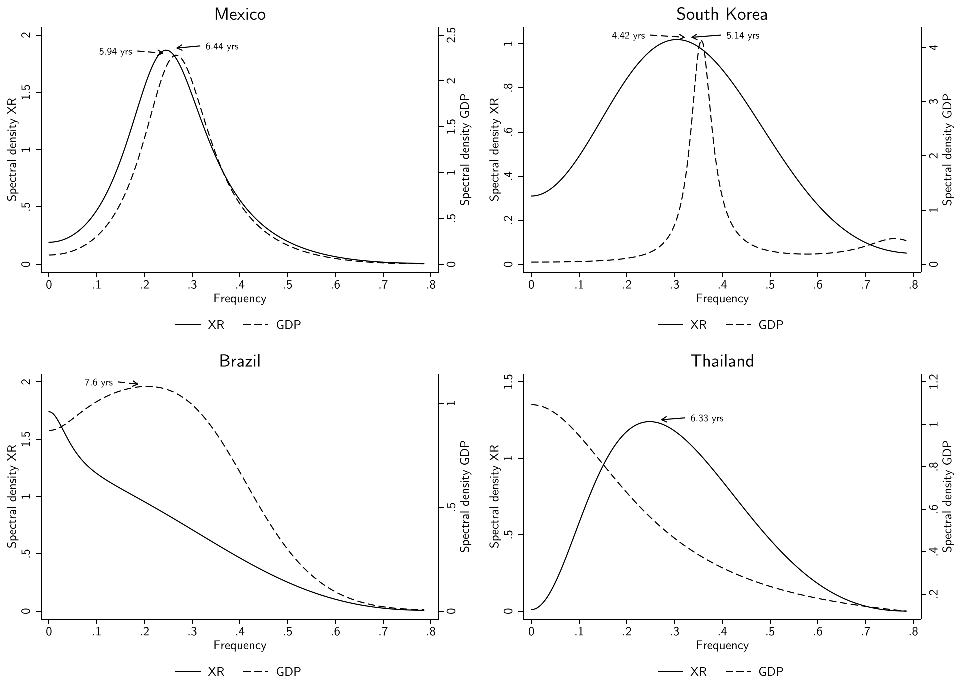

Figure 4. Spectral densities of US-dollar exchange rates and GDP: Mexico, South Korea, Brazil, Thailand
Notes: XR: logged nominal US-dollar exchange rate (cyclical components); *GDP*: logged real GDP (cyclical components). Cyclical components were extracted through Hamilton's regression filter. Spectral densities were estimated parametrically from ARMA models (see Supplementary Material C). Arrows indicate the CL (in years) associated with the peak in the spectral densities.

where the endogenous variables are collected in $\mathbf{y_t} = [Y_t, s_t]'$ and the exogenous variables in $\mathbf{z_t}$. The coefficient matrix A corresponds to the (linearized) Jacobian matrix (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.[28]

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

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

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 the two coefficients a_{12} and a_{21} of the reduced-form coefficient matrix (as opposed to contemporaneous coefficients that would be needed to identify impulse response functions).

28 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.

To estimate the VARX in (10) with real GDP and nominal US-dollar exchange rates, we use annual data.[29] 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 (at the 5%-level of significance) 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 $j = h = 1$ for parsimony.[30]

In order to maximize degrees of freedom, we use the full time span and set the sample start to 1972. As a result, crises episodes visible 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 controlling for external drivers. Second, we also report a bivariate VAR specification with step indicator saturation 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.[31]

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.[32] Second, we consider the real US monetary policy rate, defined as the Federal Funds rate minus the (annualized) US CPI inflation rate (*FFUND*).[33] 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.[34]

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.

29 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.

30 We check the robustness of our baseline specification to the case where $j = h = 0$.

31 Estimating VARs in levels is common, especially when the variables can be a mix of I(0) and I(1). A VAR 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). 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.

32 The index covers international prices of 45 commodities that are deflated by a unit value index for manufactured exports. See Supplementary Material A for further information.

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

34 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).

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

GDP	ZAF	ZAF	ZAF	CHL	CHL	CHL	PHL	PHL	PHL
L.GDP	1.227***	1.231***	1.457***	1.017***	0.883***	0.653***	1.226***	1.192***	1.196***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)
L.XR	–0.068***	–0.067***	–0.037*	–0.065**	–0.059**	–0.132**	–0.130***	–0.113**	–0.030
	(0.003)	(0.002)	(0.077)	(0.020)	(0.023)	(0.047)	(0.008)	(0.014)	(0.497)
L.CMP^W	0.003			0.009			0.016		
	(0.832)			(0.699)			(0.231)		
L.FFUND		–0.001			–0.010***			–0.004***	
		(0.643)			(0.005)			(0.008)	
L.VXO			–0.008			–0.017			–0.006
			(0.272)			(0.126)			(0.560)
XR									
L.GDP	1.863**	1.331*	1.268	0.918*	1.471***	1.497***	0.459	0.495	1.772**
	(0.019)	(0.083)	(0.216)	(0.087)	(0.005)	(0.008)	(0.351)	(0.299)	(0.017)
L.XR	1.068***	1.135***	1.188***	1.938***	1.922***	1.584***	1.210***	1.205***	1.145***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
L.CMP^W	–0.205**			0.025			–0.124***		
	(0.011)			(0.759)			(0.006)		
L.FFUND		0.015**			0.034***			0.019***	
		(0.044)			(0.009)			(0.001)	
L.VXO			–0.050			0.019			0.083**
			(0.398)			(0.539)			(0.039)
Lags	2	2	2	3	3	2	2	2	2
$a_{12}a_{21} < 0$	YES	YES	YES	YES	YES	YES	YES	YES	YES
CL 1	7.1	7.3	10	8.8	10.6	18.5	9.9	8.4	7.1
CL 2				3.9	4	2			

Notes: Sample period: 1972–2019 (with *VXO*: 1987–2019). *p*-values in parentheses. *GDP*: logged real GDP; *XR*: logged nominal US-dollar exchange rate; *CMP*^{W}: logged country-weighted 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. ***,**, and * denote statistical significance at the 1%, 5%, and 10%-level, respectively.

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

GDP	MEX	MEX	MEX	KOR	KOR	KOR
L.GDP	0.941***	0.860***	0.806***	0.839***	0.898***	0.534**
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.028)
L.XR	−0.023	−0.043	0.077**	0.001	0.022	−0.011
	(0.424)	(0.200)	(0.032)	(0.990)	(0.739)	(0.882)
L.CMP^W	0.013			−0.016		
	(0.394)			(0.154)		
L.FFUND		−0.005**			0.001	
		(0.012)			(0.529)	
L.VXO			−0.025**			−0.027*
			(0.049)			(0.096)
XR						
L.GDP	1.565	1.968*	1.278	1.080*	0.678	2.310***
	(0.164)	(0.076)	(0.136)	(0.087)	(0.260)	(0.005)
L.XR	1.527***	1.651***	1.049***	1.209***	1.050***	1.182***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
L.CMP^W	−0.180**			0.005		
	(0.044)			(0.882)		
L.FFUND		0.045***			0.012*	
		(0.000)			(0.051)	
L.VXO			0.013			0.089
			(0.837)			(0.106)
Lags	2	3	2	2	2	2
$a_{12}a_{21} < 0$	YES	YES	NO	NO	NO	YES
CL 1	16.3	8	14.7	14.5	9.5	8.6
CL 2		3.1	2			2

Notes: Sample period: 1972–2019 (with *VXO*: 1987–2019). *p*-values in parentheses. *GDP*: logged real GDP; *XR*: logged nominal US-dollar exchange rate; CMP^W: logged country-weighted 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. ***,**, and * denote statistical significance at the 1%, 5%, and 10%-level, respectively.

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) 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 by currency depreciations are contractionary ($a_{12} < 0$) and the external adjustment channel where by output expansions lead to downward pressure on currencies ($a_{21} > 0$). The results are generally robust to the use of different control variables, 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 smaller 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 Figure 2 was most suggestive of relatively stable cycles.

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 satisfied for Mexico with CMP^W and *FFUND*, and for Korea only with *VXO*, but the signs are only partially statistically significant.

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

GDP	BRA	BRA	BRA	THA	THA	THA
L.GDP	0.995***	1.066***	1.135***	1.273***	1.440***	1.174***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
L.XR	0.001	−0.005	0.011	−0.072	0.034	−0.060
	(0.891)	(0.389)	(0.188)	(0.480)	(0.709)	(0.588)
L.CMP^W	0.041*			−0.028		
	(0.062)			(0.140)		
L.FFUND		−0.003			−0.001	
		(0.203)			(0.678)	
L.VXO			−0.013			−0.011
			(0.373)			(0.567)
XR						
L.GDP	1.701	−0.532	0.782	−0.388	−0.182	0.058
	(0.463)	(0.819)	(0.792)	(0.329)	(0.604)	(0.915)
L.XR	1.638***	1.773***	1.513***	0.944***	1.038***	1.163***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
L.CMP^W	−0.885***			−0.061		
	(0.006)			(0.121)		
L.FFUND		0.026			0.005	
		(0.400)			(0.187)	
L.VXO			−0.048			0.047
			(0.857)			(0.270)
Lags	2	2	3	2	2	2
$a_{12}a_{21} < 0$	NO	NO	NO	NO	YES	YES
CL 1	26.1	34.4	28	25.7	33.2	26.7
CL 2			6.2			
CL 3			2			

Notes: Sample period: 1972–2019 (with *VXO*: 1987–2019). *p*-values in parentheses. *GDP*: logged real GDP; *XR*: logged nominal US-dollar exchange rate; CMP^W: logged country-weighted 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. ***,**, and * denote statistical significance at the 1%, 5%, and 10%-level, respectively.

Table 3 contains the results for Brazil and Thailand, which exhibited mostly chaotic or fixed exchange rates during the sample period, rendering the presence of a cycle mechanism with the nominal exchange rate 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. This suggests that our method is not strongly prone to producing false-positives when it comes to cycle mechanisms. Only the specification with the *VXO* for Thailand yields estimates in line with the hypothesized cycle mechanism, which is likely due to shorter sample period (recall that Thailand adopted a managed floating regime in 1999).

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* 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 CLs that are implied by the coefficient matrices of the VARX.[35] Focusing on the baseline with CMP^W, we find cycle frequencies of around 7 and 10 1/2 years for South Africa, Chile, and the Philippines that are very similar to the spectrally estimated frequencies in exchange rates in section 3. For the other countries, estimated frequencies are often

35 From the polar representation of the complex eigenvalues $\lambda = |\lambda|(\cos\theta \pm i.\sin\theta)$ of the companion matrix, the implied CL 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 $A_{i+1}, i = 1,...p-1$.

substantially longer, which is likely due to the presence of currency crises episodes that are not captured by the control variables.

Next, we provide 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 Supplementary Material D).[36] 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 East Asian crisis in Korea and Thailand.

The cycle condition is again satisfied and statistically significant in South Africa and Chile, and now also statistically significant for the Philippines. 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 CLs range from 3 years (South Korea) to almost 8 years (Chile), and are generally are close to the estimated frequencies displayed in Figures 3 and 4 (except for South Africa whose estimated length of around 5 years is shorter).

Further robustness tests are reported in Supplementary Material D. The main results hold up. In a VAR 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 again for South Africa, Chile, and the Philippines.

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 partly also point toward the presence of cycle mechanisms but are less robust. For Brazil and Thailand, the cycle condition is not satisfied or not significant. These results are consistent with our hypothesis that the endogenous cycle mechanism with nominal exchange rates 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 and Thailand had a pegged exchange rate throughout most of the sample period, which may explain the complete absence of a stable interaction mechanism.

We stress that our findings do not imply that fixed exchange rate regimes exhibit more macroeconomic stability. Indeed, much of the post-Keynesian and structuralist literature discussed in section 2 analyzed boom-bust cycle episodes in fixed exchange rate regimes, in which external debt accumulation played a central role. In fixed exchange rate regimes, external debt rather than the nominal exchange rate becomes the key interacting variable in an endogenous cycle mechanism. Thus, as a final exercise, we re-estimate our baseline VARX (with CMP^W) and replace the nominal exchange rate by an external debt-to-GDP ratio (*EXDEBT*) (see Table A9 in Supplementary Material D). *EXDEBT* has the expected negative effect on *GDP* in all countries. For South Africa, Chile, Mexico, Korea, and Thailand, there is a positive feedback from *GDP* to *EXDEBT*, i.e. the condition for a cycle mechanism holds (albeit not always statistically significant). Pegging the exchange rate, as did Mexico, Korea, and Thailand throughout much of the sample period, thus does not necessarily prevent the emergence of endogenous cycle mechanisms, and indeed pegged exchange rate regimes have proven to be often volatile and unstable (Ghosh *et al.*, 2015).

36 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.

However, our main results suggest that adopting floating exchange rates (managed or not) does not fully resolve the problem of endogenous cycles either.

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. However, what is the relative importance of these external factors for exchange rate cycles in DEEs? This section summarizes results from a number of additional empirical exercise that are documented in Supplementary Material E.

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.[37] 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 some 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 of the common dynamic factor as a function of external variables. We find that commodity prices are the closest correlate of the common factor in exchange rates across DEEs.[38] 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.[39] 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 emphasize 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. The results reveal that the only variable that exhibits a dominant cycle frequency is the *VXO* with an estimated CL of 13 years. This is above the estimated frequencies for exchange rates in DEEs, which range from 4 to 11 years.[40] 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

37 United Kingdom, France, Norway, Sweden, Canada, Japan, Australia.

38 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.

39 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 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.

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

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 DEEs. Drawing on post-Keynesian and structuralist frameworks as well as recent work at the BIS, 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; Ocampo, 2016; Stiglitz *et al*., 2006). This literature has mostly emphasized destabilizing balance sheet effects from fluctuating exchange rates. We show that when combined with a standard feedback mechanism from economic activity to exchange rates, an endogenous cycle mechanism emerges. Correspondingly, flexible exchange rates in emerging markets may better be described as “drivers of endogenous cycles” rather 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 emphasized global factors (Bonizzi and Kaltenbrunner, 2018; Cunha *et al*., 2020; IMF, 2012; Kalemli-Özcan, 2019; Ocampo, 2016; Rey, 2015), we argue that they only 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 can be important triggers, but only in combination with cyclical interaction 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 destabilizing role of domestic variables such as the nominal exchange rate and foreign-currency debt that are, at least to some degree, under the control of policy makers.

Third, our argument puts into question an unconditional endorsement of flexible exchange rates. Many emerging economies now operate managed floating regimes using occasional interventions in foreign exchange markets to smooth exchange rate fluctuations. While this is likely to reduce the amplitude of cycles, our results suggest that managed floating alone does not fully mute pro-cyclical effects of nominal exchange rates. In conjunction with liberalized financial accounts, managed floating can fail to curb speculative capital flows or even attract them as investors believe central banks will prevent losses (Kaltenbrunner and Painceira, 2017). Our argument supports the view that managed floating should be combined with capital controls (Botta *et al*., 2021; Frenkel, 2006; Ferrari-Filho and De Paula, 2008; Guzman *et al*., 2018; Ocampo, 2002). Capital controls 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.

Supplementary data

Supplementary data are available at *Industrial and Corporate Change* online.

Acknowledgements

We thank an anonymous referee, Bruno Bonizzi, Robert Calvert Jump, Yannis Dafermos, Alexander Guschanski, and Annina Kaltenbrunner for helpful comments. All errors are the authors'.

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