Capital Controls and the Real Exchange Rate: Do Controls Promote Disequilibria?*

Juan Antonio Montecino

This Version: August, 2017
(Click HERE for the latest version)

Abstract

The consensus view is that capital controls can effectively lengthen the maturity composition of capital inflows and increase the independence of monetary policy but are not generally effective at reducing net inflows and influencing the real exchange rate. This paper studies the adjustment dynamics of the real exchange rate towards its long-run equilibrium and presents empirical evidence that capital controls increase the persistence of misalignments. Allowing the speed of adjustment to vary according to the intensity of restrictions on capital flows, it is shown that the real exchange rate converges to its long-run level at significantly slower rates in countries with capital controls. This result is strongest when the exchange rate is undervalued and appears to operate primarily via nominal exchange rate dynamics. In addition, controls on capital inflows have consistently greater effects than controls on outflows. The results also suggest that flexible exchange rate regimes accelerate the adjustment of disequilibria relative to managed and fixed regimes.

Keywords: Capital Flows, Capital Controls, Real Exchange Rate
JEL: F2, F31, F36, F41

* ***Acknowledgments:*** I am grateful to Arslan Razmi, Gerald Epstein, and Peter Skott for their helpful comments and suggestions. I am also greatly indebted to João Paulo De Souza for his generous advice and Andrés Fernándes for kindly sharing his data. All errors and omissions, however, are solely my own.

1 Introduction

Once considered heretical to the tenets of prudent macroeconomic policy, in recent years capital controls have regained respectability in official policy circles and received fresh attention among academics as potential macro-prudential tools.[1] In the wake of the global financial crisis and mounting evidence of the destabilizing effects of unregulated international capital flows, the International Monetary Fund (IMF), previously the champion of capital account liberalization, reversed decades of official policy recommendations and declared that capital controls should once again be included in a country's "policy toolkit."[2] At the same time, as expansionary monetary policy in industrial nations has flooded emerging markets with foreign funds, a number of countries have imposed restrictions on capital inflows, specifically citing a concern with excessive exchange rate appreciation and a desire to preserve export sector competitiveness. This shift in opinion regarding the use of capital controls has taken place along with a growing recognition that some rapidly industrializing nations, in particular China, have benefitted from so-called "neo-mercantilist" policies and have used capital controls to deliberately maintain an undervalued real exchange rate.

These calls for the greater use of capital controls to manage the real exchange rate stand at odds with the empirical literature on the effectiveness of controls, which has not found clear evidence that controls can influence this variable (for detailed reviews of this literature see Engel (2015) and Magud et al. (2011)).[3] Several empirical studies (Valdés-Prieto and Soto, 1998; Edwards, 1999; De Gregorio et al., 2000; Gallego et al., 2002; Forbes, 2003) have focused on Chile's experience with capital controls during the 1990s, which sought to limit short-term capital flows in order to stabilize the economy and prevent unwanted exchange rate appreciation. While most of these studies conclude that Chile's capital controls had a meaningful impact on the maturity composition of net inflows, the results suggest either a very small and short-term effect on the real exchange rate (e.g. De Gregorio et al., 2000), or no significant effect at all (e.g. Gallego et al., 2002).[4]

Cross-country and case studies of other capital control episodes have reached similar conclusions (Levy-Yeyati et al., 2008; Baba and Kokenyne, 2011; Klein, 2012; Jinjarak et al., 2013; Alfaro et al., 2014; Forbes et al., 2015). For example, Baba and Kokenyne (2011) look at the effects of capital controls in emerging markets during three different episodes in the 2000s – the foreign exchange tax in Brazil (2008), and the URRs in Colombia (2007-08) and Thailand (2006-08) – and one episode of capital outflow liberalization – South Korea (2005-08). Their results show that controls during the 2000s appear to have successfully altered the maturity composition and lowered the overall volume of flows in Colombia and Thailand. Controls also appear to have successfully preserved monetary policy independence in Brazil and Colombia, albeit temporarily. However, their results provide no evidence that controls in any country were able to successfully influence the real exchange rate.

This paper presents new empirical evidence on the adjustment dynamics of the real exchange rate towards its long-run equilibrium in the presence of capital controls. In contrast with previous approaches, I explicitly model the adjustment dynamics of the real exchange rate as a function of the intensity of capital controls. Using a large panel of developed and developing countries, I show that capital controls can substantially slow the speed of adjustment of the real exchange rate towards its long-run level, causing disequilibria to persist

[1]A growing theoretical literature has shown that capital controls improve welfare in models featuring financial amplification dynamics arising from collateral constraints (Lorenzoni, 2008; Jeanne and Korinek, 2010; Korinek, 2012; Korinek and Sandri, 2014; Davis and Presno, 2014; Liu and Spiegel, 2015; Heathcote, 2016). In these types of models, capital flows impose externalities because private agents fail to internalize the contribution of their borrowing decisions to systemic risk. As a result, the decentralized equilibrium is characterized by "over borrowing" and is inefficient. Capital controls in this context can be seen as a Pigouvian tax to force agents to internalize the externality. Capital controls have also been shown to improve welfare in small open economies with fixed exchange rates and rigid nominal wages (Farhi and Werning, 2013; Schmitt-Grohé and Uribe, Forthcoming).

[2]Examples of work by IMF staff articulating this change in opinion are Ostry et al. (2010), Ostry et al. (2011b), and Ostry et al. (2011a). These new perspectives on the role of capital controls became part of the IMF's "institutional view" late in 2012 (IMF, 2012).

[3]A notable exception is Erten and Ocampo (2017), who find that capital account regulations reduce real exchange rate appreciation and foreign exchange pressure.

[4]It is worth noting that there exists some evidence that Chile's capital controls may have had a significant effect on the *nominal* exchange rate. Edwards and Rigobon (2009) show that capital controls slowed the appreciation of the Chilean Peso and decreased its volatility.

for extended periods of time. Specifically, this paper uses panel dynamic ordinary least-squares (DOLS) to estimate the long-run cointegrating relationship between the real exchange rate and a set of long-run determinants. This equilibrium relationship is used to calculate the extent of real under or overvaluations – that is, of disequilibria – which are then imposed on an Error-Correction Model (ECM) to study the short-run adjustment dynamics towards equilibrium.

The empirical results are consistent with the hypothesis that on average capital controls slow the speed of adjustment towards the long-run equilibrium and therefore allow real exchange rate disequilibria to persist for longer periods of time relative to the absence of controls. The point estimates from the baseline model imply half-lives for the adjustment of disequilibria of roughly 5 years in countries with stringent restrictions on international financial transactions but as short as 2 years in countries with comparatively open capital accounts. These results therefore imply considerable differences in real exchange rate adjustment dynamics between countries depending on the intensity of capital controls. Moreover, these findings do not appear to be driven by differences in exchange rate regimes and are robust to several extensions, including allowing for asymmetries between undervaluations and overvaluations, using alternative indexes of capital controls, and additional forms of error-correction heterogeneity to account for omitted variable bias.

I also present evidence of meaningful asymmetries in real exchange rate adjustment dynamics and the impact of capital controls. Although overvaluations are on average more persistent than undervaluations, capital controls appear to have greater traction when the real exchange rate is undervalued and may even modestly help to eliminate overvaluations. I also find evidence of heterogeneity in the effect of capital controls between flexible and more rigid exchange rate regimes. While capital controls increase the persistence of misalignments in countries with fixed and intermediate or managed exchange rate regimes, controls appear to lose their effectiveness under flexible regimes. The results also point to substantial differences depending on the type of financial flow restriction. Overall, controls on capital inflows appear to have a larger and more consistently statistically significant impact than do controls on outflows. Disaggregating by financial instrument, controls on equity flows, bonds, collective investments, and foreign direct investment are particularly effective.

This paper is related to the vast literature on the empirical determinants of exchange rates. While a detailed review of this literature is beyond the scope of this paper, textbook treatments are provided in Sarno and Taylor (2002, ch. 3-4) and MacDonald (2007, ch. 8-9). A recent strand in this literature argues that in the long-run the real exchange rate is pinned down by real fundamentals, including the relative productivity of the tradable sector (the Balassa-Samuelson effect), the terms of trade, and the net foreign asset position (Chinn and Johnston, 1996; Chinn, 2000; Cashin et al., 2004; Bayoumi et al., 2005; Ricci et al., 2013; Bordo et al., 2014). Although this literature is diverse, the unifying theme is to treat the real exchange rate as nonstationary and use cointegration techniques, emphasizing explicit equilibrium relationships.

My results also shed light on the policy issue of real exchange rate misalignment. It has long been recognized that real overvaluations can negatively impact growth and may precede currency crises. Moreover, a growing literature has shown that there exists a robust relationship between an undervalued real exchange rate and faster economic growth (see, for example, Rodrik (2008) and Rapetti et al. (2012)). These positive growth effects have been explained through a variety of channels: sectoral misallocation of capital due to government and market failures (Rodrik, 2008); hidden unemployment in an underdeveloped dual economy (Razmi et al., 2012); or learning by doing externalities in the tradables sector (Korinek and Serven, 2010). What all these models have in common, however, is the importance for long-run growth of the tradable sector and the potential to use undervaluation as a development tool. But how exactly should policymakers wield this new tool? It is poorly understood how a persistent undervaluation can actually be achieved and whether restrictions on capital mobility can play a role.[5] Another contribution of this paper is therefore to help fill this gap. The empirical results presented below suggest that capital controls are capable of promoting real exchange rate undervaluation for extended periods of time, and may therefore serve as an effective instrument to manage the real exchange rate.

Finally, this paper is also related in spirit and in methodology to a recent literature investigating the

[5] See Jeanne (2012) for an in depth discussion of this point and a theoretical model of real exchange rate undervaluation with "Chinese-style" capital controls.

persistence of external imbalances under alternative exchange rate and financial regimes. Beginning with the seminal work by Chinn and Wei (2013), several papers have used extended AR(1) models of the current account balance to measure the impact of exchange rate regimes and other monetary and financial policies on the speed of mean-reversion. These papers have provided conflicting results regarding the impact of financial openness and exchange rate flexibility. For example, Chinn and Wei (2013) finds no evidence that flexibility matters but that financial openness is significantly associated with slower mean-reversion. In contrast, Herrmann (2009) finds that exchange rate flexibility leads to faster mean-reversion while financial openness has no discernible effect. Eguren-Martín (2015) reports evidence in support of the link between flexibility and faster current account adjustment while offering mixed results regarding the impact of financial openness, which tends to differ according to the nominal exchange rate regime.

Following a different but related approach, Clower and Ito (2012) examine the determinants of episodes of "local non-stationarity" in current account dynamics using a Markov-Switching framework that allows for both mean-reverting and non-stationary regimes. Their results provide evidence that capital account openness and nominal exchange rate flexibility decrease the persistence of current account imbalances in the sense that these lower the probability of experiencing an episode of non-stationarity. Although my results do not directly address the adjustment of current account imbalances, they do provide support for the broader proposition that the nominal exchange rate regime influences the persistence of disequilibria. In particular, the results reported below show that real exchange rate misalignments are significantly less persistent in countries with more flexible exchange rate arrangements.

The remainder of the paper proceeds as follows. Section 2 provides a brief review of the standard empirical determinants of the real exchange rate and possible theoretical channels. Section 3 describes the dataset and econometric methodology used in the empirical analysis. Section 4 presents the benchmark results while section 5 discusses a series of extensions to the benchmark model, including asymmetries between undervaluations and overvaluations, as well as heterogeneity between exchange rate regimes. The final section provides concluding remarks.

2 The Equilibrium Real Exchange Rate and its Determinants

Purchasing power parity (PPP) is perhaps the oldest theory of exchange rate determination and states that after accounting for the domestic prices of goods and nominal exchange rates, all national currencies should have the same purchasing power.[6] This proposition is derived from the Law of One Price (LOP), which holds that in the absence of frictions such as transaction costs or other barriers to trade, international trade should cause all identical goods to trade for the same price across markets after converting into a common currency. Otherwise, it would be possible to profit through arbitrage and thus prices would eventually equalize across countries. Despite its appealing simplicity, empirical evidence suggests that PPP often fails to hold even as a long-run proposition (see, e.g., O'Connell, 1998; Engel, 2000; Pesaran, 2007).

A classic explanation for the failure of PPP is the relative productivity channel, which can be traced to Balassa (1964) and Samuelson (1964). This is the so-called Balassa-Samuelson effect, which in its simplest form predicts that countries with higher productivity in the tradable goods sector will tend to have more appreciated real exchange rates.[7] Intuitively, consider a small open economy with a tradable and non-tradable sector. Suppose further that PPP holds but only for tradable goods. Productivity growth in the tradable sector will tend to raise wages in both sectors and create upward pressure on prices. However, since the price of tradable goods is pinned down by the world market, this will lead to an increase in the relative price of non-tradables, or in other words a real exchange rate appreciation.

The Balassa-Samuelson effect has proven remarkably robust since its first test by Balassa (1964). Two examples of recent empirical confirmation of the Balassa-Samuelson effect are Lothian and Taylor (2008) and Chong et al. (2012). Employing a new semi-parametric approach, Chong et al. estimate the cointegrating relationship between the real exchange rate and productivity in a panel of 21 OECD countries at a quarterly

[6] The modern formulation of PPP is due to the Swedish economist Gustav Cassel in the early 20th century but elements of the doctrine can be traced to as far back as the Salamanca school in 16th century Spain.

[7] Some authors prefer to refer to this as the "Harrod-Balassa-Samuelson" effect due to early insights from Harrod (1933).

frequency. Their novel local projection approach makes it possible to purge the effects of short-run shocks and frictions and yields strong confirmation of the Balassa-Samuelson effect. Lothian and Taylor, use nearly two hundred years of data for the US, UK, and France to test the presence of the Balassa-Samuelson effect in an explicitly nonlinear framework that allows volatility shifts in the nominal exchange rate across monetary regimes. Their results suggest that the Balassa-Samuelson effect explains nearly 40 percent of variations in the sterling-dollar real exchange rate over the whole sample. Additional recent confirmation of the Balassa-Samuelson effect is provided by Bordo et al. (2014), who use historical data for 14 countries covering four distinct monetary regimes: the classical gold standard, the war and interwar years, Bretton Woods, and the post-Bretton Woods managed floats. They show that the traditional Balassa-Samuelson model cannot explain the small empirical effect of productivity on the real exchange rate or the substantial heterogeneity in its magnitude across monetary regimes. Modern versions of the model, including those that allow a role for product differentiation and terms of trade channels, fit the data much better. In particular, plausible shifts in structural parameters due to changes in monetary regimes can explain the historical variations in the Balassa-Samuelson effect and help reconcile discrepancies in estimates across countries. Bordo et al. conclude: "although the Balassa-Samuelson effect tends to vary across regimes, the evidence suggests that it is present, and in the long-run the real exchange rate is not constant but conditioned on relative income levels."

Another standard long-run determinant of the real exchange rate is the net foreign asset position. Interest in the impact of net foreign asset holdings on international relative prices dates back at least to the time of Keynes during the 1920's debate on the so-called *transfer problem*. Contemporary textbook models of an open economy predict a positive relationship between stocks of foreign assets and the relative price of non-tradable goods (e.g. Vegh, 2013, ch. 4). Since foreign assets represent a claim on tradable goods, an exogenous increase in foreign assets raises the supply of tradables and should lead to an increase in the relative price of non-tradables. Early empirical evidence of a positive association between net foreign asset stocks and the real exchange rate is provided by Gagnon (1996) and Lane and Milesi-Ferretti (2004). More recent studies that find a positive and significant effect include Ricci et al. (2013) and IMF (2013).

Changes in the terms of trade can also affect the real exchange rate and may help explain the long-run failure of PPP. In his 1930 *A Treatise on Money*, Keynes noted that a major problem with the theory of purchasing power parity is its neglect of the influence of the terms of trade on the real exchange rate, which "not only upsets the validity of [its] conclusions over the long period, but renders them even more deceptive over the short period..."[8] It is well understood in standard open economy macroeconomic models that improvements in the terms of trade can lead to a real appreciation of the exchange rate.[9]

Other potentially important determinants of the real exchange rate include government expenditure and demographic factors, most notably population growth. Government expenditure is expected to influence the real exchange rate through its effect on aggregate demand and the price level. It may also produce a real appreciation since public spending tends to be more concentrated on non-tradable goods and services (see, for example, De Gregorio and Wolf, 1994; Arellano and Larrain, 1996; Chinn, 2000). Although demographic factors have not received much attention in the equilibrium real exchange rate literature, higher fertility may appreciate the real exchange rate by raising consumption associated with child-rearing, which mainly consists of non-tradables. Rose et al. (2009) present a formal model and empirical evidence of this channel.

3 Data and Empirical Framework

In order to estimate the effect of capital controls on the persistence of real exchange rate disequilibria, I construct a dataset consisting of a balanced panel of 77 countries observed at a yearly frequency over the period 1980-2011. The sample is largely dictated by data availability and contains a mix of high, middle,

[8] Originally cited by Cashin et al. (2004).

[9] One such textbook treatment is Vegh (2013, ch. 4), which presents a simple intertemporal model of a small open endowment economy with three sectors: exportables, importables, and non-tradables. In this simple setup, wealth and intertemporal substitution effects both lead to a real appreciation following an improvement in the terms of trade and all that is required is for all goods to be normal.

Table 1: Sample description

Number of countries by World Bank group classifications			
Geographic Classification		Income Classification	
East Asia & Pacific	8	High Income: OECD	24
Europe & Central Asia	2	High Income: nonOECD	9
Industrial	25	Upper Middle Income	19
Latin America & Caribbean	19	Lower Middle Income	17
Middle East & North Africa	7	Low Income	8
South Asia	1		
Sub-Saharan Africa	14		

Table 2: Summary statistics

	Mean	Std. Dev.	Min	Max
Long-run Variables				
Log Real Effective Exchange Rate (RER)	4.668	0.336	3.545	7.685
Log PPP GDP per capita (LNY)	8.997	1.211	6.117	11.212
Net Foreign Assets / Imports (NFA)	-1.168	1.887	-15.853	5.989
Short-run Variables				
Log Commodity Terms of Trade (TOT)	4.735	0.351	3.413	6.193
Government Expenditure / GDP (GOV)	0.168	0.057	0.032	0.545
Population Growth (POP)	0.015	0.011	-0.018	0.082
Financial and/or Currency Crises ($CRISIS$)	0.081	0.274	0	1
Capital Control Indices				
Schindler Index – Overall (SCH)	0.276	0.338	0	1
Schindler Index – Inflows (SCH_{IN})	0.240	0.312	0	1
Schindler Index – Outflows (SCH_{OUT})	0.311	0.390	0	1
Schindler Index – Equity (SCH_{EQ})	0.277	0.350	0	1
Schindler Index – Collective Investments (SCH_{CI})	0.248	0.349	0	1
Inverse Chinn-Ito Financial Openness Index ($CHITO$)	0.509	0.369	0	1
Klein Episodic Index ($KLEIN$)	0.056	0.168	0	1

Note: Each variable was obtained from the following sources. $REER$: IMF International Financial Statistics. LNY: World Development Indicators. NFA: External Wealth of Nations Database (Lane and Milesi-Ferretti, 2007). TOT: IMF World Economic Outlook Database. GOV: World Development Indicators. POP: World Development Indicators. $CRISIS$: Broner et al. (2013). SCH_j: Fernández et al. (2014). $KLEIN$: Klein (2012). $CHITO$: Chinn and Ito (2008).

and low income countries. Table 1 provides a breakdown of the sample composition by geographic regions and country income groups.[10]

Most of the variables come from the IMF's *International Financial Statistics* (IFS) and the World Bank's *World Development Indicators* (WDI).[11] The dependent variable of interest is the natural logarithm of the real effective exchange rate (RER), which is an index constructed on the basis of a weighted average of each country's bilateral exchange rates vis-á-vis its trading partners deflated by its relative price level, where the weights reflect the importance of trade with each partner. The long-run variables included in the cointegrating relationship are the following: log PPP GDP per capita (LNY) , and net foreign assets divided by total imports (NFA). The short-run determinants of the real exchange rate are: log commodity terms of trade (TOT), government expenditure to GDP (GOV), annual population growth (POP), and a dummy variable for the advent of currency crises ($CRISIS$). Summary statistics are presented in Table 2.

Naturally, a key consideration is the appropriate measurement of capital controls. In the broadest sense, capital controls refer to any administrative or market-based restriction on cross-border financial flows. These

[10] The full set of countries included in the sample are listed in Appendix E.

[11] See the note for Table 2 for further details.

can range from outright prohibitions on the ownership of domestic assets by foreigners, to simple taxes on foreign exchange transactions or international borrowing. Capital controls may also be imposed either on a small subset of specific assets categories, or across the board, restricting or otherwise regulating international transactions in all types of financial instruments. A further distinction can be made between controls on capital inflows – that is, when foreigners acquire domestic assets – and controls on capital outflows – when domestic residents increase their holdings of foreign assets.[12]

Measures of capital controls fall into two broad categories: so-called de jure and de facto indexes. De jure indexes attempt to measure legal or regulatory barriers to international financial transactions while de facto measures, on the other hand, capture the actually existing level of financial integration in a given country, often by observing macroeconomic outcomes. The vast majority of de jure-type indexes are based on information contained in the IMF's Annual Report on Exchange Arrangements and Restrictions (AREAR), a yearly publication documenting changes in IMF member country laws and regulations governing international financial transactions. A major problem with de facto measures of capital controls is that they are potentially as much an endogenous outcome variable as they are an indicator of restrictions on capital flows. As such, de facto indexes are poorly suited for empirical studies where the aim is to ascertain the effect of a policy change since they do not actually measure changes in a government's *intention* to restrict flows.

For this reason, I will primarily use the de jure index developed by Schindler (2009), the so-called "Schindler index", which is based on detailed textual analysis of the AREAR and has also recently been updated to cover a larger number of countries and years by Fernández et al. (2014). The Schindler index is an average of the number of international transaction categories with any restrictions for a given year and country. Thus, the index ranges from zero – indicating that the country has no capital controls on any category – to one – when a country has controls on every transaction category. For example, between 1995 and 2011 Mexico's Schindler index averaged 0.5, suggesting that half of all transaction categories had some form of restriction during that time period. As noted by Quinn et al. (2011) in a thorough assessment of the most common measures of capital controls, the Schindler index is by far the most granular, covering a large range of disaggregated financial instruments and distinguishing between controls on inflows and outflows. There is, however, one major drawback of using the Schindler index that is worth noting. The AREAR only started publishing the detailed country reports on which the Schindler index is based starting in 1995 and as a result the index is only available in subsequent years. Moreover, because the index is based on textual analysis, its construction is labor intensive and does not include all IMF member countries. Thus, the sample used when investigating the short-run adjustment dynamics with capital controls is shorter, spanning 1995 to 2011, and includes less countries (43 compared to the 77 for the long-run analysis).

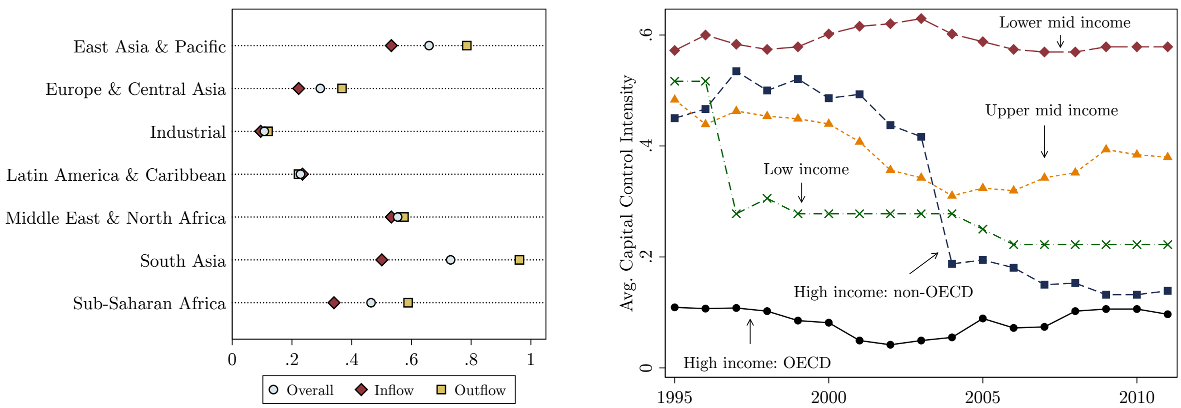
Figure 1 provides a broad overview of the relative prevalence of capital controls across regions and levels of development. As can be seen in Panel (a), large differences in the extent of capital account liberalization persist, on average, across country income groups. Perhaps not surprisingly, OECD countries had the most open capital accounts throughout the sample, with average restrictions on roughly 10 percent of instrument categories. In contrast, lower middle income countries had tighter capital controls on the books throughout the 1995-2011 period, with restrictions on roughly 60 percent of transaction categories. As a group, low income countries appear to have rapidly liberalized their capital accounts throughout the period, as did non-OECD high income countries. Large variation in the prevalence of capital controls is also evident across regions. South Asian economies had the tightest capital controls on average, with restrictions on roughly 75 percent of transaction categories. Latin American and Caribbean countries, in contrast, had nearly the loosest capital controls, second only to industrialized economies.

The order of integration of each variable was determined using the second generation panel unit root tests proposed by Pesaran (2007). Pesaran's Cross-Sectionally Augmented Dickey-Fuller (CADF) test tests the null hypothesis that all panels contain a unit root against the alternative that a fraction of panels are

[12] In addition to these distinctions, capital controls can also cover a wider and more subtle range of regulations governing capital inflows. For example, domestic monetary authorities may require firms to deposit a fraction of funds borrowed abroad in non-interest bearing accounts for a specified period of time. These "unremunerated reserve requirements" or URR, as they have come to be known, have been used most famously in Chile during the 1990s and in Colombia during the 2000s. Countries may also enforce so-called "minimum stay" requirements on foreign direct investment, barring the entry of short-term and potentially speculative investments.

Figure 1: Average intensity of capital controls by income group and region (1995-2011)

(a) By income groups **(b)** By geographic groups



Note: This figure reports the average value of the Schindler index for overall capital controls (SCH) broken up into the World Bank's geographic group (panel a) and country income (panel b) classifications. In panel (a), "Overall" refers to the Schindler index for both controls on capital inflows and outflows while "inflow" and "outflow", respectively, refer to the disaggregated indexes for restrictions on capital inflows and outflows.

stationary. Test results are reported in Table 3. The CADF tests fail to reject the null hypothesis that the level of the real exchange rate is non-stationary. This indicates, consistent with the literature discussed above, that the real exchange rate is likely $I(1)$ and therefore it will be treated as such in the empirical analysis that follows. The tests also suggest that LNY, NFA and TOT are first-difference stationary. As such, these are also treated as $I(1)$.

The variables were tested for cointegration using the panel error-correction tests proposed by Westerlund (2007) and implemented by Westerlund and Edgerton (2008). These tests are derived from a panel error-correction model that allows for heterogeneity in the error-correction dynamics, including panel-specific intercepts, trends, and slopes. The test statistics are based on the idea that if the series are cointegrated, the coefficient on the error-correction term should be significantly negative. Westerlund develops four alternative statics, two of which are constructed by averaging the estimated coefficients (G_α) and t-statistics (G_t) from each panel-specific error-correction term. The latter two are calculated by pooling observations across panels and estimating the error-correction term (P_α) and t-statistic (P_t).

Test results are shown in Table 4. Three of the four test statistics reject the null hypothesis of no cointegration for the model including RER, LNY, and NFA. Results for the model including the log terms of trade are inconclusive: only one of the three test statistics rejects the null of no cointegration. This is consistent with results presented by Cashin et al. (2004), who showed that the real exchange rate may only be cointegrated with the terms of trade in so-called commodity currency countries.[13] Given the inconclusive evidence of a cointegrating relationship, TOT is omitted from the specification of the long-run level.

The cointegrating relationship is estimated using the method of dynamic ordinary least-squares (DOLS) proposed by Saikkonen (1991). As Saikkonen shows, the cointegrating relationship can be consistently and efficiently estimated by OLS adding leads and lags of the first differenced cointegrated variables with Newey-West heteroskedasticity and autocorrelation consistent (HAC) standard errors. Because $RER_{i,t}$ is an index

[13] Cashin et al. (2004) uncover evidence of significant cross-country heterogeneity in the relationship between the real exchange rate and the terms of trade. They find significant cointegrating relationships between the real exchange rate and the terms of trade but only in around one third of the countries in the sample. This suggests that the long-run equilibrium exchange rate is only driven by the terms of trade in so-called "commodity currency" countries. However, for these commodity currencies, movements in the terms of trade explain a remarkably large amount of the variation in the real exchange rate. Their estimates imply that nearly 85% of real exchange rate variations are due to the terms of trade.

Table 3: Panel Unit Root Tests

	RER		*LNY*		*NFA*		*TOT*	
	Level	FD	Level	FD	Level	FD	Level	FD
t-bar statistic	-2.409	-2.981	-1.487	-2.635	-2.090	-3.017	-2.183	-3.344
Zt-bar statistic	-0.790	-6.496	8.070	-2.920	2.181	-6.245	1.462	-10.107
Critical Values for t-bar statistic								
1%	-2.650	-2.650	-2.650	-2.650	-2.680	-2.680	-2.650	-2.650
5%	-2.560	-2.560	-2.560	-2.560	-2.580	-2.580	-2.560	-2.560
10%	-2.510	-2.510	-2.510	-2.510	-2.530	-2.530	-2.510	-2.510
Countries	77	77	71	71	64	64	77	77
Years	34	33	33	32	32	31	32	31

Note: Pesaran's cross-sectionally augmented Dickey-Fuller (CADF) test was implemented in Stata by Lewandowski (2006). The CADF test tests the null hypothesis that all panels contain a unit root against the alternative that a fraction of panels are stationary. All reported test results consider the case with 2 lags, cross-sectional demeaning, and country-specific deterministic trends. The "t-bar statistic" diverges to negative infinity under the alternative hypothesis. Exact critical values for each combination of countries and years are provided. The "Zt-bar statistic" is distributed standard normal under the non-stationarity null hypothesis.

and does not contain information about the relative level of the real exchange rate, the model includes country fixed effects. The inclusion of country fixed effects also addresses potential omitted variable bias. Year dummies are also included to control for common time factors. The estimated long-run equilibrium equation is given by

$$RER_{i,t} = \gamma_i + \alpha_t + \beta x_{i,t} + \sum_{j=-\rho}^{\rho} \eta \Delta x_{i,t-j} + e_{i,t} \tag{1}$$

where γ_i and α_t are vectors of country and year fixed effects, respectively, $x_{i,t}$ is a vector of $I(1)$ variables cointegrated with $RER_{i,t}$, and the fourth term on the right hand side is the set of leads and lags of $\Delta x_{i,t}$. The error term $e_{i,t}$ captures short-run deviations from the long-run relationship and can be interpreted as the extent of real exchange rate disequilibria. A positive $e_{i,t}$ implies the real exchange rate is overvalued while a negative value implies an undervaluation. To estimate how fast deviations from the long-run equilibrium are eliminated, the estimated residuals, $\hat{e}_{i,t}$, are imposed on the error-correction model (ECM) in equation (2):

$$\Delta RER_{i,t} = \Theta_{i,t} \hat{e}_{i,t-1} + \alpha \Delta x_{i,t} + \beta z_{i,t} + u_{i,t} \tag{2}$$

where

$$\Theta_{i,t} = \theta_1 + \theta_2 K_{i,t} \tag{3}$$

The ECM is augmented with a vector of short-run stationary variables $z_{i,t}$. These include the annual change in the government expenditure to GDP ratio (ΔGOV), the log growth of the commodity terms of trade (ΔTOT), and a dummy for domestic and/or international financial crises ($CRISIS$). The coefficient $\Theta_{i,t}$ measures the speed of adjustment towards the long-run equilibrium and varies across both countries and years. Consistency between equations (1) and (2) requires $\Theta_{i,t} < 0$. Otherwise, $e_{i,t}$ would be non-stationary and therefore $RER_{i,t}$ and $x_{i,t}$ cannot be cointegrated. The term $K_{i,t}$ is a binary variable equal to one for countries with a "high" amount of restrictions on capital flows and equal to zero otherwise. Rather than allowing unlimited heterogeneity, the speed of adjustment is modeled as a function of a constant base-rate θ_1 and an additional term that depends on the intensity of capital controls. Hence, the speed of adjustment

Table 4: Westerlund Panel Cointegration Tests

	Panel-specific		Pooled	
	G_t	G_α	P_t	P_α
RER, LNY	-1.705	-3.979	-12.881	-3.995
	(0.000)	(0.380)	(0.000)	(0.000)
RER, LNY, NFA	-1.753	-3.754	-11.420	-3.631
	(0.003)	(0.999)	(0.001)	(0.034)
RER, LNY, NFA, TOT	-1.932	-2.463	-11.932	-4.204
	(0.043)	(1.000)	(0.117)	(0.533)

Note: This table reports the Z-values from the Westerlund (2007) panel cointegration tests. The null hypothesis is no cointegration. All tests consider the case with one lag and panel specific intercepts. These tests were implemented in Stata by Westerlund and Edgerton (2008). P-values are reported in parenthesis.

is captured by the marginal effect of $\hat{e}_{i,t}$ on $\Delta RER_{i,t}$:

$$\frac{\partial \Delta RER_{i,t}}{\partial \hat{e}_{i,t-1}} = \begin{cases} \theta_1 & \text{if } K_{i,t} = 0 \text{ (low capital controls)} \\ \theta_1 + \theta_2 & \text{if } K_{i,t} = 1 \text{ (high capital controls)} \end{cases} \tag{4}$$

If capital controls slow the speed of adjustment and cause disequilibria to persist for longer periods of time, then $\Theta_{i,t}$ should be smaller in absolute value when controls are present. This requires $\theta_1 < 0$, $\theta_2 > 0$, and $\Theta_{it} < 0$. In addition, another possibility is that real exchange rate misalignments may not disappear when capital controls are sufficiently tight. This case would entail $\Theta_{it} \geq 0$ and implies the absence of cointegration as long as strict controls remain in place. Intuitively, this would imply that countries with restrictive capital controls may be able to completely shield themselves against real exchange rate adjustment pressures.

As a first pass, "high" capital controls are defined as values of the Schindler index greater than or equal to 0.8. This corresponds roughly to the 90th percentile in this sample. Reassuringly, my results are robust to several different definitions, as well as to an alternative specification treating the intensity of capital controls as a semi-continuous variable. As an additional robustness check, I also carry out a sensitivity analysis using a simple threshold model that optimally chooses the threshold cutoff for high and low capital control regimes.[14]

Identification of the effect of capital controls may be confounded by the role of the nominal exchange rate regime. The exchange rate regime is an important omitted variables because it may directly influence the persistence of misalignments and is correlated with capital account policies. More rigid or managed regimes are by definition intended to dampen exchange rate fluctuations or to target specific levels of the exchange rate. As a result, more fixed regimes place the burden of adjustment towards the long-run equilibrium on changes in relative prices, which may be sluggish due to nominal rigidities. In addition, capital controls are often imposed in conjunction with more managed exchange rate arrangements. Therefore, if managed regimes tend to slow the adjustment of real misalignments, a naive model that omits the exchange rate regime will overstate the true effect of capital controls.

To address this concern, I rely on the *de facto* regime classifications constructed by Ilzetzki et al. (2010). This index categorizes exchange rate regimes by increasing degrees of flexibility, ranging from hard pegs and the absence of a national currency on one end of the spectrum, to freely floating on the other end. Following Eguren-Martín (2015), I use this data to construct dummy variables for relatively fixed and more flexible regimes. Details on the classification scheme and construction of this variable are provided in Appendix F. In general, flexible regimes tend to have fewer restrictions on capital flows than do fixed and managed regimes. Moreover, the average Schindler index is highest in managed regimes, regardless of whether the

[14]See Appendix B for more information.

Table 5: Long-run cointegrating relationship

Dependent Variable: RER_t	(1)	(2)	(3)
Log PPP GDP per capita (LNY)	0.213***	0.192**	0.182**
	(0.081)	(0.075)	(0.073)
Net foreign assets / imports (NFA)		0.028***	0.025**
		(0.011)	(0.011)
Log terms of trade (TOT)			0.089
			(0.063)
Error-Correction Model			
$\hat{e}_{t-1}$	-0.165***	-0.166***	-0.169***
	(0.021)	(0.022)	(0.022)
Country FE?	Yes	Yes	Yes
Time FE?	Yes	Yes	Yes
Observations	2109	1945	1945
Adj R^2	0.526	0.572	0.576

Note: The benchmark DOLS specification includes two leads and two lags of the differenced long-run explanatory variables. Results are robust to different lag lengths. The coefficient and standard error estimates for the leads and lags are not reported. Full results are available upon request. Robust HAC standard errors are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1

overall index or subindexes for restrictions on inflows or outflows are considered.[15]

The exchange rate regime indicators are then included in an extended version of the ECM in (2) allowing for additional heterogeneity in the speed of adjustment depending on the exchange rate regime. The error-correction term in the extended model is given by:

$$\Theta_{it} = \theta_1 + \theta_2 K_{i,t} + \theta_3 FLEX_{i,t} + \theta_4 FIX_{i,t} + \phi w_{i,t} \tag{5}$$

where $FLEX_{i,t}$ and $FIX_{i,t}$ denote, respectively, dummy variables for flexible and fixed exchange rate regimes, and managed or intermediate regimes are treated as the base category. The term $w_{i,t}$ refers to additional controls for the speed of adjustment, which include country income group dummies to control for differences in the level of development, and whether or not a country has adopted an inflation targeting regime.

Putting the pieces together, the empirical strategy is to estimate the long-run equilibrium relationship (1) and use the residuals to estimate the ECM in (2). To estimate the effect of differences in capital controls on the speed of adjustment, the different measures of capital control intensity are interacted with the lagged residuals. Therefore, a positive and statistically significant coefficient on the interaction term would confirm the hypothesis. The ECM is augmented with a lagged dependent variable to account for potential persistence in short-run real exchange rate movements and, in some specifications, I introduce a full set of country and time dummies to deal with unobservable short-run time-invariant and country-invariant factors. Since the introduction of a lagged dependent variable in a fixed-effects framework introduces dynamic panel bias (Nickell bias), these ECMs are estimated using two-step GMM.

4 Benchmark Results

The results for the benchmark equilibrium real exchange rate level regressions are presented in Table 5. I consider a variety of specifications for the long-run relationship, including a simple model where the long-run real exchange rate only depends on log GDP per capita. These results appear in column (1). The coefficient is positive, indicating that an increase in productivity leads to a real appreciation, and statistically significant at the one percent level. Its magnitude is also economically significant and consistent with the existing literature: a one percent increase in GDP per capita leads to roughly a fifth of a percent increase in the real exchange rate. The specification in column (2), which will serve as the baseline for the error-correction models estimated below, includes both LNY and NFA simultaneously. Both coefficients have the expected signs and are significant at standard significance levels. Consistent with the literature, a higher net foreign assets position has a statistically significant positive effect on the real exchange rate. In particular, a one standard deviation increase of NFA leads to a six percent real appreciation. Next, column (3) considers a model including the log commodity terms of trade. Although, consistent with the literature, the coefficient on the terms of trade is positive, the estimate is not statistically significant at standard confidence levels.

To compare with the results below, in each level specification I report the error-correction term for a simple ECM with homogenous adjustment dynamics. The speed of adjustment is roughly 0.17 in all three specifications, indicating that, for example, a one percent overvaluation produces a 0.17 percent offsetting depreciation the following year. These estimates are consistent with previous studies and, in particular, are very close to those reported by Ricci et al. (2013), who report an adjustment speed of 0.2. As a reference, these estimated adjustment speeds imply half-lives on average of roughly 3.7 years.[16]

The results for the ECM with heterogenous adjustment dynamics in (2) are shown in Table 6. The lagged residual $\hat{e}_{i,t-1}$ corresponds to the baseline level specification in column (2) of Table 5. As described above, the ECM is augmented with a lagged dependent variable and, in specifications (5) through (8), a full set of country and time dummies. Because the combination of a lagged dependent variable and fixed country effects introduces Nickell bias, these model are estimated using two-step GMM. In addition, several forms of additional slope heterogeneity in the error-correction dynamics are introduced, including unobservable differences across country income groups, nominal exchange rate arrangements, and inflation targeting regimes.

Our main coefficient of interest is for the interaction term between lagged real exchange rate misalignment and the binary measure of capital control intensity, $\hat{e}_{i,t-1} \cdot K_{i,t}$. The first thing to note is that the results appear to support the hypothesis that capital controls slow the speed of adjustment towards long-run equilibrium. Specifically, the interaction term has a positive and significant coefficient that is smaller in absolute value than the coefficient on $\hat{e}_{i,t-1}$. Controlling for additional forms of error-correction heterogeneity strengthens the results, increasing both the size of the point estimates and the probability of rejecting the null. Moreover, models (3), (4), (7), and (8) provide evidence that not only do capital controls impact the adjustment dynamics independently of the exchange rate regime, but that exchange rate flexibility matters in its own right. Specifically, the interaction term for flexibility ($\hat{e}_{i,t-1} \cdot FLEX_{i,t}$) is negative and statistically significant, indicating that real disequilibria are less persistent under flexible regimes. In contrast, inflation targeting does not appear to have a consistently significant impact on the speed of adjustment. These benchmark results are robust to using another popular measure of *de jure* capital mobility: the "Chinn-Ito" index from Chinn and Ito (2008).[17]

Next, Table 7 examines the effects of capital controls under different intensity thresholds of the Schindler index – columns (1)-(3) – as well as an alternative specification treating capital controls as a continuous variable – columns (4) and (5). The term K_{mean} refers to a dummy variable for countries with a Schindler index greater than or equal to the sample mean. In models with the continuous capital control intensity, it is necessary to also introduce the standalone effect of capital controls in order to properly identify the

[15] See Table A7 in Appendix F for summary statistics of the Schindler index by exchange rate regime classification.

[16] The adjustment speed half-life can be calculated as follows. Setting all short-term covariates equal to zero and assuming no further shocks to the real exchange rate, the half-life is given by: $HL = \ln(1/2)/\ln(1 + \hat{\theta})$, where $\hat{\theta}$ is the estimated error-correction term.

[17] See Table A4 in Appendix C.

Table 6: Benchmark Error-Correction Models

Dependent Variable: ΔRER_t					*Two-step GMM*			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\hat{e}_{t-1}$	-0.291***	-0.468***	-0.484***	-0.513***	-0.316***	-0.497***	-0.475***	-0.508***
	(0.037)	(0.067)	(0.087)	(0.090)	(0.037)	(0.070)	(0.078)	(0.082)
$\hat{e}_{t-1} \cdot K$	0.152*	0.197***	0.160**	0.170**	0.255***	0.301***	0.249***	0.246***
	(0.081)	(0.072)	(0.081)	(0.081)	(0.088)	(0.077)	(0.085)	(0.084)
$\hat{e}_{t-1} \cdot FIX$			0.075	0.111			-0.013	0.027
			(0.138)	(0.142)			(0.126)	(0.132)
$\hat{e}_{t-1} \cdot FLEX$			-0.195**	-0.213**			-0.183**	-0.198**
			(0.086)	(0.083)			(0.082)	(0.081)
$\hat{e}_{t-1} \cdot IT$				0.116*				0.118
				(0.067)				(0.083)
Controls								
Inc. Group Slope?	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Country FE?	No	No	No	No	Yes	Yes	Yes	Yes
Time FE?	No	No	No	No	Yes	Yes	Yes	Yes
Half-Lives (years)								
$K = 0$	2.014	1.098	1.046	0.963	1.822	1.008	1.075	0.978
$K = 1$	4.624	2.191	1.766	1.653	10.996	3.174	2.704	2.284
$H_0 : \hat{\Theta}_{it} = 0$ (p-value)	0.053	0.005	0.006	0.003	0.449	0.047	0.034	0.016
Observations	643	643	585	585	643	643	585	585
Adj R^2	0.267	0.281	0.305	0.308	0.282	0.294	0.315	0.316

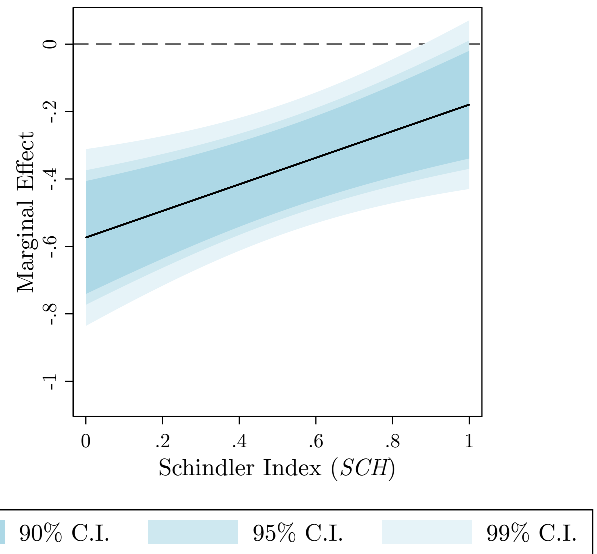
Note: Each ECM is estimated using the residuals from the DOLS specification (2) in Table 5. All specifications include first-differences of the long-run variables LNY and NFA and short-run covariates: GOV, TOT, $CRISIS$. The base groups for models including exchange rate regime and income group heterogeneity are, respectively, managed regimes and upper middle-income countries. The single lag-order was chosen using the AIC and BIC. The list of countries adopting inflation targeting regimes was taken from Roger (2010). Robust HAC standard errors are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

interaction effect. This inclusion also makes it possible to directly judge whether capital controls influence the real exchange rate through the error-correction dynamics or as separate short-run effects. Therefore, models (4) and (5) include both the interaction and the standalone Schindler index.

Incorporating varying degrees of capital control intensity does not alter our conclusions that controls increase in the persistence of disequilibria. The interaction terms for the presence of capital controls are positive and significant at standard confidence levels, though the point estimates and significance of the two different thresholds varies depending on whether or not the exchange rate regime is included. Turning to the continuous specification in columns (4) and (5), the interaction term is positive and significant at the 1 percent level. In contrast, the standalone short-run effect, though negative in sign, is not significant at standard confidence levels. Since the Schindler index ranges between 0 to 1, with a one indicating the highest intensity of capital controls, models (4) and (5) make it possible to judge the speed of error-correction given intermediate degrees of capital mobility. This is depicted graphically in Figure 2, which shows the speed of error-correction as a function of the Schindler index.

These estimates imply significant heterogeneity in the speed of adjustment across both countries and time. To illustrate these differences in speed, Table 6 also reports the estimated half-lives for the persistence of disequilibria. The real exchange rate converges to its equilibrium level at a very high speed in countries with relatively low control intensities. For instance, in model (2), which includes country income group heterogeneity, it takes roughly 1 year for half of a deviation to be eliminated in countries with no controls.

Figure 2: Error-Correction Speed as a Function of Capital Control Intensity



Note: This figure depicts the estimated error-correction speed as a function of the Schindler index (SCH): $\hat{\Theta}_{it} = \hat{\theta}_1 + \hat{\theta}_2 \cdot SCH_{it}$. The results correspond to the specification from model (5) in Table 7.

On the other hand, the half-life is more than twice as large, at 2.2 years, in countries with tight capital controls. The differences are even starker in models with country and year fixed effects: the half-life when controls are strict is as high as 3.2 years.

Taking full advantage of the granularity of the Schindler index, I also examine if controls on some types of financial instruments are more effective than others. To this end, I construct separate indicator variables for the presence of restrictions on a wide range of financial instrument categories distinguishing between restrictions on inflows and outflows. I then estimate separate ECMs with interactions for each instrument category. The full set of instrument categories considered are: equities, bonds, collective investments, direct investment, money market instruments, and financial credit. These results are reported in Tables A1 and A2 in Appendix A. In general terms, the results imply substantially different adjustment speeds depending on the type of restriction imposed. For instance, controls on equity flows appear to be particularly effective, whether inflow or outflow restrictions are considered. Controls on collective investments, on the other hand, only appear to be effective when inflows are restricted. Moreover, neither money market instruments nor financial credits appear to have statistically significant effects.

5 Extensions

This section extends the benchmark results in several directions. First, I extend the basic ECM to allow for asymmetries between real exchange rate overvaluations and undervaluations. Next, I take a closer look at the role of the exchange rate regime and how it may interact with capital account policies to influence the adjustment dynamics. Finally, I examine potential theoretical channels through which capital controls may influence the speed of adjustment.

5.1 Misalignment Asymmetries

While the models I have considered until now treat the effect of capital controls and real exchange rate adjustment as symmetric, capital controls may be expected to have an asymmetric effect depending on

Table 7: ECMs with Varying Capital Control Intensities

Dependent Variable: ΔRER_t	(1)	(2)	(3)	(4)	(5)
$\hat{e}_{t-1}$	-0.498***	-0.528***	-0.541***	-0.567***	-0.587***
	(0.078)	(0.071)	(0.083)	(0.084)	(0.100)
$\hat{e}_{t-1} \cdot K_{mean}$	0.175**	0.106	0.207**		
	(0.089)	(0.090)	(0.106)		
$\hat{e}_{t-1} \cdot K$		0.237***	0.105		
		(0.085)	(0.108)		
$\hat{e}_{t-1} \cdot SCH$				0.310**	0.361***
				(0.135)	(0.136)
SCH				-0.022	-0.015
				(0.019)	(0.020)
Controls					
Inc. Group Slope?	Yes	Yes	Yes	Yes	Yes
Exchange Rate Regime Slope?	No	No	Yes	No	Yes
Country FE?	Yes	Yes	Yes	Yes	Yes
Time FE?	Yes	Yes	Yes	Yes	Yes
Test H_0: $\hat{\Theta}_{it} = 0$ (p-value)					
Mean K_{it}/SCH_{it}	0.000	0.000	0.000	0.000	0.000
High K_{it}/SCH_{it}		0.058	0.028	0.014	0.016
Observations	643	643	585	643	585
Adj R^2	0.289	0.296	0.323	0.294	0.321

Note: Each ECM is estimated using the residuals from the DOLS specification (2) in Table 5. All specifications include first-differences of the long-run variables LNY and NFA and short-run covariates: GOV, $LTOT$, $CRISIS$. The base groups for models including exchange rate regime and income group heterogeneity are, respectively, managed regimes and upper middle-income countries. The single lag-order was chosen using the AIC and BIC. Robust HAC standard errors are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

whether the exchange rate is overvalued or undervalued. In particular, if a policymaker's aim is to maintain a competitive real exchange rate, successful capital controls should increase the persistence of undervaluations while facilitating the correction of overvaluations. Moreover, one should expect asymmetries between the effects of controls on capital inflows and outflows depending on the sign of the misalignment. I now turn to these issues and present evidence that capital controls do indeed have asymmetric effects and that these are most pronounced when the real exchange rate is undervalued.

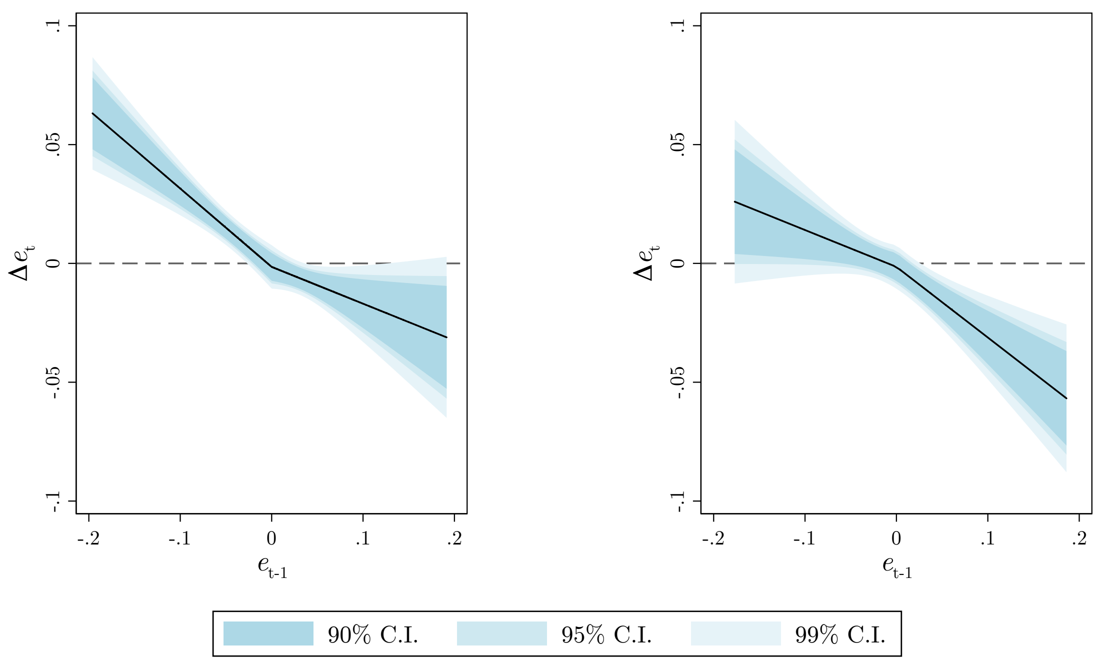
I start with a stripped down AR(1) model for real exchange rate disequilibria that incorporates a "kink" between overvaluations and undervaluations:

$$\Delta \hat{e}_{i,t} = (\theta_1 + \theta_2 K_{i,t}) \cdot \hat{e}_{i,t-1} + \mathbb{1}\{\hat{e}_{i,t-1} \geq 0\} \cdot (\theta_3 + \theta_4 K_{i,t}) \cdot \hat{e}_{i,t-1} + \gamma z_{i,t} + u_{i,t} \tag{6}$$

where $\mathbb{1}\{\hat{e}_{i,t-1} \geq 0\}$ is an indicator variable that is equal to one when the real exchange rate is overvalued and zero otherwise. This entails that, for example, the adjustment speed during undervaluations and with no capital controls is captured by the coefficient θ_1. The change in the adjustment speed when capital controls are introduced is captured by θ_2 and by $\theta_2 + \theta_4$ during undervaluations and overvaluations, respectively. The different adjustment dynamics based on the intensity of capital controls can be depicted graphically as a phase-diagram in $(\hat{e}_{i,t-1}, \Delta\hat{e}_{i,t})$-space. A dynamically stable equilibrium relationship requires a downward sloping curve, where steeper slopes correspond to faster adjustment dynamics. This is shown in Figure 3 for two cases: low capital controls ($K_{i,t} = 0$) and high capital controls ($K_{i,t} = 1$). Whenever $\hat{e}_{i,t-1} < 0$ and the

Figure 3: Misalignment Phase Diagram

(a) Low Capital Controls　　　**(b)** High Capital Controls



Note: This figure depicts the adjustment dynamics of real exchange rate disequilibria in a simple model of the form: $\Delta \hat{e}_{it} = (\theta_1 + \theta_2 K_{it}) \cdot \hat{e}_{it-1} + \mathbb{1}\{\hat{e}_{it-1} \geq 0\} \cdot (\theta_3 + \theta_4 K_{it}) \cdot \hat{e}_{it-1}$. The left panel (a) considers the low capital controls regime ($K_{it} = 0$) while the right panel (b) depicts the case with high capital controls ($K_{it} = 1$).

real exchange rate is undervalued, $\Delta \hat{e}_{i,t} > 0$ and thus the undervaluation is gradually eliminated.

As can be seen in Panel (a), when controls are absent the real exchange rate rapidly adjusts to eliminate undervaluations while overvaluations appear to be much more persistent, as indicated by the flatter curve to the right of zero. However, when capital controls are tightened in Panel (b), the adjustment curve is flatter in the $\hat{e}_{i,t-1} < 0$ region and therefore undervaluations are corrected more sluggishly. In contrast, the adjustment curve actually steepens in the $\hat{e}_{i,t-1} \geq 0$ region following the tightening of controls, which suggests that overvaluations become less persistent. These results therefore support the idea that capital controls can help countries "lean against the wind" by increasing the persistence of undervaluations while helping to avoid an overvaluation.

To assess the robustness of the finding and examine further asymmetries between controls on capital inflows and outflows, I introduce asymmetries into the error-correction model considered above in equation (2). Specifically, the error-correction term now takes the following form:

$$\Theta_{i,t} = \theta_1 + \theta_2 K_{i,t} + \mathbb{1}\{\hat{e}_{i,t-1} \geq 0\} \cdot (\theta_3 + \theta_4 K_{i,t}) + \phi w_{i,t} \tag{7}$$

where, as before, $w_{i,t}$ is a vector of control variables to account for additional forms of heterogeneity in the adjustment dynamics. In addition, I consider models with indicators for controls on capital inflows (K_{IN}) and capital outflows (K_{OUT}).

The results from the asymmetric ECMs are reported in Table 8. Starting with the stripped down model without controls, we can see in column (1) that there exists an asymmetry between overvaluations and undervaluations, as indicated by the positive and significant interaction term $\hat{e}_{t-1} \cdot \mathbb{1}\{\hat{e}_{t-1} \geq 0\}$. Consistent with the phase-diagram in Figure 3, we can also see evidence of significant differences in the effects of capital controls depending on whether the exchange rate is undervalued or overvalued. Indeed, the negative

Table 8: Error-Correction Models with Overvaluation Asymmetries

Dependent Variable: ΔRER_t	(1)	(2)	(3)	(4)	(5)	(6)	(7)
$\hat{e}_{t-1}$	-0.390***	-0.487***	-0.491***	-0.660***	-0.574***	-0.660***	-0.476***
	(0.048)	(0.075)	(0.084)	(0.100)	(0.102)	(0.106)	(0.108)
$\hat{e}_{t-1} \cdot \mathbb{1}\{\hat{e}_{t-1} \geq 0\}$	0.194**	0.321*	0.391**	0.487***	0.309*	0.406**	0.244
	(0.082)	(0.178)	(0.174)	(0.184)	(0.169)	(0.181)	(0.177)
$\hat{e}_{t-1} \cdot K$	0.340***	0.334***	0.345***				
	(0.097)	(0.085)	(0.100)				
$\hat{e}_{t-1} \cdot \mathbb{1}\{\hat{e}_{t-1} \geq 0\} \cdot K$	-0.428**	-0.413**	-0.424**				
	(0.185)	(0.193)	(0.199)				
$\hat{e}_{t-1} \cdot K_{IN}$				0.191***		0.190**	0.179**
				(0.067)		(0.077)	(0.083)
$\hat{e}_{t-1} \cdot \mathbb{1}\{\hat{e}_{t-1} \geq 0\} \cdot K_{IN}$				-0.130		-0.206*	-0.169
				(0.107)		(0.117)	(0.115)
$\hat{e}_{t-1} \cdot K_{OUT}$					0.105	0.001	-0.013
					(0.069)	(0.076)	(0.075)
$\hat{e}_{t-1} \cdot \mathbb{1}\{\hat{e}_{t-1} \geq 0\} \cdot K_{OUT}$					0.052	0.163	0.205*
					(0.102)	(0.115)	(0.110)
Controls							
Inc. Group Slope?	No	Yes	Yes	Yes	Yes	Yes	Yes
Time FE?	No	No	Yes	Yes	Yes	Yes	Yes
Short-Run Variables?	No	No	Yes	Yes	Yes	Yes	Yes
XR Regime Slope?	No	No	No	No	No	No	Yes
Observations	643	643	643	643	643	643	585
Adj R^2	0.250	0.262	0.323	0.315	0.315	0.316	0.336

Note: Each ECM is estimated using the residuals from the DOLS specification (2) in Table 5. All specifications include first-differences of the long-run variables LNY and NFA. The base groups for models including exchange rate regime and income group heterogeneity are, respectively, managed regimes and upper middle-income countries. The single lag-order was chosen using the AIC and BIC. $\mathbb{1}\{\cdot\}$ is an indicator function that is equal to one when the real exchange rate is overvalued and equal to zero otherwise. Robust HAC standard errors are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

and significant coefficient for the term $\hat{e}_{t-1} \cdot \mathbb{1}\{\hat{e}_{t-1} \geq 0\} \cdot K$ indicates that the asymmetry is statistically significant. Moreover, since $\hat{\theta}_2 + \hat{\theta}_4 < 0$, the point estimate for the effect of capital controls *during* on overvaluation is actually negative and controls may therefore accelerate the correction of overvaluations, though the combined estimate is not statistically significant at standard confidence levels.[18] Failure to reject the null that $\hat{\theta}_2 + \hat{\theta}_4 \neq 0$ suggests that capital controls may have no effect on the adjustment dynamics when the real exchange rate is overvalued, even if it has large and significant effects during an undervaluation. Models (2) and (3) in Table 8 add increasingly rich sets of controls, including income group slope heterogeneity, year fixed effects, and the short-run covariates ΔGOV and ΔTOT. As can be seen in the table, the estimates are qualitatively very similar to those reported for the stripped down model in column (1).

Next, models (4)-(7) introduce separate indicators for controls on capital inflows and outflows to examine if these have asymmetric effects. Controls on capital inflows appear to have a consistent and significant effect on the persistence of real disequilibria. What is less clear is whether or not they have asymmetric effects depending on the sign of the misalignment. Although the point estimates on $\hat{e}_{t-1} \cdot \mathbb{1}\{\hat{e}_{t-1} \geq 0\} \cdot K_{IN}$ are

[18] This can be ascertain through simple F-tests of the null hypothesis that $\hat{\theta}_2 + \hat{\theta}_4 \neq 0$. These results are not reported here but are available upon request.

negative, these are not statistically significant except for one specification, which is only significant at the 10 percent level. Results are similarly ambiguous for controls on capital outflows. In general, point estimates for the effects of controls on outflows are not statistically significant with the exception of model (7), where outflow restrictions appear to be associated with more persistent overvaluations. Overall, these results provide suggestive, if mixed, evidence of asymmetries between the impact of capital controls on inflows and outflows depending on whether the real exchange is overvalued or undervalued.

5.2 Exchange Rate Regime Heterogeneity

In this section, I take a closer look at the nominal exchange rate regime and test if capital controls have different effects on the persistence of misalignments across regimes. This type of heterogeneity could be important if, for example, capital controls enhance the effectiveness of foreign exchange market interventions or if policymakers across regimes have different policy objectives or preferences. To this end, I consider models where the error-correction term is given by:

$$\Theta_{i,t} = \theta_1 + \theta_2 SCH_{i,t} + \theta_3 SCH_{i,t} \cdot FIX_{i,t} + \theta_4 SCH_{i,t} \cdot FLEX_{i,t} + \theta_5 FIX_{i,t} + \theta_6 FLEX_{i,t} \tag{8}$$

where, as above, managed or intermediate exchange rate regimes are treated as the base group. Therefore, the coefficients θ_3 and θ_4 measure, respectively, the difference in the effects of capital controls under fixed and flexible exchange rate regimes relative to a managed regime. If the coefficients θ_3 and θ_4 are statistically significant, this can be interpreted as evidence of heterogeneity in the effect of controls between different exchange rate regimes. It is also worth noting that the specification in (8) includes the continuous measure of the Schindler index of capital control intensity, SCH, as in the models reported above in Table 7. This exercise is similar in spirit to estimates for the persistence of current account imbalances reported by Eguren-Martín (2015), which allow for heterogeneity in the effect of financial openness across exchange rate regimes.[19]

As Klein (2012) argues, it is potentially important to distinguish between permanent and episodic capital controls. According to Klein, this is because domestic financial institutions with experience in international financial markets may find it easier to evade short-term restrictions and taxes on capital flows, rendering episodic capital controls less effective than long-term ones. Another reason is that the episodic imposition of capital controls may reflect differences in the underlying motive or policymaker objective function behind the imposition of capital controls. That is, since episodic controls are by definition temporary, countries may use them to insulate themselves against short-term shocks as opposed to more long-term structural issues. Thus, as an additional robustness exercise, I consider the impact of episodic capital controls as defined by Klein.[20] The Klein index is simply an episodic counterpart of SCH that, in line with Klein's work, excludes permanent restrictions on capital flows.

Columns (1)-(3) in Table 9 report the results of extended ECMs with exchange rate regime heterogeneity. The first thing to notice is that, as above, capital controls appear to increase the persistence of real exchange rate misalignments under managed regimes, as indicated by the consistently positive and significant interaction terms. There is no evidence of significant differences in the effect of capital controls between managed or intermediate regimes and more fixed regimes, as indicated by the failure to reject the null hypothesis that the term $\hat{e}_{i,t-1} \cdot SCH \cdot FIX$ is equal to zero. I do, however, find evidence that capital controls may loose their effectiveness under more flexible nominal exchange rate regimes. The term $\hat{e}_{i,t-1} \cdot SCH \cdot FLEX$ is negative and statistically significant at standard confidence levels in models (2) and (3), which control for differences in the level of development and for inflation targeting regimes, and include country and year fixed effects, respectively. In fact, in these models the effect of capital controls under a flexible regime $(\hat{\theta}_2 + \hat{\theta}_4)$ is not statistically different from zero, as reported at the bottom of Table 9. This suggest that capital controls may lose their effectiveness when local authorities allow the exchange rate to float.

Results for specifications including episodic capital controls are reported in columns (4)-(6). The key takeaway from this exercise is that episodic capital controls also appear to slow the speed of adjustment in

[19]These results are robust to alternative specifications controlling for differences in the adjustment speed due to crises or sudden stops. These alternative specifications are available upon request.

[20]The episode dates and instruments covered correspond to those reported in Table A.1 in Klein (2012).

Table 9: Error-Correction Models – Exchange Rate Regime Heterogeneity

Dependent Variable: ΔRER_t				Episodic Controls		
	(1)	(2)	(3)	(4)	(5)	(6)
$\hat{e}_{t-1}$	-0.293***	-0.702***	-0.776***	-0.238***	-0.488***	-0.463***
	(0.073)	(0.115)	(0.102)	(0.049)	(0.103)	(0.091)
$\hat{e}_{t-1} \cdot SCH$	0.216**	0.421***	0.650***	0.293***	0.259**	0.376**
	(0.107)	(0.098)	(0.097)	(0.099)	(0.121)	(0.148)
$\hat{e}_{t-1} \cdot SCH \cdot FIX$	-0.106	-0.159	-0.430	-0.005	-0.413	-0.785
	(0.434)	(0.370)	(0.351)	(0.696)	(0.578)	(0.544)
$\hat{e}_{t-1} \cdot SCH \cdot FLEX$	-0.518	-0.801**	-1.010***	-0.535	-0.598	-0.813**
	(0.347)	(0.349)	(0.315)	(0.400)	(0.415)	(0.316)
SCH	-0.017***	-0.018***	-0.018	-0.013	-0.016	-0.028
	(0.006)	(0.006)	(0.018)	(0.011)	(0.012)	(0.019)
$SCH \cdot FIX$	0.017	0.018	0.029	0.012	0.023	0.084*
	(0.019)	(0.018)	(0.019)	(0.034)	(0.031)	(0.044)
$SCH \cdot FLEX$	-0.039	-0.043	-0.080*	-0.054	-0.058	-0.140***
	(0.035)	(0.033)	(0.044)	(0.043)	(0.042)	(0.048)
Controls						
Inc. Group Slope?	No	Yes	Yes	No	Yes	Yes
Inflation Target?	No	Yes	Yes	No	Yes	Yes
Country FE?	No	No	Yes	No	No	Yes
Year FE?	No	No	Yes	No	No	Yes
Effect of SCH						
$FIX = 1 \quad (\hat{\theta}_2 + \hat{\theta}_3)$	0.110	0.263	0.221	0.288	-0.154	-0.409
	(0.421)	(0.371)	(0.349)	(0.688)	(0.570)	(0.516)
$FLEX = 1 \ (\hat{\theta}_2 + \hat{\theta}_4)$	-0.302	-0.379	-0.359	-0.242	-0.339	-0.437
	(0.321)	(0.331)	(0.303)	(0.378)	(0.371)	(0.311)
Observations	585	585	585	585	585	585
Adj R^2	0.306	0.341	0.353	0.296	0.317	0.340

Note: Each ECM is estimated using the residuals from the DOLS specification (2) in Table 5. All specifications include first-differences of the long-run variables LNY and NFA and short-run covariates: GOV, $LTOT$, $CRISIS$. The base groups for the exchange rate regime and income group heterogeneity are, respectively, managed regimes and upper middle-income countries. Due to the combination of a lagged dependent variable and country fixed effects, models (3) and (6) are estimated with two-step GMM. Robust HAC standard errors are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

managed regimes, as indicated by the positive and significant coefficients on the interaction term $\hat{e}_{t-1} \cdot K$. As with non-episodic controls, there is no evidence of meaningful differences in the effect of controls between managed and fixed exchange rate regimes. In addition, the evidence regarding the effect of episodic capital controls under flexible regimes is more ambiguous than for their non-episodic counterparts. Indeed, only one of the reported specifications yields a statistically significant coefficients for the term $\hat{e}_{i,t-1} \cdot SCH \cdot FLEX$.

5.3 Nominal Exchange Rate Dynamics and Domestic Inflation

In this final section, I now present suggestive evidence that capital controls primarily affect the real exchange rate through their impact on nominal exchange rate dynamics. The analysis proceeds in two steps. First, I unpack how much of the variation in the persistence of disequilibria is due to nominal exchange rate volatility and how much is attributable to domestic inflation. Second, I then examine reduced form correlations between capital controls and periods of high exchange rate volatility and inflation through a simple Probit model controlling for differences in levels of development and exchange rate regimes.

In order to judge the relative impact of nominal exchange rate volatility and inflation on the persistence of real exchange rate disequilibria, the error-correction term in the benchmark ECM from equation (2) is modified as follows:

$$\Theta_{i,t} = \theta_1 + \theta_2 \cdot Channel_{i,t} \tag{9}$$

where $Channel_{i,t}$ is an indicator variable for high nominal exchange rate volatility or inflation. The idea is that if the correction of real exchange rate misalignments takes place through nominal exchange rate and relative price adjustments, countries with higher nominal exchange rate volatility and inflation should exhibit faster error correction. The coefficient θ_2 is therefore expected to have a negative sign. Periods of high nominal exchange rate volatility and inflation are identified using a five year rolling window. Countries are categorized as "high" if, for example, average inflation over the 5-year period exceeds the sample's 75th percentile.

Results are reported for each channel in the upper panel of Table 10. Models (1)-(3) consider error-correction heterogeneity due to high nominal volatility, while models (4)-(6) consider heterogeneity due to high inflation periods. As expected, the coefficient $\hat{\theta}_2$ is negative for both channels, indicating that greater exchange rate volatility and high inflation are associated with faster error-correction. Nevertheless, there are large differences in the magnitude of the point estimates and these are not statistically significant for periods of high inflation. Although not definitive, this provides some evidence that the correction of real exchange rate disequilibria occurs largely through nominal exchange rate adjustment.

The next step is to examine if capital controls are associated with either of these two channels. This is accomplished by way of a standard Probit model with the indicator for either high nominal volatility or high inflation as the dependent variable and capital controls as the explanatory variable:

$$\Pr(Channel_{i,t}|K_{i,t-5}) = \Phi(\beta_0 + \beta_1 K_{i,t-5}) \tag{10}$$

The indicator for the presence of capital controls, $K_{i,t-5}$, is appropriately lagged five years to match the timing of the 5-year rolling window of the $Channel_{i,t}$ variable.

The marginal effects from the Probit models are reported in the bottom panel of Table 10, where each specification corresponds in the included controls and channel to the ECM in the panel above. As can be seen in the table, lagged capital controls have a strong negative correlation with nominal exchange rate volatility. The results imply that the introduction of capital controls are associated with a reduced probability of experiencing an episode of high nominal volatility of between -0.125 and -0.222. In contrast, there is no clear association between periods of high inflation and lagged capital controls. Although clearly these results should not be interpreted causally, they nevertheless provide suggestive evidence that capital controls slow the adjustment of real exchange rate disequilibria through nominal exchange rate dynamics.

Table 10: ECMs with Channel Heterogeneity and Probit Analysis

Channel:	*Nominal Exchange Rate Volatility*			*Inflation*		
	(1)	(2)	(3)	(4)	(5)	(6)
Error-Correction Model						
$\hat{e}_{t-1}$	-0.182***	-0.271***	-0.292***	-0.235***	-0.383***	-0.372***
	(0.041)	(0.084)	(0.073)	(0.037)	(0.084)	(0.099)
$\hat{e}_{t-1} \cdot Channel$	-0.285***	-0.256*	-0.270**	-0.128	-0.016	-0.108
	(0.084)	(0.144)	(0.137)	(0.098)	(0.096)	(0.085)
Observations	628	628	570	628	628	570
Adj R^2	0.261	0.260	0.285	0.236	0.243	0.268
Channel Probit Model (MFX)						
K_{t-5}	-0.194***	-0.222***	-0.125***	-0.020	-0.060	0.032
	(0.039)	(0.030)	(0.019)	(0.051)	(0.050)	(0.049)
Observations	657	610	469	657	610	469
Controls						
Inc. Group Slope?	No	Yes	Yes	No	Yes	Yes
XR Regime Slope?	No	No	Yes	No	Yes	Yes

Note: Each ECM is estimated using the residuals from the DOLS specification (2) in Table 5. All specifications include first-differences of the long-run variables LNY and NFA. The base groups for models including exchange rate regime and income group heterogeneity are, respectively, managed regimes and upper middle-income countries. *Channel* refers to indicator variables for the periods of high exchange rate volatility or inflation over a 5-year window. The Probit results are reported as marginal effects. Robust HAC standard errors are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

6 Conclusion

This paper has examined the relationship between capital controls and the real exchange rate. The consensus among empirical studies on the effects of capital controls is that these enable domestic authorities to maintain an independent monetary policy and shield countries from short-term, speculative flows. The evidence is far less conclusive when it comes to limiting the overall volume of flows and influencing the real exchange rate. Previous studies, however, have largely overlooked the long-run determinants of the real exchange rate and are therefore misspecified. Taking the determinants of the real exchange rate seriously, I have presented evidence that capital controls may have very dramatic effects on real exchange rate dynamics, especially if controls are sufficiently strict. Specifically, controls increase the persistence of real exchange rate misalignments, and these effects appear to be strongest during an undervaluation.

My results also lend support to the proposition that flexible exchange rate regimes accelerate the adjustment of misalignments relative to managed and fixed regimes. In addition, I find some evidence that controls may lose traction under flexible regimes and that controls appear to operate primarily through the nominal exchange rate, as opposed to through relative price adjustment. Together, these results suggest that capital controls should be regarded as complements to traditional forms of exchange rate policy to the extent that they enhance the ability of managed and fixed exchange rate regimes to slow the adjustment of the real exchange rate towards its long-run equilibrium.

The broader lesson to take from this study is that capital controls are an effective policy tool for managing the real exchange rate. In other words, controls can help achieve policy objectives in addition to the macro-prudential concerns stressed by the recent literature. In particular, capital controls can be of use to countries seeking to deliberatively maintain a real exchange rate undervaluation. Nevertheless, strictly speaking, the empirical results presented above to do not explain how an undervaluation is initially achieved but rather suggest that the real exchange rate, once already undervalued, will take longer to converge to its long-run level. How the undervaluation is originally achieved and how this affects the short-run dynamics of the real exchange rate requires further research.

References

Laura Alfaro, Anusha Chari, and Fabio Kanczuk. The real effects of capital controls: financial constraints, exporters, and firm investment. NBER Working Paper 20726, December 2014.

Soledad Arellano and Felipe Larrain. Tipo de cambio real y gasto publico: Un modelo econometrico para chile. *Cuadernos de Economia*, 33(98):47–75, April 1996.

Chikako Baba and Annamaria Kokenyne. Effectiveness of capital controls in selected emerging markets in the 2000s. *IMF Working Paper*, 11(281), December 2011.

Bela Balassa. The purchasing power parity doctrine: a reappraisal. *Journal of Political Economy*, 72(6): 584–596, 1964.

Tamim Bayoumi, Hamid Faruqee, and Jaewoo Lee. A fair exchange? theory and practice of calculating equilibrium exchange rates. IMF Working Papers 05/229, 2005.

Michael D. Bordo, Ehsan U. Choudhri, Giorgio Fazio, and Ronald MacDonald. The real exchange rate in the long run: Balassa-samuelson effects reconsidered. NBER Working Paper 20228, 2014.

Fernando Broner, Tatiana Didier, Aitor Erce, and Sergio L. Schmukler. Gross capital flows: dynamics and crises. *Journal of Monetary Economics*, 60(1):113–133, 2013.

Paul Cashin, Luis F. Céspedes, and Ratna Sahay. Commodity currencies and the real exchange rate. *Journal of Development Economics*, 75(1):239–268, October 2004.

Menzie Chinn and Louis Johnston. Real exchange rate levels, productivity and demand shocks: evidence from a panel of 14 countries. *National Bureau of Economic Research*, Working Paper 5907, 1996.

Menzie D. Chinn. The usual suspects? productivity and demand shocks and asia-pacific real exchange rates. *Review of International Economics*, 8(1):20–43, 2000.

Menzie D. Chinn and Hiro Ito. A new measure of financial openness. *Journal of Comparative Policy Analysis*, 10(3):309–322, September 2008.

Menzie D. Chinn and Shang-Jin Wei. A faith-based initiative meets the evidence: Does a flexible exchange rate regime really facilitate current account adjustment? *Review of Economics and Statistics*, 95(1): 168–184, 2013.

Yanping Chong, Òscar Jordà, and Alan M. Taylor. The harrod-balassa-samuelson hypothesis: real exchange rates and their long-run equilibrium. *International Economic Review*, 53(2):609–634, 2012.

Erica Clower and Hiro Ito. The persistence of current account balances and its determinants: The implications for global rebalancing. Asian Development Bank Institute, 2012.

Scott Davis and Ignacio Presno. Capital controls as an instrument of monetary policy. Technical report, Federal Reserve Bank of Dallas, 2014. URL `https://www.dallasfed.org/assets/documents/institute/wpapers/2014/0171.pdf`.

José De Gregorio and Holger C. Wolf. Terms of trade, productivity, and the real exchange rate. NBER Working Paper 4807, 1994.

José De Gregorio, Sebastian Edwards, and Rodrigo O. Valdes. Controls on capital inflows: do they work? *Journal of Development Economics*, 63(1):59–83, 2000.

Sebastian Edwards. How effective are capital controls? *Journal of Economic Perspectives*, 13(4):65–84, 1999.

Sebastian Edwards and Roberto Rigobon. Capital controls on inflows, exchange rate volatility and external vulnerability. *Journal of International Economics*, 78(2):256–267, July 2009. ISSN 00221996. doi: 10.1016/j.jinteco.2009.04.005. URL `http://linkinghub.elsevier.com/retrieve/pii/S0022199609000518`.

Fernando Eguren-Martín. Exchange rate regimes and current account adjustment: an empirical investigation. *Bank of Englad Staff Working Papers*, (544), 2015.

Charles Engel. Long-run PPP may not hold after all. *Journal of International Economics*, 57(2):243–273, 2000.

Charles Engel. Macroprudential policy under high capital mobility: Policy implications from an academic perspective. Working Paper, February 2015.

Bilge Erten and José A. Ocampo. Macroeconomic effects of capital accont regulations. *IMF Economic Review*, 65(2):193–240, June 2017.

Emmanuel Farhi and Iván Werning. A theory of macroprudential policies in the presence of nominal rigidities. NBER Working Paper 19313, Augest 2013.

Andrés Fernández, Alessandro Rebucci, and Martín Uribe. Are capital controls countercyclical? Manuscript, December 2014.

Kristin Forbes, Marcel Fratzscher, and Roland Straub. Capital flow management measures: what are they good for? NBER Working Paper 20860, January 2015.

Kristin J. Forbes. One cost of the chilean capital controls: Increased financial constraints for smalles traded firms. Technical report, National Bureau of Economic Research, 2003. URL `http://www.nber.org/papers/w9777`.

Joseph Gagnon. Net foreign assets and equilibrium exchange rates: Panel evidence. International Finance Discussion Paper No. 574, December 1996.

Francisco Gallego, Leonardo Hernandez, and Klaus Schmidt-Hebbel. Capital controls in chile: were they effective? In Leonardo Hernandez and Klaus Schmidt-Hebbel, editors, *Banking, Financial Integration, and Crises*. Santiago: Central Bank of Chile, Santiago, 2002.

Roy Harrod. *International Economics*. Nisbet and Cambridge University Press, London, 1933.

Jonathan Heathcote. On the desirability of capital controls. NBER Working Paper 21898, January 2016.

Sabine Herrmann. Do we really know that flexible exchange rates facilitate current account adjustment? some new empirical evidence for cee countries. Deutsche Bundesbank - Economic Research Centre, 2009.

Ethan O. Ilzetzki, Carmen M. Reinhart, and Kenneth S. Rogoff. Exchange rate arrangement entering the 21st century: Which anchor will hold? *Unpublished*, 2010.

IMF. The liberalization and management of capital flows: an institutional view, November 2012.

IMF. External balance assessment (EBA) methodology: Technical background. Technical report, International Monetary Fund, Washington, D.C., June 2013.

Olivier Jeanne. Capital account policies and the real exchange rate. NBER Working Paper 18404, 2012.

Olivier Jeanne and Anton Korinek. Excessive volatility in capital flows: A pigouvian taxation approach. *American Economic Review Papers & Proceedings*, 100(2):403–407, May 2010.

Yothin Jinjarak, Ilan Noy, and Huanhuan Zheng. Capital controls in Brazil – stemming a tide with a signal? NBER Working Paper 19205, July 2013.

Michael W. Klein. Capital controls: Gates versus walls. *Brookings Papers on Economic Activity*, 2012.

Anton Korinek. Capital flows, crises and externalities. In Franklin Allen, Masahiko Aoki, Nobuhiro Kiyotaki, Roger Gordon, Joseph E. Stiglitz, and Jean-Paul Fitoussi, editors, *The Global Macro Economy and Finance*. Palgrave Macmillan, 2012.

Anton Korinek and Damiano Sandri. Capital controls or macroprudential regulation? NBER Working Paper 20805, December 2014.

Anton Korinek and Luis Serven. Undervaluation through reserve accumulation: Static losses, dynamic gains. *World Bank Policy Research Working Papers*, (5250), March 2010.

Philip R. Lane and Gian M. Milesi-Ferretti. The transfer problem revisited: Net foreign assets and real exchange rates. *Review of economics and statistics*, 86(4):841–857, November 2004.

Philip R. Lane and Gian M. Milesi-Ferretti. The external wealth of nations mark ii: revised and extended estimates of foreign assets and liabilities, 1970-2004. *Journal of International Economics*, 73(2):223–250, November 2007.

Eduardo Levy-Yeyati, Sergio L. Schmukler, and Neeltje Van Horen. Crises, capital controls, and financial integration. *World Bank Policy Research Working Papers*, (4770), November 2008.

Piotr Lewandowski. Pescadf: Stata module to perform pesaran's cadf panel unit root test in presence of cross section dependence. Statistical Software Components, Boston College Department of Economics, Jun 2006.

Zheng Liu and Mark M. Spiegel. Optimal monetary policy and capital account restrictions in a small open economy. Federal Reserve of San Francisco Working Paper, February 2015.

Guido Lorenzoni. Inefficient credit booms. *Review of Economic Studies*, 75:809–833, 2008.

James R. Lothian and Mark P. Taylor. Real exchange rates over the past two centuries: How important is the harrod-balassa-samuelson effect? *The Economic Journal*, 118(532):1742–1763, 2008.

Ronald MacDonald. *Exchange Rate Economics*. Routledge, New York, 2007.

Nicolas E. Magud, Carmen M. Reinhart, and Kenneth Rogoff. Capital controls: myth and reality - a portfolio balance approach. NBER Working Paper 16805, 2011.

Paul GJ O'Connell. The overvaluation of purchasing power parity. *Journal of international economics*, 44 (1):1–19, 1998. URL `http://www.sciencedirect.com/science/article/pii/S0022199697000172`.

Jonathan D Ostry, Atish R Ghosh, Marcos Chamon, and Mahvash S Qureshi. Capital controls: When and why? *IMF Economic Review*, 59(3):562–580, August 2011a.

Jonathan D. Ostry, Atish R. Ghosh, Karl Habermeier, Luc Laeven, Marcos Chamon, Mahvash S. Qureshi, and Annamaria Kokenyne. Managing capital inflows: What tools to use? *IMF Staff Discussion Note*, 2011b.

Mr Jonathan David Ostry, Mahvash Saeed Qureshi, Mr Karl Friedrich Habermeier, Dennis BS Reinhardt, Mr Marcos Chamon, and Mr Atish R. Ghosh. *Capital inflows: The role of controls*. International Monetary Fund, 2010.

M. Hashem Pesaran. A simple panel unit root test in the presence of cross-section dependence. *Journal of Applied Econometrics*, 22(2):265–312, March 2007.

Dennis Quinn, Martin Schindler, and A Maria Toyoda. Assessing measures of financial openness and integration. *IMF Economic Review*, 59(3):488–522, August 2011.

Martin Rapetti, Peter Skott, and Arslan Razmi. The real exchange rate and economic growth: are developing countries different? *International Review of Applied Economics*, 26(6):735–53, 2012.

Arslan Razmi, Martin Rapetti, and Peter Skott. The real exchange rate and economic development. *Structural Change and Economic Dynamics*, 23(2):151–169, 2012.

Luca Antonio Ricci, Gian M. Milesi-Ferretti, and Jaewoo Lee. Real exchange rates and fundamentals: A cross-country perspective. *Journal of Money, Credit and Banking*, 45(5):845–865, 2013.

Dani Rodrik. The real exchange rate and economic growth: theory and evidence. *Brookings Papers on Economic Activity*, 39(2):365–439, 2008.

Scott Roger. Inflation targeting turns 20. *Finance and Development*, 47(1):46–49, 2010.

Andrew K. Rose, Saktiandi Supaat, and Jacob Braude. Fertility and the real exchange rate. *Canadian Journal of Economics*, 42(2):496–518, 2009.

Petti Saikkonen. Asymptotic efficient estimation of cointegration regressions. *Econometric Theory*, 7(1): 1–21, 1991.

Paul A. Samuelson. Theoretical notes on trade problems. *The Review of Economics and Statistics*, 46(2): 145, May 1964.

Lucio Sarno and Mark P. Taylor. *The Economics of Exchange Rates.* Cambridge University Press, New York, 2002.

Martin Schindler. Measuring financial integration: A new data set. *IMF Staff Papers*, 56(1), 2009.

Stephanie Schmitt-Grohé and Martín Uribe. Downward nominal wage rigidity, currency pegs, and involuntary unemployment. *Journal of Political Economy*, Forthcoming.

Salvador Valdés-Prieto and Marcelo Soto. The effectiveness of capital controls: theory and evidence from chile. *Empirica*, 25(2):133–164, 1998.

Carlos A. Vegh. *Open Economy Macroeconomics in Developing Countries.* MIT Press, August 2013.

Joakim Westerlund. Testing for error correction in panel data. *Oxford Bulletin of Economics and Statistics*, 69(6):709–748, December 2007.

Joakim Westerlund and David L. Edgerton. A simple test for cointegration in dependent panels with structural breaks. *Oxford Bulletin of Economics and Statistics*, 70(5):665–704, 2008.

A ECMs with Additional Capital Control Measures

This appendix presents additional error-correction models with finer breakdowns of capital control transaction categories. Tables A1 and A2 report the error-correction term and its interaction with various measures of capital controls. The ECM specification is the same as the benchmark model reported in column (2) of Table (6). Each specification in Tables A1 and A2 features a different indicator variable constructed from the Schindler subindexes corresponding to the presence of restrictions on a financial instrument category j. The full list of instruments reported are: equities, bonds, collective investment instruments, foreign direct investment, money market instruments, and financial credit.

Table A1: Error-Correction Models by Various Capital Control Measures

Dependent Variable: ΔRER_t									
K-control measure j:	Equity			Bonds			Collective Investment		
		IN	OUT		IN	OUT		IN	OUT
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$\hat{e}_{t-1}$	-0.518***	-0.479***	-0.548***	-0.511***	-0.422***	-0.574***	-0.446***	-0.438***	-0.417***
	(0.100)	(0.085)	(0.101)	(0.107)	(0.086)	(0.118)	(0.104)	(0.064)	(0.104)
$\hat{e}_{t-1} \cdot K_j$	0.178**	0.161**	0.211***	0.118*	0.035	0.198**	0.109	0.250**	0.077
	(0.075)	(0.077)	(0.079)	(0.062)	(0.095)	(0.084)	(0.086)	(0.104)	(0.090)
Observations	643	643	643	558	558	558	643	643	643
Adj R^2	0.251	0.249	0.253	0.271	0.268	0.278	0.245	0.261	0.244

Table A2: Error-Correction Models by Various Capital Control Measures (continued)

Dependent Variable: ΔRER_t									
K-control measure j:	Direct Investment			Money Market			Financial Credit		
		IN	*OUT*		*IN*	*OUT*		*IN*	*OUT*
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$\hat{e}_{t-1}$	-0.436***	-0.363***	-0.397***	-0.408***	-0.408***	-0.406***	-0.384***	-0.367***	-0.341***
	(0.069)	(0.083)	(0.098)	(0.067)	(0.064)	(0.067)	(0.060)	(0.060)	(0.088)
$\hat{e}_{t-1} \cdot K_j$	0.143**	0.027	0.090	0.110	0.152	0.107	0.053	0.025	-0.030
	(0.060)	(0.074)	(0.092)	(0.074)	(0.111)	(0.074)	(0.077)	(0.089)	(0.086)
Observations	643	643	642	643	643	643	643	643	643
Adj R^2	0.251	0.242	0.244	0.247	0.249	0.247	0.243	0.242	0.242

B Sensitivity Analysis Using Threshold ECMs

This appendix briefly presents the results from a series of threshold error-correction models intended to assess the sensitivity of my benchmark results to alternative thresholds for regimes with "high" restrictions on capital mobility. In other words, instead of *a priori* specifying the threshold for the Schindler index, the idea is to determine the threshold with a data-driven criteria. Specifically, these models choose the "optimal" threshold, denoted as k^*, as the value of the Schindler index that minimizes the root mean squared error (RMSE). This procedure is illustrated for several different specifications in Figure A1. For each specification, I show the estimated interaction term $\hat{\theta}_2$ from model (2) for a wide range of thresholds, as well as the RMSE corresponding to that threshold value. The RMSE-minimizing threshold is indicated by a red line.

Table A3 reports the results from ECMs with optimally chosen thresholds for several different measures of capital controls. As can be seen in the table, the estimated effect of capital controls on the persistence of real exchange rate misalignments is both qualitatively and quantitatively similar to the benchmark specifications presented in Section 4 in the main text. Table A3 also reports the value of the RMSE-minimizing threshold, k^*, which in general differs depending on the capital control measure considered.

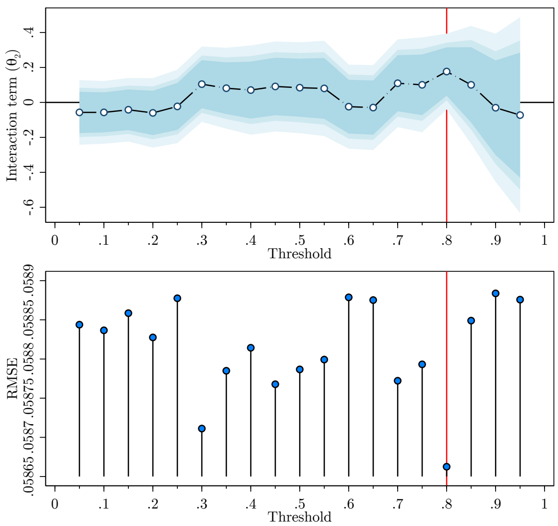
Table A3: Threshold Error-Correction Models

Dependent Variable: ΔRER_t								
K-control measure j:	Overall	Inflow	Outflow	Chinn-Ito	Equity	Bonds	Col. Invest.	FDI
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\hat{e}_{t-1}$	-0.470***	-0.393***	-0.455***	-0.266***	-0.377***	-0.411**	-0.385***	-0.441***
	(0.129)	(0.114)	(0.143)	(0.064)	(0.134)	(0.188)	(0.129)	(0.112)
$\hat{e}_{t-1} \cdot K_j$	0.234***	0.350***	0.139	0.223**	0.142*	0.174	0.348***	0.194**
	(0.078)	(0.129)	(0.101)	(0.098)	(0.084)	(0.108)	(0.124)	(0.083)
Controls								
Inc. Group Slope?	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country FE?	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time FE?	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Optimal Threshold (k^*)	0.3	0.625	0.25	0.8	0.65	0.65	0.65	0.275
Observations	643	643	643	1185	643	558	643	643
Adj R^2	0.261	0.261	0.251	0.226	0.244	0.246	0.257	0.254

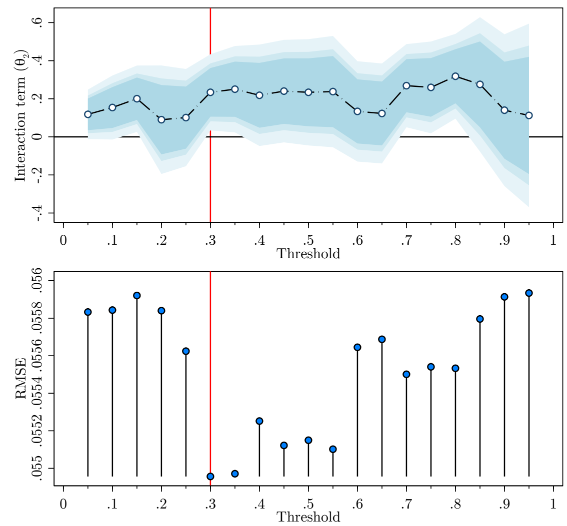
Note: Each ECM is estimated using the residuals from the DOLS specification (2) in Table 5. All specifications include first-differences of the long-run variables LNY and NFA. The optimally chosen threshold to construct the capital control indicator variable K_j is reported as k^*. All models are estimated using two-step GMM. The base income group is upper middle-income. Robust HAC standard errors are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Figure A1: Capital control threshold sensitivity analysis

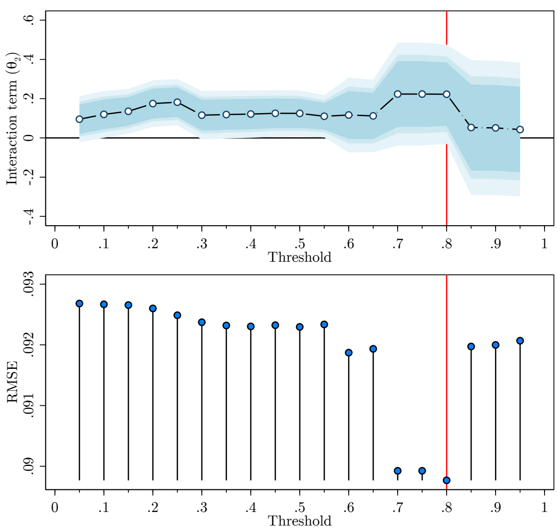
(a) *Model 1:* Benchmark

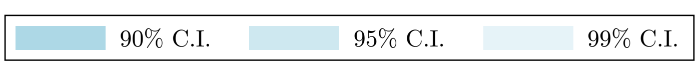


(b) *Model 2:* GMM, inc. group slope



(c) *Model 3:* Chinn-Ito index





Note: This figure reports the estimated interaction term ($\hat{\theta}_2$) and root mean squared error (RMSE) for a wide range of alternative threshold values of capital control intensity and three alternative model specifications. The optimal threshold k^* that minimizes the RMSE is shown for each model in red. *Model 1* considers the benchmark model with no additional slope heterogeneity. *Model 2* includes country income group slope heterogeneity, year and country fixed effects, and is estimated with two-step GMM. *Model 3* is the same as *Model 2* but uses the Chinn-Ito index.

C ECMs with Chinn-Ito Index

This appendix reports the results from estimating the benchmark ECMs using the Chinn-Ito index of financial openness instead of the Schindler index. To ensure comparability with the benchmark results, I normalize and use the inverse of the Chinn-Ito index so that it ranges from zero to one, where a greater value corresponds to less financial openness. As in the benchmark models, I define a regime of "high" capital controls as those with a normalized Chinn-Ito index exceeding 0.8.

Table A4: Benchmark Error-Correction Models With Chinn-Ito Index

Dependent Variable: ΔRER_t								
					Two-step GMM			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\hat{e}_{t-1}$	-0.280***	-0.397***	-0.403***	-0.404***	-0.332***	-0.449***	-0.452***	-0.450***
	(0.073)	(0.093)	(0.093)	(0.096)	(0.074)	(0.091)	(0.106)	(0.111)
$\hat{e}_{t-1} \cdot K$	0.144	0.201**	0.241**	0.242**	0.189*	0.244**	0.281***	0.279**
	(0.098)	(0.096)	(0.099)	(0.105)	(0.100)	(0.099)	(0.109)	(0.114)
$\hat{e}_{t-1} \cdot FIX$			0.076	0.077			0.058	0.055
			(0.122)	(0.126)			(0.124)	(0.129)
$\hat{e}_{t-1} \cdot FLEX$			-0.038	-0.038			-0.037	-0.037
			(0.046)	(0.046)			(0.042)	(0.043)
$\hat{e}_{t-1} \cdot IT$				0.007				-0.026
				(0.080)				(0.082)
Controls								
Inc. Group Slope?	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Country FE?	No	No	No	No	Yes	Yes	Yes	Yes
Time FE?	No	No	No	No	Yes	Yes	Yes	Yes
Half-Lives (years)								
$K = 0$	2.109	1.370	1.342	1.340	1.716	1.163	1.153	1.161
$K = 1$	4.722	3.171	3.918	3.925	4.480	3.019	3.702	3.698
$H_0 : \hat{\Theta}_{it} = 0$ (p-value)	0.004	0.001	0.002	0.002	0.005	0.000	0.000	0.000
Observations	1182	1182	1066	1066	1182	1182	1066	1066
Adj R^2	0.273	0.292	0.294	0.293	0.268	0.287	0.287	0.286

Note: Each ECM is estimated using the residuals from the DOLS specification (2) in Table 5. All specifications include first-differences of the long-run variables LNY and NFA and short-run covariates: GOV, TOT, $CRISIS$. The base groups for models including exchange rate regime and income group heterogeneity are, respectively, managed regimes and upper middle-income countries. The single lag-order was chosen using the AIC and BIC. The list of countries adopting inflation targeting regimes was taken from Roger (2010). Robust HAC standard errors are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

D Control Group Refinement Using Propensity Score

In this appendix, I present a robustness exercise intended to account for selection bias by refining the sample used in the estimation of the ECMs to ensure that the countries without capital controls are as comparable as possible to those with controls. In particular, in the language of the program evaluation literature, I estimate the propensity score of receiving treatment (imposing capital controls) conditional on pre-treatment observables. I then use the propensity score to drop countries from the control group with a very low probability of imposing capital controls. The idea here is that non-treatment countries with a high propensity score do not systematically differ from treatment countries and as such yield a better comparison group to judge the causal impact of imposing capital controls.

Let T_i denote an indicator for whether or not a country belongs to the treatment group, that is, that it will impose capital controls at some point in the sample. The propensity score is defined as the probability of belonging to the treatment group conditional on observables at the beginning of the sample, $p(T_i|\boldsymbol{X}_{i,1995})$, where $\boldsymbol{X}_{i,1995}$ is a vector of observables at the beginning of the sample. I estimate the propensity score using Logit regressions for two alternative models. The first considers a range of pre-treatment macroeconomic indicators, while the second model conditions treatment on a set of policy and institutional variables. These are summarized below.

- **Model 1:** log PPP GDP per capita, net foreign assets as share of imports, extent of real exchange rate misalignment, government expenditure to GDP ratio, log commodity terms of trade, current account balance to GDP ratio.
- **Model 2:** log PPP GDP per capita, current account balance to GDP ratio, nominal exchange rate flexibility, prior IMF agreement, average inflation over the previous 10 years, any bilateral investment treaty (BIT) in force.

After estimating the propensity score, I drop control group countries from the sample with propensity scores below certain thresholds (e.g. dropping all countries for which $p(T_i) < 0.1$). Armed with the refined sample, I then estimate the benchmark ECM from equation (2). The results are reported below in Table A5 for both models and for propensity score cutoffs $p(T_i) < 0.1$, $p(T_i) < 0.25$, and $p(T_i) < 0.5$, representing increasingly stricter criteria for the selection of the control group sample.

Table A5: ECMs with Refined Control Groups

Dependent Variable: ΔRER_t	***Model 1***			***Model 2***		
	$p(T) < 0.1$	$p(T) < 0.25$	$p(T) < 0.5$	$p(T) < 0.1$	$p(T) < 0.25$	$p(T) < 0.5$
	(1)	(2)	(3)	(4)	(5)	(6)
$\hat{e}_{t-1}$	-0.297***	-0.295***	-0.340***	-0.284***	-0.291***	-0.309***
	(0.040)	(0.046)	(0.063)	(0.041)	(0.046)	(0.055)
$\hat{e}_{t-1} \cdot K$	0.178**	0.188**	0.249***	0.167**	0.181**	0.201**
	(0.085)	(0.083)	(0.090)	(0.085)	(0.087)	(0.093)
Half-Lives (years)						
No K-Controls	1.968	1.986	1.667	2.076	2.012	1.873
With K-Controls	5.497	6.139	7.203	5.586	5.910	6.073
Observations	583	478	313	568	463	359
Adj R^2	0.245	0.262	0.318	0.241	0.261	0.281

Note: Each ECM is estimated using the residuals from the DOLS specification (2) in Table 5. All specifications include first-differences of the long-run variables LNY and NFA. $p(T)$ refers to the estimated propensity score of imposing capital controls, as defined in the text. Robust HAC standard errors are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

E Full List of Sample Countries

High Income: OECD – Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Hungary, Iceland, Italy, Japan, Luxembourg, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, United Kingdom, United States.

High Income: non-OECD – Antigua and Barbuda, The Bahamas, Bahrain, Cyprus, Israel, Malta, Saudi Arabia, Singapore, Trinidad and Tobago.

Upper Middle Income – Algeria, Brazil, Chile, Colombia, Costa Rica, Dominica, Dominican Republic, Fiji, Gabon, Grenada, Malaysia, Mexico, Poland, South Africa, St. Kitts and Nevis, St. Lucia, St. Vincent and the Grenadines, Uruguay, Venezuela.

Lower Middle Income – Belize, Bolivia, Cameroon, China, Côte d'Ivoire, Ecuador, Iran, Lesotho, Morocco, Nigeria, Pakistan, Papua New Guinea, Paraguay, Philippines, Samoa, Solomon Islands, Tunisia.

Low Income – Burundi, Central African Republic, The Gambia, Malawi, Sierra Leone, Togo, Uganda, Zambia.

F Exchange Rate Regime Classification

Data on *de facto* exchange rate regime classifications was obtained from Ilzetzki et al. (2010). Details on the different classification codes are provided in Table A6. Following Eguren-Martín (2015), "fixed" regimes are defined as those with category codes 1 through 4. "Managed" or "intermediate" regimes are those with codes 5 through 11, while "flexible" regimes correspond to codes 12 through 14. Countries with code 15, dual market in which parallel market data is missing, are excluded from the analysis.

Table A6: Ilzetzki, Reinhart, and Rogoff (2010) *de facto* exchange rate regime classification

Code	Description
1	No separate legal tender
2	Pre announced peg or currency board arrangement
3	Pre announced horizontal band that is narrower than or equal to +/-2%
4	De facto peg
5	Pre announced crawling peg
6	Pre announced crawling band that is narrower than or equal to +/-2%
7	De factor crawling peg
8	De facto crawling band that is narrower than or equal to +/-2%
9	Pre announced crawling band that is wider than or equal to +/-2%
10	De facto crawling band that is narrower than or equal to +/-5%
11	Moving band that is narrower than or equal to +/-2% (i.e., allows for both appreciation and depreciation over time)
12	Managed floating
13	Freely floating
14	Freely falling
15	Dual market in which parallel market data is missing.

As noted in the text, Table A7 provides summary statistics for the Schindler according to the three-way exchange rate regime classification. In general, flexible regimes tend to have fewer restrictions on capital flows than do fixed and managed regimes. Moreover, the average Schindler index is highest in managed regimes.

Table A7: Schindler index average intensity and standard deviation by exchange rate regime

Capital control measure:	*Overall*		*Inflows*		*Outflows*	
	mean	sd	mean	sd	mean	sd
Exchange Rate Regime						
Fixed	0.214	0.316	0.183	0.293	0.245	0.365
Managed	0.329	0.370	0.303	0.344	0.356	0.421
Flexible	0.196	0.243	0.163	0.231	0.230	0.287