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Real Exchange Rate and International Reserves in the Era of Financial Integration
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ABSTRACT

The global financial crisis has brought increased attention to the consequences of international reserves holdings. In an era of high financial integration, we investigate the relationship between the real exchange rate and international reserves using nonlinear regressions and panel threshold regressions over 110 countries from 2001 to 2020. We find the buffer effect of international reserves is more pronounced in Europe and Central Asia above a threshold of 17% of international reserves over GDP. Our study shows the level of financial-institution development plays an essential role in explaining the buffer effect of international reserves. Countries with a low development of their financial institutions may manage the international reserves as a shield to deal with the negative consequences of terms-of-trade shocks on the real exchange rate. We also find the buffer effect is stronger in countries with intermediate levels of financial openness.

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1. Introduction

The current surge in international reserves hoarding has become a topic of debate in international economics, although it is not a new phenomenon in the history of international economics. Economists consider the cost-benefit model to analyze the relationship between changes in international services holding and other macroeconomic indicators such as exchange-rate intervention policy (Levy, 1983), real exchange rate (Aizenman and Riera-Crichton, 2008), commodity terms of trade shocks (Aizenman et al., 2012), and so on. Hoarding of international reserves could be considered a self-insurance tool or buffer against external shocks.

Note that holding international reserves is not a free lunch for countries (Ben-Bassat and Gottlieb, 1992; Rodrik, 2006; Korinek and Serven, 2016); however, policymakers are also keen on making use of this tool to cope with the terms of trade shocks by managing the real exchange. Given the divergence of monetary and trade policies, many countries have followed their approaches to managing their macroeconomic indicators. Therefore, motivated by the broader scope, this study evaluates the interrelationship among international reserves holding, terms of trade shocks, and real exchange rates in the global scope. Similar to Aizenman and Riera-Crichton (2008), our study is also based on the presumption that these effects may be of a first-order magnitude for developing economies. In doing so, our study attempts to disentangle these relationships by clustering different country groups. Noticeably, the question of defining the optimal threshold for international reserves has not been fully answered.

This study is different from the existing literature for two main reasons. First, although we draw on the global sample for further analyses, the single-country characteristics are disaggregated. Since Aizenman and Riera-Crichton (2008) and Aizenman et al. (2012) claimed the effects are different between advanced and developing economies (e.g., most emerging countries are exposed to terms-of-trade shocks due to the composition of their exports), our study attempts to explain the heterogeneity from geographical and economic perspectives. Second, this study provides a benchmark for each country to reconcile their policies in the general context. Once we consider the threshold approach, we extend the existing literature that shows international reserves and real exchange rates are associated with the nonlinear shape. In doing so, we construct and present more reasons for policymakers to intervene in the macroscopic economy with cautious actions. As mentioned earlier, this study incorporates other macroeconomic indicators, and we examine the threshold under the other constraints to strengthen the understanding of different contexts.

We summarized our findings in three main points. First, the buffer effect of international reserves is more pronounced in Europe and Central Asia at a certain level above 17% of international reserves over GDP. Our first finding complements and extends Aizenman and Hutchison (2012) regarding the exchange-rate shocks and loss of international reserves. We also echo Aizenman, Cheung and Ito (2015) on international reserves holding in Europe and Asia. Second, the level of financial-institution development plays an important role in explaining the buffer effect of international reserves. To be more precise, hoarding international reserves could be beneficial for countries experiencing slower financial development. Third, the buffer effect of international reserves is stronger for intermediate levels of financial openness. For observations (countries and periods) associated with a low level of financial openness, the buffer effect is six times lower than for intermediate openness. For the advanced level of financial openness, the effect is also six times lower than for intermediate openness, but only significant at the 10% level. Our last two findings are consistent with the existing literature which is debating about the reserve holdings (Obstfeld et al., 2010; Aizenman and Riera-Crichton, 2008; Alberola et al., 2016).

This study contributes to the existing literature by exploring the threshold of international reserves over GDP necessary to activate buffer effects. In addition, by using linear and nonlinear estimates, our study explores the role of financial-institution development and financial openness in driving the effects of international reserves and the real exchange rate. Whereas the previous study emphasizes the linear estimation, our study extends and discusses the existence of nonlinear effects. We organize this paper as follows: Section 2 reviews the contemporary literature. Section 3 presents the methodology. The main analysis is provided in section 4. Section 5 concludes.

2. Literature review

In the literature, we find two main motivations that led countries to hold international reserves, as noted by Aizenman and Lee (2007). The first motive — the mercantilist one — consists of holding international reserves to weaken the domestic currency and promote an export-led growth strategy. The second motive — the precautionary or deterrent one — consists of holding international reserves as a form of self-insurance against foreign financial shortfalls. Several studies have undertaken empirical investigations to estimate the relative importance of these two main motives (Aizenman and Lee, 2007; Delatte and Fouquau, 2011; Ghosh et al., 2017; Cabezas and De Gregorio, 2019; Choi and Taylor, 2022). To test the relative importance of these two motives, we can recall that Aizenman and Lee (2007) use two types of variables for a panel of 53 developed and developing economies between 1980 and 2000: (i) lagged exports and deviation of the national price from the trends, based on income levels for the mercantilist motive; and (ii) capital account liberalization and dummy variables for the adjustments in the aftermath of unanticipated sudden-stop crises. They find limited support for the mercantilist motive and strong support for the precautionary or deterrent motive.

The subsequent studies of Delatte and Fouquau (2011) and Ghosh et al. (2017) examine the possibility of nonlinear behaviors and time-varying patterns in the motivation for holding international reserves. In an attempt to explain the increase in reserves holding at that time, Delatte and Fouquau (2011) use a panel threshold model for 20 emerging countries between 1981 and 2004. They explain, for example, that the change in the configuration of the capital market for emerging countries (switching from debtor to creditor) could explain why the elasticity may change in the model. They use several candidates for the threshold variable that correspond to the mercantilist motive, precautionary motive, and the influence of the US external position. In their main results, they find a deterioration in the US external position increases the weight of the mercantilism concern relative to the precautionary concern.

Ghosh et al. (2017) recall that the holding of international reserves by emerging economies dramatically increased between 1980 and 2010.[1] In their empirical study, they examine the motives

[1]These trends are confirmed in the 2010s, as shown by Arslan and Cantú (2019) The level of reserves held by the emerging economies reached 30% of GDP in the 2008-2017 period, when Asian emerging economies and oil exporters (notably Saudi Arabia and Algeria) held the largest stocks.

behind these holdings, using a panel of 43 emerging economies between 1980 and 2010. In their baseline specification, they exclude country and time fixed effects to capture both cross-country and cross-period correlations. Additionally, to consider time-varying patterns and asymmetries, they use three subsamples (pre-Asian crisis (1980-97), post-Asian crisis (1998-2004), and a surge in reserve holdings and global imbalances (2005-2010)) together with quantile regressions. They find the relative importance of each motive may change over time. For example, the protection against capital account shocks becomes more important in the immediate aftermath of the East-Asian crisis. Similar to the empirical work of Delatte and Fouquau (2011), they find mercantilist motivations were probably more important during the 2000s. They also note emerging countries may have become more risk averse at the end of their sample.

Cabezas and De Gregorio (2019) explore an interesting research avenue. In addition to the traditional motives for holding international reserves, they investigate the precautionary motive as a deterrent role of speculation on the domestic currency. Indeed, they note that a large stock of reserves also acts as a deterrent for speculation. Thus, even if they are not used, international reserves may still have a smoothing function when shocks to foreign financing occur. Similar to Dominguez et al. (2012), they find the accumulation of reserves was massive before the Global Financial Crisis (GFC), but many countries were reluctant to use these reserves in the immediate aftermath of the crisis. In their empirical specification, they use a sample of 52 emerging and developing economies observed between 2000 and 2013. They include country fixed effects in their regressions and use two subsamples (2000-2008 and 2009-2013) to consider the GFC. They find the mercantilist motive was important in the first subsample. The deterrent motive (precautionary) is significant over the whole sample, along with the self-insurance motive (precautionary) for financial account variables. Overall, both the mercantilist and the precautionary motives have similar explanatory power.

Choi and Taylor (2022) find increases in publicly held assets (reserves) are associated with depreciation of the domestic currency, especially combined with capital controls. They propose the following rationale: combined reserves and capital controls can affect trade balances via undervaluation (mercantilist motive), whereas reserves without controls can insure against crises (precautionary motive) independently of exchange rates.[2] In their empirical investigation, they use a sample of 22 advanced and 30 developing and emerging economies observed between 1980 and 2015. They distinguish between the net external assets held by the private sector and those held by the public sector. In the baseline regressions, a dummy variable based on the *KAOPEN* index captures the financial openness (Chinn and Ito, 2006). They split their sample into subperiods (1980-1986 (1), 1987-1993 (2), 1994-2000 (3), 2001-2007 (4), and 2008-2015 (5)) to detect time-varying patterns. In their pooled cross-sectional regressions between 1994 and 2007, they find the increase in publicly held reserves and, even more, for financially closed economies is associated with an exchange-rate depreciation, confirming the main predictions of their theoretical model. After controlling for the GDP and terms of trade over the period 1980-2007, a one-percentage-point increase in international reserves is associated with a 0.63% depreciation. In financially closed countries, the depreciation of the real exchange rate is even larger (0.94%).

[2] See Propositions 2 and 3 in Choi and Taylor (2022) for more details.

After this review of the recent literature on the different motives behind the accumulation of international reserves, we focus on studies that analyze the interaction between real exchange rates, international reserves, and terms of trade (Aizenman and Riera-Crichton, 2006, 2008; Aizenman et al., 2012; Al-Abri, 2013; Coudert et al., 2015; Adler et al., 2018; Aizenman and Jinjarak, 2020).

Our empirical investigation aims at exploring the relationship between the real exchange rate and international reserves. In this respect, Aizenman and Riera-Crichton (2006, 2008) investigate whether the accumulation of international reserves helps mitigate the consequence of terms of trade shocks on the real exchange rate. Indeed, using panel data regressions for 60 developed countries and 20 emerging countries over the period of 1970-2004, they find the reduction in the magnitude of the real-exchange-rate adjustment triggered by capital flows may contribute to the mitigation of terms of trade shocks. This *buffer effect* of international reserves is especially strong for emerging Asia. They find financial depth significantly reduces the role of reserves as a shock absorber for developing economies (Aizenman and Riera-Crichton, 2006).[3]

In our empirical investigation, we deepen these last results of Aizenman and Riera-Crichton (2006, 2008) for a large macroeconomic panel of 110 countries over a more recent period spanning from 2001 to 2020, where financial integration has known several evolutions. To avoid ad hoc country grouping, we use panel threshold regressions (Hansen, 1999). Thus, we provide more systematic evidence about the existence of financial-indicator threshold effects. We offer a more complete view of financial development by using the aggregated and disaggregated financial indexes introduced by Svirydzenka (2016). To provide a multidimensional view of financial development, these financial indexes go beyond the traditional variables used to measure the development of financial markets and institutions (e.g., private-sector credit to GDP and stock market capitalization to GDP). We find countries with a low development of their financial institutions may use the international reserves as a shield to deal with the negative consequences of terms-of-trade shocks on the real exchange rate. Thus, together with the more common recommendation of better management of international reserves, the development of sound financial institutions could be important to deal with negative consequences of terms-of-trade shocks.

Aizenman et al. (2012) focus on the commodity terms-of-trade shocks[4] for several Latin American countries. They recall the buffer-stock approach to international reserves goes back to the Bretton Woods era. The prevailing rule of thumb for an adequate level of international reserves was that four months of imports was considered adequate at that time. They use panel data over the period of 1970-2009, with quarterly data, to understand (i) how real exchange rate reacts to commodity terms of trade shocks in several Latin America countries; and (ii) the influence of international reserves in this adjustment. They use two versions for their error-correction model. In the first one, the international-reserve-to-GDP ratio is a long-run determinant of the real exchange rate. In the second one, the international reserve-to-GDP ratio also affects the adjustment speed toward the long-run equilibrium. To illustrate the second version of the error-correction model, they give the example of a commodity-terms-of-trade shock that implies a real appreciation of the domestic currency. If the central bank absorbs part of the shock in relative export revenue by increasing the stock of international reserves, the subsequent expansion of the domestic currency

[3]See Appendix C (Table C.7) and D (Table D.1.1 and D.1.2) of Aizenman and Riera-Crichton (2006).

[4]Which are generally more volatile than global terms of trade shocks.

will push toward a real depreciation, thereby softening the real appreciation. In this sense, international reserves could influence the adjustment speed toward the long-run equilibrium.

In the fixed-effects panel regressions, they include a third version of the error-correction model, where the international-reserves-to-GDP ratio interacts with the transitory commodity-terms-of-trade shocks. Indeed, they find that international reserves do act as a buffer for commodity-terms-of-trade shocks (Model 3). They also find an increase in the persistence of real-exchange-rate deviations (Model 1) and a reduction in the error-correction term for the real exchange rate (Model 2) after a term-of-trade shock.[5] In the country analysis, they use a seemingly unrelated regression (SURE) procedure to consider the slope heterogeneity in this group of countries. On the whole, the mitigation effect of international reserves after terms of trade shocks and the reduction in the adjustment speed are confirmed for a majority of countries. Interestingly, they also consider the influence of the quality of institutions on their various specifications. They use a dummy variable for the bad and good institutions based on a transformation of the International Country Risk Guide (hereafter ICRG). In countries with good institutions, we observe an increase in the persistence of real-exchange-rate deviation associated with a decrease in the adjustment speed that corresponds to a reduction in exchange-rate volatility.

The empirical study of Al-Abri (2013) focuses on the volatility of the real exchange rate in commodity exporting economies. For a dynamic panel of 53 economies and 5-year averaged data between 1980 and 2007, he finds that greater financial integration mitigates the effect of terms-of-trade shocks on the volatility of the real exchange rate. Additionally, he uses five different variables for financial integration, including international reserves. Interestingly, the mitigation effect of international integration is larger when the author uses foreign direct investment integration rather than portfolio integration. Indeed, long-run capital flows, such as foreign direct investment, could help stabilize the price of non-tradable goods. As shown by Ouyang and Rajan (2013), the fluctuations in the price of non-tradable goods explain a large portion of exchange rate volatility, especially in commodities exporting countries. Consequently, from the perspective of the mitigation of terms-of-trade shocks, better financial integration could be an alternative policy to the accumulation of international reserves. This conclusion echoes one of our main results, where the buffer effect of international reserves is larger for observations (countries and / or periods) with low development of their financial institutions.

Coudert et al. (2015) analyze the impact of terms-of-trade volatility on the real exchange rate for a panel of 68 commodity exporters, which are not homogeneous in terms of economic development. They also use panel-cointegration techniques along with panel-threshold-regression techniques to examine the relationship between the real exchange rate and the terms of trade. While they use a yearly sample that spans between 1980 and 2012 for the estimation of the long-run determinants of the real exchange rate, they use a monthly sample that spans from January 1994 to December 2012 for the short-run impact on terms of trade volatility. Interestingly, they find the terms-of-trade volatility plays an important role as a determinant of the real exchange rate in the short-run, but only in the regime where the commodity and financial volatility is high (measured by the S&P GSCI and the VIX) and for advanced commodities exporters, namely, Australia, Canada, and Norway.

Adler et al. (2018) find the holding of international reserves is a key tool to smooth the

[5] See Aizenman et al. (2012) for more details.

adjustments after large terms-of-trade shocks for a large macroeconomic panel of 150 countries observed between 1960 and 2015. In their empirical investigations, they rely on the estimation of a Markov-switching process with level shifts in the terms of trade. Indeed, they identify regimes of *low* and *high* terms of trade for each country. Then, they successively estimate a set of dynamic panel equations to examine the dynamic impact of the terms of trade shifts after the identification of these regimes of boom and bust for the terms of trade. They find countries with a high level of international reserves can smooth (delay) the adjustment of the current account during regimes of falling terms of trade. However, no statistical differences exist between countries with low and high levels of international reserves for boom episodes, where constraints on reserve accumulation are absent. Unfortunately, they do not report the results for the real exchange rate. However, we can reasonably infer these asymmetrical effects of international reserves are also present in the exchange-rate adjustment.

To provide some perspectives on the efficient use of international reserves, we close this review of the literature with the work of Aizenman and Jinjarak (2020). They evaluate the opportunity costs of buffer-stock services for several emerging markets over the 2000-2019 period with quarterly data. As noted in Rodrik (2006), the opportunity cost of reserves in terms of foreign currency can be measured as the sovereign spread between the private-sector cost of short-term borrowing abroad and the yield on international reserves. Although it is a second-best policy,[6] they found that a counter-cyclical management of international reserves (i.e., hoarding reserves in times of plenty and selling them on rainy days) may generate sizeable benefits, especially for countries with highly volatile real exchange rate and large sovereign spread.

3. Methodology and data

3.1. Data

We use annual data for a macroeconomic panel of 110 countries from 2001 to 2020. Along with the countries' list, the definitions and sources of the data are provided in Table A.1 in Appendix A. We follow Aizenman and Riera-Crichton (2008) to construct our variables, such as the real effective exchange rate, *rer*; trade openness, *to*; terms of trade *tot*; effective terms of trade, *etot*; and international reserves, *res*. We also add some common determinants of the real effective exchange rate, namely, GDP per capita, *gdppk*, and government expenditures as a percent of GDP, *govexp*. More precisely, providing some details for the main variables may be useful: *rer* is the natural log of the real effective exchange rate (an increase amounts to appreciation); *to* is the natural log of one plus the sum of the export-to-GDP and import-to-GDP ratios; *tot* is the natural log of the ratio between the export-value unit and the import-value unit; and, finally, *res* is the natural log of one plus the reserves-to-GDP ratio expressed as a percentage. Afterward, the effective terms of trade, *etot* is obtained by multiplying *to* by *tot*. We present the descriptive statistics in Table 1.

In Figure 1, we follow Arslan and Cantú (2019) to visualize the evolution of international reserves for a sample of the largest holders of international reserves in emerging and developing economies.[7] We see a number of emerging countries hold a sizeable amount that represent a large

[6]The first-best policy calls for prudential regulations.

[7]We provide the figure for the full sample in Appendix A.2.

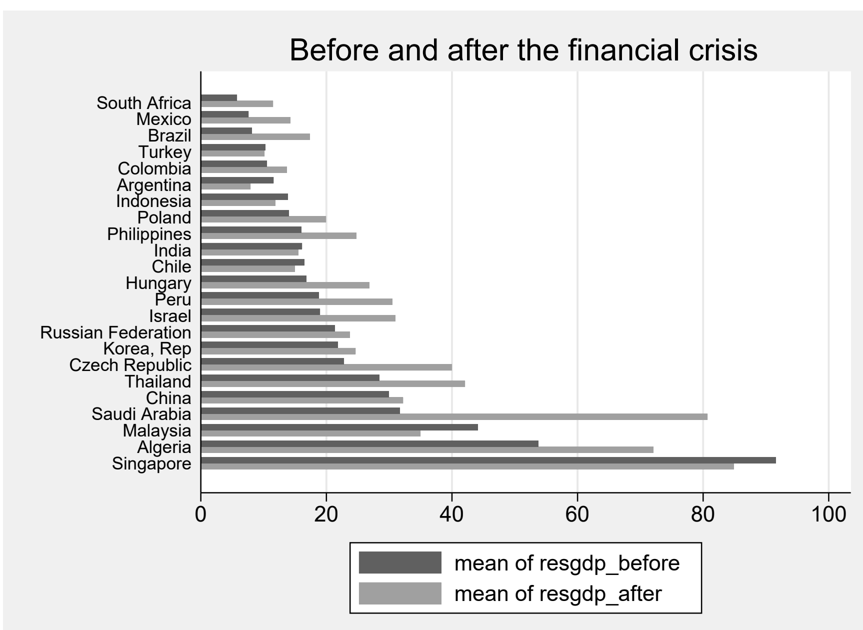
Table 1: Descriptive statistics

	Observations	**Mean**	**Standard deviation**	**Minimum**	**Maximum**
rer	2,200	4.633	0.183	2.847	5.567
to	2,200	3.650	0.482	2.378	5.392
tot	2,200	-0.015	0.371	-2.112	2.513
res	2,200	2.523	0.893	0.093	4.697
govexp	2,127	2.696	0.371	-0.050	3.565
gdppk	2,200	4.605	0.541	3.159	5.775

Source: Authors' computations.

share of their GDP, confirming the trends in the accumulation of foreign reserves. Additionally, we can observe that the trends observed by Arslan and Cantú (2019) are confirmed in the more recent period. A majority of countries belonging to this group have more reserves after the financial crisis. For Eastern European countries, we can note the Czech Republic and Hungary have largely increased their holding of international reserves. Also, for the oil exporters, Algeria and Saudi Arabia have become two of the largest holders relative to their GDP, as evidenced by a higher average after the GFC.

Figure 1: Large holders of international reserves as percent of GDP (before and after the GFC)



Notes: We select a sample of emerging and developing economies as in Arslan and Cantú (2019). We split the sample into two sub-periods, 2001-2007 and 2010-2020, to observe the consequences of the GFC on reserves accumulation. Source: Authors' calculations.

To provide a better understanding of the relationship between the real exchange rate and international reserves, we use two kinds of financial variables to assess the influence of financial development and openness. First, we use three indexes of financial development, financial institution development, and financial market development where several characteristics of financial markets are considered, namely depth, access, and efficiency (Svirydzenka, 2016). Second, we use the *KAOPEN* index of Chinn and Ito (2006), which is a measure of the inverse of capital controls' intensity. For clarity purposes, providing some explanations about the construction of these financial indexes may be useful.

First, we start with the financial-development index constructed by Svirydzenka (2016). The empirical literature on financial development pays particular attention to two measures of financial depth, namely, the ratio of private credit to GDP and stock market capitalization, also as a ratio to GDP. However, modern financial systems are multifaceted, and a growing constellation of financial institutions and markets facilitates the provision of financial services. As underlined by Cihak et al. (2012) and Aizenman, Jinjarak and Park (2015), the effect of financial development on growth is nonlinear and uneven across sectors. To capture several dimensions of the financial development, Svirydzenka (2016) constructs a series of the financial-development indexes aimed at capturing the financial development of institutions and markets in terms of depth, access and efficiency.[8]

Moreover, we can describe the financial institutions and markets covered by these several indexes, where the financial institutions include banks, insurance companies, mutual funds, and pensions funds; and the financial markets include stock and bond markets. Following the matrix of financial systems developed by Cihak et al. (2012), Svirydzenka (2016) defines financial development as a combination of depth (size and liquidity of markets), access (ability of individuals and companies to access financial systems), and efficiency (ability of institutions to provide financial services at a low cost and with sustainable revenues, and the activity level of capital markets).

The financial-development index is constructed using a standard three-step approach: (i) normalization of variables, (ii) aggregation of normalized variables into subindices representing a particular functional dimension, and (iii) aggregation of the subindices into the final index. A total of nine indexes are constructed, namely, *FID*, *FIA*, *FIE*, *FMD*, *FMA*, and *FME*, where the letters *I* and *M* indicate institutions and markets, and the letters *D*, *A*, and *E* indicate depth, access, and efficiency. The first three indexes are aggregated in the financial-institution index (*FI*) and the last three indexes are aggregated in the financial-market index (*FM*). These two last indexes are aggregated into the overall measure of financial development (*FD*). In the Table 2, we present the variables selected for each category to ease the reading of the interpretations in the results section.

After the treatment of missing data, they normalize winsorized indicators with the following min-max procedure in equations (1) and (2):

$$I_x = \frac{x - x_{\min}}{x_{\max} - x_{\min}} \tag{1}$$

$$I_x = 1 - \frac{x - x_{\min}}{x_{\max} - x_{\min}}, \tag{2}$$

where x stands for the raw data and I_x stands for the transformed continuous indicator. The formula in equation (2) is used in the case of indicators, where a higher value indicates worse performance.

[8]This series of financial indexes and subindexes first appeared in Sahay et al. (2015).

Table 2: Selected variables in the financial development indexes

Category	**Indicator**
Financial Institutions	
Depth	Private-sector credit to GDP Pension fund assets to GDP Mutual fund assets to GDP Insurance premiums, life and non-life to GDP
Access	Bank branches per 100,000 adults ATMs per 100,000 adults
Efficiency	Net interest margin Lending-deposits spread Non-interest income to total income Overhead costs to total assets Return on assets Return on equity
Financial Markets	
Depth	Stock market capitalization to GDP Stocks traded to GDP International debt securities of government to GDP Total debt securities of financial corporations to GDP Total debt securities of non-financial corporations to GDP
Access	Percent of market capitalization outside of top 10 largest companies Total number of issuers of debt (domestic and external, corporations)
Efficiency	Stock market turnover ratio (stocks traded to capitalization)

Source: Reproduced from Svirydzenka (2016).

Then the transformed continuous indicators are used to construct the six subindexes, namely, *FID* (financial-institution depth), *FIA* (financial-institution access), *FIE* (financial-institution efficiency), *FMD* (financial-market depth), *FMA* (financial-market access), and *FME* (financial-market efficiency):

$$FI_j = \sum_{i=1}^{n} w_i I_i \tag{3}$$

$$FM_j = \sum_{i=1}^{n} w_i I_i, \tag{4}$$

where the weights w_i are obtained by a principal component analysis in equations (3) and (4). These subindexes are renormalized by the following equation (1).

Finally, the financial-institution index (*FI*), the financial-market index (*FM*), and the financial-development index are obtained using the following linear aggregation in equation (5), (6) and (7):

$$FI = \sum_{j=1}^{n} w_j FI_j \tag{5}$$

$$FM = \sum_{j=1}^{n} w_j FM_j \tag{6}$$

$$FD = w_{FI} FI + w_{FM} FM. \tag{7}$$

Again, the series are re-normalized following equation (1) and the weights are obtained using principal component analysis across time and countries. More precisely, the weights are the square factor loadings of the first principal component. As noted by Svirydzenka (2016), the banking-system credit to the private sector has a weight of 25% in the depth subcomponent of the financial-institution index and only 40% in the financial-institution subcomponent.

Second, we continue with the *KAOPEN* index, which is a measure of financial openness (i.e., openness of the capital account). Introduced by Chinn and Ito (2002), this index aims at measuring the extensity of the capital controls (because it is an inverse measure of the intensity of capital controls) based on the information in the IMF's Annual Report on Exchange Rate Arrangements and Exchange Restrictions (*AREAR*).

The *KAOPEN* index is computed from binary dummy variables. These dummy variables are used to codify the restrictions on cross-border financial transactions reported in the *AREAR*. Until 1996, the *AREAR* assigned dummy variables for the four major categories of the restriction on the capital account (the existence of multiple exchange rates (k_1), restrictions on current account transactions (k_2), restrictions on capital account transactions (k_3), and the requirement of the surrender of export proceeds (k_4)). To understand the complexity of capital control policies, these four categories have been more disaggregated for 1996 (the variables indicating restrictions on current account transactions have been divided into 13 categories).

Because they are focused on the effect of financial openness, Chinn and Ito reverse these binary variables. When variables are equal to zero, the capital account restrictions exist. In addition, for the k_3 category, Chinn and Ito use a five-year window during which capital controls were not in effect ($SHAREk_3$):

$$SHAREk_{3,t} = \left(\frac{k_{3,t} + k_{3,t-1} + k_{3,t-2} + k_{3,t-3} + k_{3,t-4}}{5} \right)$$

Then they construct their index for capital account "openness", which is the first standardized principal component of k_1, k_2, $SHAREk_3$, k_4 (Chinn and Ito, 2006). The more the country is open to cross-border capital flows, the higher the *KAOPEN* index. This index tries to measure the intensity of capital restriction. The index was first designed to measure the extensity of capital controls, but because it incorporates various kinds of restrictions, it may be a good proxy to gauge the intensity of capital account restrictions. Note that the *KAOPEN* index is highly correlated with other measures of financial openness (Chinn and Ito, 2006).

3.2. Methodology

Along with panel nonlinear regressions with interaction terms, we use panel threshold regressions introduced by Hansen (1999) in the empirical literature.[9] For clarity purposes, we will briefly expose the single threshold model as in Wang (2015). We consider the following model to investigate the possibility of threshold effects in the relationship between the real exchange rate and international reserves, as suggested by Aizenman and Riera-Crichton (2008) and Aizenman et al. (2012):

$$rer_{i,t} = \mu + \beta_1 etot_{i,t} I\left(res_{i,t-1} \leq \gamma\right) + \beta_2 etot_{i,t} I\left(res_{i,t-1} > \gamma\right) + \alpha_1 x_{i,t} + u_i + e_{i,t}, \tag{8}$$

where subscripts $i = 1, \ldots, n$ represents the country and $t = 1, \ldots, T$ index the time. μ is a constant term, u_i is the country-specific fixed effect, and ε_{it} is the error term. The involved variables are presented in Table A.1 and in the previous section. Indeed, *rer* is the real effective exchange rate and $I(.)$ is an indicator function indicating the regime defined by the threshold variable; *to* stands for the trade openness, *res* represents the international reserves. Here, the threshold variable and the regime-dependent variable are different.[10] The effect of the effective terms of trade *etot* depends on the lagged level of international reserves, as we can see in equation (8). The independent regime control variables, x, include the GDP per capita and the government expenditure as a percent of GDP. We can also write equation (8) as:

$$rer_{i,t} = \begin{cases} \mu + \beta_1 etot_{i,t} + \alpha_1 x_{i,t} + u_i + e_{i,t}, & res_{i,t-1} \leq \gamma, \\ \mu + \beta_2 etot_{i,t} + \alpha_1 x_{i,t} + u_i + e_{i,t}, & res_{i,t-1} > \gamma. \end{cases} \tag{9}$$

In a first step, the search is restricted to a certain interval of quantiles for the threshold variables to estimate the threshold value γ. The estimator value for the threshold is the value that minimizes the residual sum of square,[11] that is,

$$\widehat{\gamma} = \arg\min_{\gamma} S_1(\gamma). \tag{10}$$

If the value of the threshold is known, the model is similar to a linear model. But if γ is unknown, a nuisance parameter problem arises, which makes the distribution of the threshold nonstandard. To test the true value of the threshold $\gamma = \gamma_0$,[12] Hansen (1999) proposes to form the confidence interval using the "no-rejection" method with likelihood-ratio (LR) statistics, as follows:

[9] One important advantage of this approach is to test the statistical significance of the threshold values. Determining whether thresholds are statistically significant when thresholds are chosen in an ad hoc manner is difficult.

[10] Later, we use the financial-development indicators as the threshold variable to test the presence of thresholds in the buffer effect of international reserves.

[11] The models are estimated in within-group deviation form.

[12] The threshold effect may be detected only in the investigated sample and not in the statistical population.

$$\mathrm{LR}_1(\gamma) = \frac{\{\mathrm{LR}_1(\gamma) - \mathrm{LR}_1(\widehat{\gamma})\}}{\widehat{\sigma}^2} \xrightarrow{\text{Pr}} \xi \tag{11}$$
$$\Pr(x < \xi) = \left(1 - e^{\frac{-x}{2}}\right)^2.$$

The α quantile can be computed from the following inverse function of equation (11):

$$c(\alpha) = -2\log(1 - \sqrt{1-\alpha}). \tag{12}$$

For example, for $\alpha = 0.1, 0.05,$ and 0.01, the quantile is 6.53, 7.35 and 10.59, respectively. If the LR_1 statistic does not exceed $c(\alpha)$, we do not reject H_0.

In a second step, testing for a threshold amounts to testing whether the coefficients are the same in each regime. Under the null, the coefficients are the same, and we revert to a linear model, whereas the coefficients are significantly different in each regime under the alternative hypothesis. We test the linear model versus the single threshold model:

$$H_0 : \beta_1 = \beta_2 \quad H_a : \beta_1 \neq \beta_2. \tag{13}$$

The F statistic is constructed as

$$F_1 = (S_0 - S_1(\hat{\gamma}))\,/\hat{\sigma}^2, \tag{14}$$

where S_0 is the RSS for the model without a threshold, S_1, is the RSS for the model with a specific threshold $\hat{\gamma}$, $\hat{\sigma}^2$ is the residual variance for a specific threshold. Under H_0, the threshold is not identified, and F_1 has nonstandard asymptotic distribution. Hansen (1996) uses a bootstrapped likelihood-ratio test (asymptotically valid).[13]

4. Empirical results

In our empirical approach, we use two kinds of regressions, namely panel nonlinear regressions with interaction terms and panel threshold regressions with country fixed effects, to capture the ability of lagged international reserves to deal with the consequence of terms-of-trade increase on the real effective exchange rate.[14] From a mathematical perspective, nonlinear regressions are more general than threshold regressions, such as piecewise linear regressions. Additionally, threshold regressions can provide clearer interpretations, especially in the case of interactions between two continuous variables. These two types of regressions can be seen as complementary. In the following, we present the results of the panel nonlinear regressions for the buffer effect for different levels of financial development, financial-institution development and financial-market development.

Then, we present the results of the panel threshold regressions where the threshold variable is the lagged reserves, in order to identify a threshold of international reserves from which the

[13] See Hansen (1999) and Wang (2015) for further details.

[14] The real exchange rate is stationary in all the tests we conduct. These tests are available upon request.

buffer effect can be different. Next, we test the buffer-effect intensity for different levels of financial developments. The threshold variable will be the financial-development indexes. Our central hypothesis is that several countries could use international reserves as a substitute for sound financial institutions in order to deal with the consequences of terms-of-trade shocks on the real effective exchange rate. Lastly, we provide a robustness check with the *KAOPEN* index of financial openness (Chinn and Ito, 2006) as the threshold variable.

4.1. Nonlinear regressions

We test the following baseline regressions (see Table 3), where the buffer effect is captured by the negative coefficient on the interaction term between lagged reserves and terms of trade:

$$rer_{i,t} = \mu_i + \beta_1 gdppk_{i,t} + \beta_2 govexp_{i,t} + \beta_3 etot_{i,t} + \beta_4 res_{i,t-1} + \beta_5 etot_{i,t} \times res_{i,t-1} + u_{i,t}, \tag{15}$$

The results are the following for countries with complete observations.[15] In the baseline regressions, the coefficients for the main variables have the expected signs. For clarity purposes, we describe the interpretation of the signs of the coefficients as follows. First, the positive coefficient on the GDP per capita variable intends to capture the Harrod-Balassa-Samuelson effect, that is, the effect of relative productivity on the real exchange rate.[16] Second, the positive coefficient on the government consumption variable is also related to the Harrod-Balassa-Samuelson effect. Indeed, government consumption is related to a real-exchange-rate appreciation, because government consumption is typically associated with the consumption of non-tradable goods.[17] Third, the positive coefficient on the third explanatory variable captures the effect of terms of trade shocks on the real exchange rate. When the terms of trade increase, the price of exports grows more rapidly than the price of imports. Thus, this increase in international purchasing power thereby induces an increase in consumption of both domestic and foreign goods. In turn, this led to an increase in the price of domestic goods and a real appreciation. The empirical literature generally finds the income effect is stronger than the substitution effect.[18] Fourth, the positive coefficient on the lagged reserves shows that holding reserves may lead to a real appreciation.[19]

Finally, the fifth coefficient is the main coefficient of interest in this study. It captures the effect of the interaction between the effective terms of trade and the level of lagged reserves on the real exchange rate. The buffer effect of exchange-rate reserves may be presented as follows: we may expect a real appreciation when countries face positive terms-of-trade shocks. In turn, countries may seek to insulate themselves from the negative consequences of exchange appreciation.

[15] We use the IRR classification (Ilzetzki et al., 2019) to control for the influence of exchange rate regimes. Our results are robust to the inclusion of exchange-rate regimes in the baseline equation. Additionally, the very low p-value of the Ramsey RESET test indicates this interaction model may not be sufficient to capture the threshold effects in the buffer effect. This test leads us to use threshold regressions in the next sections.

[16] See Lothian and Taylor (2008) for long-run evidence on the link between productivity differentials and equilibrium exchange rates.

[17] Interestingly, Galstyan and Lane (2009) provide empirical evidence that government consumption is associated with real-exchange-rate appreciation, and government investment may be related to real-exchange-depreciation.

[18] See De Gregorio and Wolf (1994) and Mendoza (1995) for instance.

[19] This positive effect is robust to the inclusion of exchange-rate-regime variables.

Table 3: Baseline nonlinear regression

	(1)
Variables	*rer*
gdppk	0.6589***
	(0.0725)
govexp	0.1435***
	(0.0292)
etot	0.0369***
	(0.0134)
L.res	0.0266***
	(0.0098)
etot × *L.res*	-0.0196***
	(0.0047)
Constant	1.1186***
	(0.3733)
Observations	1,900
Number of countries	100
Adjusted R-squared	0.4395
RMSE	0.1198

Note: Bootstrapped standard errors in parentheses where 10,000 replications have been used. Fixed effects are included, but not shown. ***, **, * indicate statistical significance at the 1%, 5% and 10% levels, respectively. L stands for the lag operator. The results are very similar when we use lagged values for all the explanatory variables. Source: Authors' estimates.

Building up reserves may be used as a shield to lessen real appreciations after positive terms-of-trade shocks. We interpret our result as the buffer effect being observed when the coefficient of the interaction term is negative and statistically significant. Indeed, we find the buffer effect is statistically significant for this large macroeconomic panel over the last 20 years (see Table 3).[20]

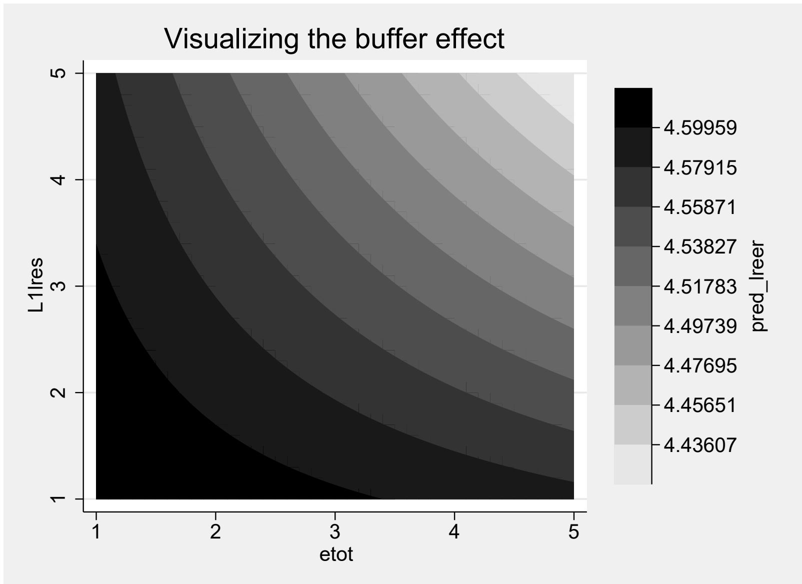
In Figures 2 and 3, we provide the contour plot and the 3-D plot to illustrate the interaction between the effective terms of trade and the lagged international reserves.[21] As mentioned previously, identifying the interaction between two continuous variables may be difficult because there are no discrete values for which we could interpret the influence of a first explanatory variable, which is the effective terms of trade, for different levels of a second explanatory variable, which is the lagged level of reserves, on the real exchange rate. Several points are of interest regarding Figure 2. On a Cartesian plane, we can see the real appreciation is more limited when countries hold more lagged reserves. This observation is represented by the lighter areas in Figure 2. Conversely, the effect of a terms-of-trade shock will be stronger in countries with a low level of lagged reserves. This observation is represented by the darker areas in Figure 2. Furthermore, we provide a 3-D plot to illustrate the same interaction for the buffer effect.

On the one hand, we can see in Figure 3 that the effect of terms-of-trade shocks are stronger when countries have a low level of reserves. In the 3-D plot, the areas in red correspond to the darker areas in the contour plot. On the other hand, we can see in Figure 3 that the effect of terms-of-trade shocks are weaker when countries have a high level of reserves. In the 3-D plot,

[20] In Appendix B.3, we provide empirical evidence showing our results are robust when the common factors (with homogeneous or heterogeneous factor loadings) are considered. The results are very similar when we lag all the variables.

[21] All the calculations were performed with Stata 17.0.

Figure 2: Contour plot for the buffer effect



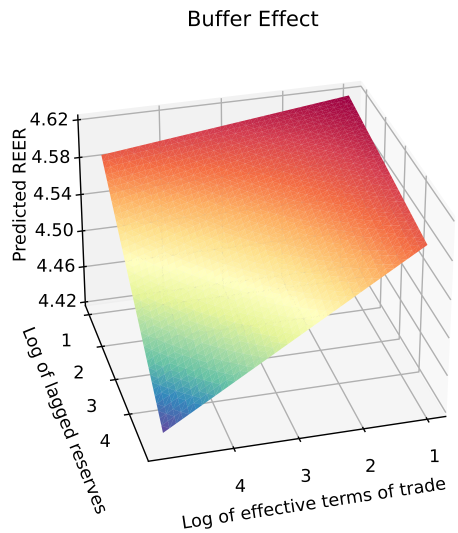
Note: The lighter areas indicate the buffer effect (i.e., the mitigation of real-exchange-rate appreciation after a terms-of-trade shock) is stronger when the level of reserves is higher. We include year fixed effects in the regressions. The results are similar without the year fixed effects. Source: Authors' estimates.

the areas in blue correspond to the lighter areas in the contour plot. In the spirit of Aizenman and Riera-Crichton (2008), we may conjecture that the buffer effect is stronger for some regions of the world economy. Indeed, some regions could be more affected by terms-of-trade shocks than others. As a corollary, we can also conjecture that the buffer effect is statistically different for various levels of lagged reserves. We investigate these conjectures in the following tables. The second type of conjectures that we investigate are related to the level of financial development and financial openness. We can reasonably infer that countries with low levels of financial development and unsound institutions will experience a stronger buffer effect. Indeed, the central contribution of this paper to the literature is to provide empirical evidence that countries with a low development of their financial markets and institutions may use the international reserves as a shield to deal with the negative consequences of terms-of-trade shocks on the real exchange rate.

In Appendix B.4, we find the buffer effect has a higher level of significance after the GFC. This finding is consistent with those of Aizenman and Sun (2012), Cabezas and De Gregorio (2019), and Dominguez et al. (2012). During the first phase of the GFC, several countries have experienced large depletions in their international reserves to cope with the consequences of the crisis, especially due to the exposure to short-term debt in times of falling export revenues. Aizenman and Sun (2012) find countries with a majority of emerging economies in their sample used less than one-fourth of their pre-crisis international reserves stocks. In the aftermath of the crisis, many countries have stopped using their reserves, fearing an almost complete depletion of their international reserves stocks and the capacity to buffer external financing shocks.

To quantify the regional heterogeneity, we test the baseline regressions for various country

Figure 3: 3-D plot for the buffer effect



Note: The blue areas indicate the buffer effect (i.e., the mitigation of real exchange rate appreciation after a terms-of-trade shock) is stronger when the level of reserves is higher. We include year fixed effects in the regressions. The results are similar without the year fixed effects. Source: Authors' estimates.

groups based on the World Bank's classification in Table 4. On the whole, the R-squared and the RMSE are quite close to those of the baseline regression. The coefficients have the expected signs for the GDP per capita, the government consumption, the effective terms of trade, and the lagged reserves in these regressions, with the notable exception of countries in the Middle East and North Africa group. For East Asia and Pacific (EAS), Europe and Central Asia (ECS), and Sub-Saharan Africa (SSF), the buffer effect is around -0.111, -0.018, and -0.023, respectively. We also do not detect any buffer effect for Latin America and the Caribbean (LAC), Middle East and North Africa (MEA), and South Asia (SAS). These results are in line with Aizenman and Riera-Crichton (2008), who find the highest value for the buffer effect in Asia. This regional heterogeneity of the buffer effect may be due to different levels of developments of the financial markets and financial institutions. Additionally, we can argue that the level of financial openness may also greatly influence the buffer effect.

In Appendix B.5, we estimate the buffer effect for other country groups, namely, OECD countries, non-OECD countries, ECS countries outside the eurozone (hereafter EZ), and commodity exporters. Although we do not find any buffer effect for OECD countries, we do find a similar buffer effect (-0.0198) for non-OECD countries to those in the baseline regression in Table 3. Because the ECS countries group includes the euro area, controlling for the presence of these countries in this group may be worthwhile. Indeed, the European Central Bank's policies provide buffers for most of the EZ countries, especially to deal with intra-zone capital flights through the TARGET 2 system (Cheung et al., 2020). Instead of running down reserves like Mexico in 1994,

Table 4: Regional baseline regressions

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	EAS	ECS	LCN	MEA	NAC	SAS	SSF
Variables	*rer*	*rer*	*rer*	*rer*	*rer*	*rer*	*rer*
gdppk	1.0095***	0.6223***	1.1065***	-0.4581*	0.7047	1.5699***	0.1675
	(0.1097)	(0.0757)	(0.2752)	(0.2510)	(0.6906)	(0.1093)	(0.1995)
govexp	0.3070***	0.1519***	0.1998***	-0.1076	-1.0568***	0.2116***	0.1245***
	(0.0639)	(0.0529)	(0.0664)	(0.1015)	(0.2320)	(0.0395)	(0.0415)
etot	0.3412***	0.0527***	0.0124	-0.1240	0.4374*	-0.0908*	0.0413**
	(0.1003)	(0.0136)	(0.0540)	(0.0919)	(0.2394)	(0.0549)	(0.0205)
L.res	0.0891***	-0.0103	0.1052***	-0.0425	-0.5427***	0.0529	0.0837***
	(0.0264)	(0.0087)	(0.0379)	(0.0274)	(0.0940)	(0.0427)	(0.0259)
etot × *L.res*	-0.1109***	-0.0175***	-0.0225	0.0184	-0.5321**	0.0185	-0.0229***
	(0.0323)	(0.0060)	(0.0196)	(0.0215)	(0.2160)	(0.0163)	(0.0073)
Constant	-1.1045**	1.0721**	-1.1372	7.3190***	4.4000	-2.3250***	3.4647***
	(0.4665)	(0.4366)	(1.2672)	(1.3201)	(3.2728)	(0.4312)	(0.8148)
Observations	247	760	323	114	38	95	304
Nb. of countries	13	40	17	6	2	5	16
R-squared	0.6595	0.3296	0.4721	0.3850	0.7476	0.7930	0.3839
RMSE	0.0933	0.0938	0.1378	0.0979	0.0614	0.0699	0.1474

Note: Bootstrapped standard errors in parentheses where 10,000 replications have been used. Fixed effects are included, but not shown. ***, **, * indicate statistical significance at the 1%, 5% and 10% levels, respectively. *L* stands for the lag operator. Source: Authors' estimates.

EZ peripheral countries run up of TARGET2 liabilities vis-à-vis the Eurosystem, and Germany is accumulating corresponding claims.[22] Consequently, finding that the buffer effect is two times larger for the ECS country group without the EZ is not surprising (see Appendix B.5). Finally, we focus on emerging and developing economies for commodity exporters, because the buffer effect may be accomplished by sovereign wealth for advanced economies. The buffer effect is also two times larger for commodity exporters than in the baseline regression.

Thus, investigating the influence of financial development may help us to understand the regional heterogeneity in the buffer effect.[23] In the following, we test for the existence of financial-development thresholds in the buffer effect in an ad hoc manner. We test the baseline regression (equation (15)) for different quartiles of financial-development indicators. We expect that the buffer effect will be observed for low levels of financial development (*FD*), financial-institution development (*FI*) and financial-market development (*FM*), because the international reserves are used as a substitute for sound financial institutions and markets. We also test the buffer effect for different levels of financial openness (*KAOPEN*), as a complementary empirical exercise. We test

[22] Steiner et al. (2019) explore the implications of permanent current account deficits within the EZ with a modified version of the Trilemma for Europe. In scenario 3, the accumulation of TARGET imbalances due to flight to safety forces the European Central Bank to pursue an expansionary monetary policy. This Trilemma generates break-up expectations on the side of investors.

[23] The regional heterogeneity could also be explained by the existence of economic development thresholds.

Table 5: The buffer effect for low levels of financial indicators

	(1)	(2)	(3)	(4)
	FD	*FI*	*FM*	*KAOPEN*
Variables	*rer*	*rer*	*rer*	*rer*
gdppk	0.814***	0.815***	0.761***	0.961***
	(0.0949)	(0.0849)	(0.103)	(0.0729)
govexp	0.135***	0.133***	0.136***	0.140***
	(0.0295)	(0.0301)	(0.0343)	(0.0276)
etot	0.0453***	0.0418***	0.0473***	0.0379**
	(0.0161)	(0.0159)	(0.0161)	(0.0173)
L.res	0.0360***	0.0383***	0.0345***	0.0317***
	(0.0117)	(0.0126)	(0.0114)	(0.0113)
etot × *L.res*	-0.0231***	-0.0221***	-0.0229***	-0.0226***
	(0.00539)	(0.00535)	(0.00550)	(0.00535)
Constant	0.575	0.557	0.800	-0.161
	(0.449)	(0.403)	(0.489)	(0.335)
Observations	1,373	1,381	1,379	1,306
Nb. of countries	80	82	83	99
R-squared	0.4497	0.4559	0.4383	0.4310
RMSE	0.1303	0.1291	0.1291	0.1224

Note: Bootstrapped standard errors in parentheses where 10,000 replications have been used. Fixed effects are included, but not shown. ***, **, * indicate statistical significance at the 1%, 5% and 10% levels, respectively. L stands for the lag operator. Source: Author's estimates.

the following equations with $k = \{FD, FI, FM, KAOPEN\}$:

$$rer_{i,t} = \mu_i + \beta_1 gdppk_{i,t} + \beta_2 govexp_{i,t} + \beta_3 etot_{i,t} + \beta_4 res_{i,t-1} + \beta_5 etot_{i,t} \times res_{i,t-1} + u_{it} \quad \text{if} \quad k < Q1k, Q2k, Q3k \tag{16}$$

$$rer_{i,t} = \mu_i + \beta_1 gdppk_{i,t} + \beta_2 govexp_{i,t} + \beta_3 etot_{i,t} + \beta_4 res_{i,t-1} + \beta_5 etot_{i,t} \times res_{i,t-1} + u_{it} \quad \text{if} \quad k \geq Q1k, Q2k, Q3k, \tag{17}$$

where $Q1k, Q2k,$ and $Q3k$ are respectively the first, second and third quartile of the financial indicator k.[24]

We can see in Tables 5 and 6 that the buffer effect is only observed for low levels of financial development. Indeed, the buffer effect is almost the same as in the baseline nonlinear regression in Table 5. The coefficient for the buffer effect fluctuates around -0.022 and is statistically significant for the observations (countries and periods) below the third quartile of the aforementioned financial-development indicators. Additionally, the coefficient for the buffer effect is no longer statistically significant for the observations (countries and periods) above the third quartile of the financial development indicators, with the notable exception of the financial-market index (FM). These first pieces of evidence may indicate the existence of threshold effects in this relationship.

After the regional regressions, the regressions at different quartiles of financial development and openness are a useful way to explore the existence of threshold effects in the relationship between the

[24]In Tables 5 and 6, we focus on the results below the third quartile and above the third quartile for the financial-development indicators to provide preliminary evidence about our main hypothesis. We also run the tests for the first quartile ($Q1k$) and the median ($Q2k$) for these indicators of financial development and openness. We do not detect any difference in the buffer effect for these ad hoc thresholds ($Q1k$ and $Q2k$).

Table 6: The buffer effect for high levels of financial indicators

	(1)	(2)	(3)	(4)
	FD	*FI*	*FM*	*KAOPEN*
Variables	*rer*	*rer*	*rer*	*rer*
gdppk	0.125*	0.00404	0.353***	0.167**
	(0.0680)	(0.0630)	(0.0812)	(0.0831)
govexp	0.0678	-0.0169	0.172**	0.131***
	(0.0604)	(0.0527)	(0.0689)	(0.0478)
etot	-0.000934	0.0137	0.0245*	0.00407
	(0.0147)	(0.0143)	(0.0145)	(0.0139)
L.res	-0.0421***	-0.0475***	-0.0310**	-0.0441***
	(0.0113)	(0.00949)	(0.0155)	(0.0137)
etot × *L.lres*	0.00661	-0.00805	-0.0182***	-0.00304
	(0.00833)	(0.00705)	(0.00573)	(0.00672)
Constant	3.843***	4.729***	2.357***	3.542***
	(0.467)	(0.384)	(0.546)	(0.477)
Observations	527	519	521	594
Nb. of countries	34	35	36	100
R-squared	0.5389	0.5534	0.4817	0.7413
RMSE	0.0701	0.0703	0.0819	0.0788

Note: Bootstrapped standard errors in parentheses where 10,000 replications have been used. Fixed effects are included, but not shown. ***, **, * indicate statistical significance at the 1%, 5% and 10% levels, respectively. *L* stands for the lag operator. Source: Author's estimates.

real exchange rate and international reserves. The preliminary results seem to validate our second type of conjectures about the use of international reserves as a substitute for soundly developed financial systems. However, the use of an ad hoc threshold like the different quartiles for the financial-development indicators may be questioned, because we cannot formally test whether the coefficients for the buffer effect are significantly different in each regime of financial development. To estimate the value for the thresholds and to formally test the difference between the coefficient in the different regimes, we estimate panel threshold regressions in the next subsection.

4.2. Threshold regressions

In Table 7, we move to the threshold regressions for the whole panel and for some of the regional panels. We test the following regression:

$$rer_{i,t} = \mu + \beta_1 etot_{i,t} I\left(res_{i,t-1} \leq \gamma\right) + \beta_2 etot_{i,t} I\left(res_{i,t-1} > \gamma\right) + \alpha_1 x_{i,t} + u_i + e_{i,t} \tag{18}$$

In these panel threshold regressions, we test the first type of conjecture where the buffer effect could be different for different levels of international reserves. Indeed, the holding of foreign reserves may mitigate the magnitude of exchange-rate adjustments triggered by capital flow movements,[25] as underlined by Aizenman and Riera-Crichton (2008). Consequently, we expect that the buffer effect will be stronger from a certain level of reserves. This threshold of reserves will be estimated from the data, as described in section 3.2. In this case, the coefficient for the buffer effect is negative after the threshold. However, we could also expect that some regions of

[25] Devereux and Wu (2022) find holdings of foreign reserves are associated with an exchange rate that is less sensitive to global shocks.

Table 7: Panel threshold regressions

Variables	(1) FULL *rer*	(2) EAS_SAS *rer*	(3) ECS *rer*	(4) LAC *rer*	(5) MEA *rer*
Estimated threshold	1.4260*	–	2.9058**	–	3.3463***
95% confidence interval	[1.2928; 1.4643]	–	[2.8780; 2.9323]	–	[3.2554; 3.3566]
gdppk	0.7004***	1.2468***	0.5618***	1.1271***	-0.2885
	(0.0523)	(0.0759)	(0.0603)	(0.2170)	(0.1931)
govexp	0.1498***	0.2434***	0.1790***	0.2500***	-0.0462
	(0.0209)	(0.0470)	(0.0420)	(0.0683)	(0.0732)
$etot.I\ (L.res \leq \gamma)$	0.0405***	-0.0265***	0.0353***	-0.0475***	-0.1378***
	(0.0106)	(0.0081)	(0.0066)	(0.0140)	(0.0223)
$etot.I\ (L.res > \gamma)$	-0.0237***	-0.2889***	-0.0208***	0.0084	-0.0217
	(0.0040)	(0.0844)	(0.0076)	(0.0315)	(0.0144)
Constant	0.9753***	-1.5495***	1.2702***	-1.0935	6.1917***
	(0.2520)	(0.3559)	(0.3449)	(1.0091)	(0.9715)
Observations	1,900	342	760	323	114
Observation below threshold	300	-	503	-	66
Number of countries	100	18	40	17	6
RMSE	0.120	0.0930	0.0922	0.139	0.0913

Note: Bootstrapped standard errors in parentheses where 10,000 replications have been used. Fixed effects are included, but not shown. ***, **, * indicate statistical significance at the 1%, 5% and 10% levels, respectively. *L* stands for the lag operator. Source: Authors' estimates.

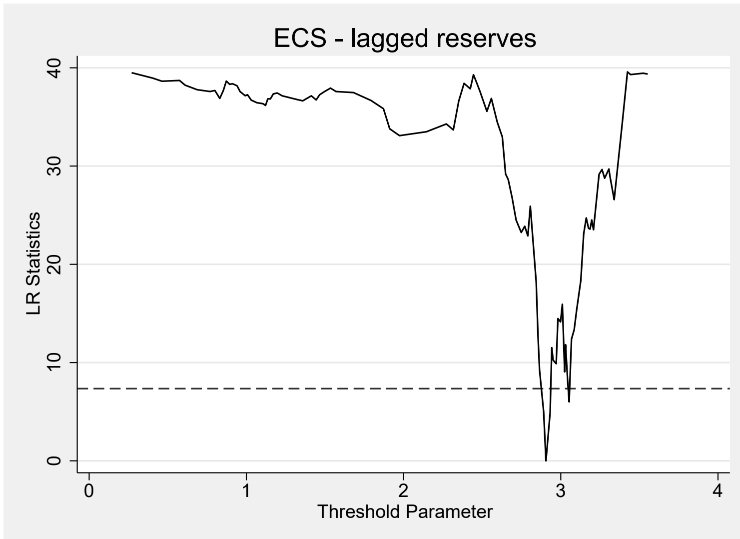
the world economy, especially the ones with high amounts of commodities, accumulate too many international reserves. In this case, we could observe the buffer effect up to a certain level of reserves. Indeed, we do not observe any buffer effect after this threshold. Here, the coefficient for the buffer effect is negative before the threshold and non-significant after the threshold.

In Table 7, we can see the full sample. Europe and Central Asia (ECS), East Asia and Pacific together with South Asia (EAS_SAS) are in the first case where the buffer effect is stronger after the estimated threshold. However, the threshold is significant at the 5% level only for the ECS region. As Figure 4 shows, the threshold value corresponds to 17.28% for the international reserves. These results mean that for observations (countries and periods) above 17.28% for the international reserves,[26] we have a statistically significant buffer effect, which is different from the effect that we have for observations below or equal to this estimated threshold. As Figure 5 shows, a majority of emerging and developing ECS countries have a mean value for their international reserves holdings superior to the value of the threshold for this region, especially after the GFC. We may conjecture that these countries have been more careful since the EZ crisis. Countries like Hungary, Croatia, Bulgaria, and the Czech Republic have substantially crossed the threshold after the EZ crisis. Still, in Table 7, we can see that in Latin America and the Caribbean (LAC) region and in the Middle East and North Africa (MEA) the accumulation of international reserves is below the threshold associated with effective real-exchange-rate mitigation, because the coefficient for the buffer effect is negative before the threshold and non-significant after the threshold.

Having established this first set of results, which provides some support for our first set of

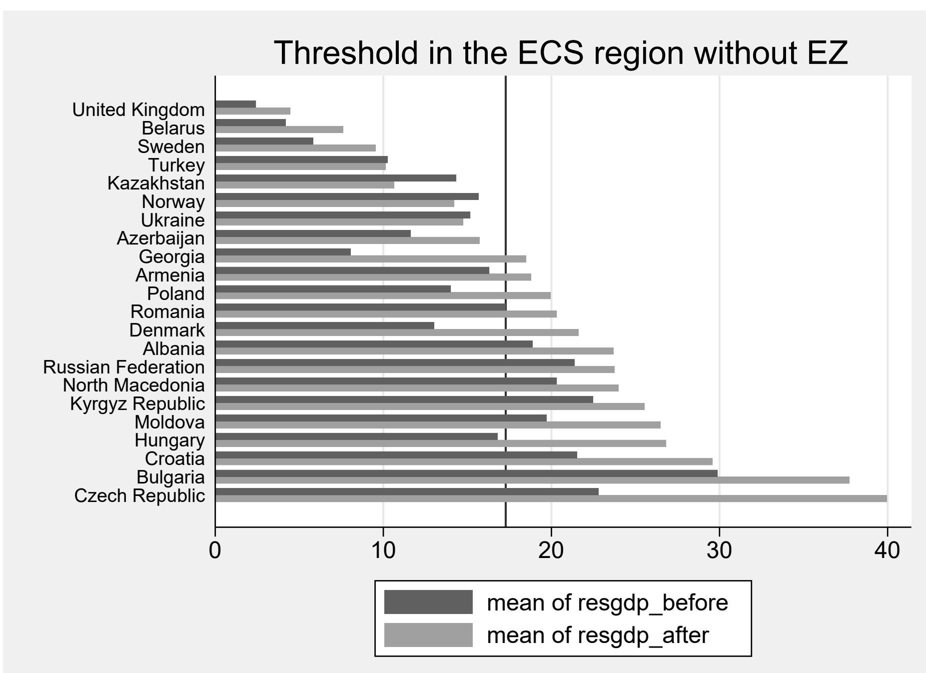
[26]In this respect, Jeanne and Ranciere (2011) introduce a model of the optimal level of international reserves for small open economies. The optimality of the estimated threshold level is beyond the scope of this paper.

Figure 4: Construction of the confidence interval in the threshold model – ECS region



Notes: The estimation for the threshold value is the point where the LR statistic is equal to zero. We obtain a value of 2.91 for the threshold. This value corresponds to a value of 17.28% for the reserves-to-GDP ratio ($\ln(1 + 100 \times x) = 2.9058 \Leftrightarrow x = 0.1728$). When the LR curve crosses the horizontal line for the first time, the lower limit of the CI is obtained. When the LR curve crosses the horizontal line for the second time, the upper limit of the CI is obtained. Source: Authors' estimations.

Figure 5: Threshold effect in the ECS region



Notes: We use a selection of emerging and developing ECS countries to compare the value of the threshold (17.28% of GDP) found in this region with the evolution of international reserves holding (mean value) before and after the GFC. Source: Authors' estimations.

conjectures, we want to explore the reasons behind the existence of these threshold effects. To do so, we explore the existence of financial-indicator thresholds. Indeed, as mentioned before in our second set of conjectures, countries may use international reserves as a substitute for a soundly developed financial system to protect themselves from negative consequences of commodity shocks. To identify the thresholds for financial-development indicators, we test the following equation with $k = \{FD, FI, FM, KAOPEN\}$ in Table 8:

$$rer_{i,t} = \mu + \beta_1 etot_{i,t} \times res_{i,t-1} I\left(k_{i,t-2} \leq \gamma\right) + \beta_2 etot_{i,t} \times res_{i,t-1} I\left(k_{it-2} > \gamma\right) + \beta_3 x_{i,t} + u_i + e_{i,t}, \quad (19)$$

Table 8: Panel threshold regressions and financial development

Variables	(1) *FD* *rer*	(2) *FI* *rer*	(3) *FM* *rer*	(4) *FM - ECS* *rer*	(5) *FMD – ECS* *rer*
Estimated threshold	–	0.4806**	–	0.0217***	0.0256***
95% confidence interval	–	[0.479; 0.4814]	–	[0.0210; 0.0220]	[0.0166; 0.0282]
gdppk	0.6930***	0.7113***	0.7140***	0.6172***	0.5944***
	(0.0552)	(0.0548)	(0.0552)	(0.0633)	(0.0633)
gov	0.1470***	0.1538***	0.1441***	0.1521***	0.1587***
	(0.0218)	(0.0217)	(0.0218)	(0.0409)	(0.0409)
$etot \times L.res.I\ (L2.k \leq \gamma)$	0.0035	-0.0096***	-0.0044***	-0.0135***	-0.0121***
	(0.0034)	(0.0014)	(0.0015)	(0.0030)	(0.0028)
$etot \times L.res.I\ (L2.k > \gamma)$	-0.0089***	0.0078***	-0.0145***	0.0144***	0.0129***
	(0.0014)	(0.0029)	(0.0022)	(0.0027)	(0.0025)
Constant	1.0207***	0.9178***	0.9325***	1.0763***	1.1718***
	(0.2654)	(0.2637)	(0.2651)	(0.3554)	(0.3552)
Observations	1,800	1,800	1,800	720	720
Observation below threshold	-	1180	-	122	123
Number of countries	100	100	100	42	42
RMSE	0.117	0.116	0.117	0.0866	0.0866

Note: Bootstrapped standard errors in parentheses where 10,000 replications have been used. Fixed effects are included, but not shown. ***, **, * indicate statistical significance at the 1%, 5% and 10% levels, respectively. L, $L2$, are the first and second lag operators, respectively. Source: Authors' estimates.

As Table 8 and Figure 6 show, we find a significant threshold effect for the financial-institution index (*FI*). For observations (countries and periods) with a low development level of their financial institutions, the buffer effect is stronger; that is, the coefficient is negative for observations inferior or equal to the threshold (around 0.48) in column (2). This central result provides some empirical support for the second types of conjectures[27]. Indeed, countries with a low development of their financial institutions may use the international reserves as a shield to deal with the negative consequences of terms of trade shocks on the real exchange rate. For the sake of completeness, we can recall that the bank credit to the private sector has a weight of only 40% in the financial-institution subcomponent of the financial-development index, as mentioned in section 3. Thus,

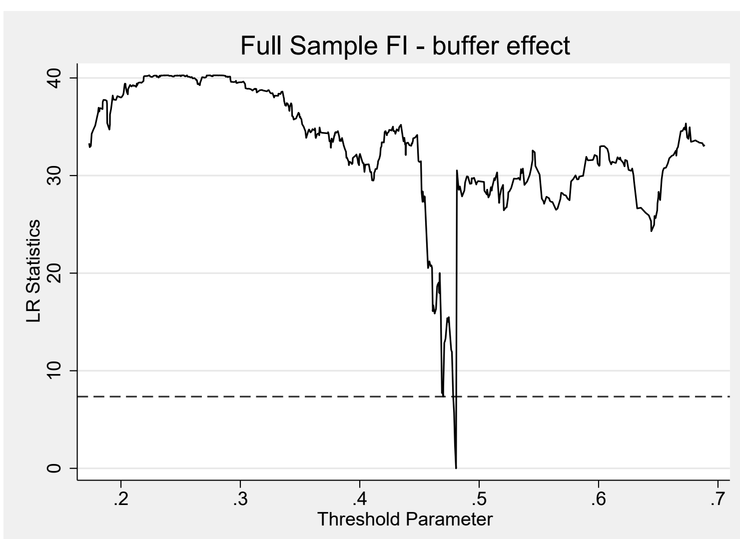
[27]We also found that the results are robust to endogeneity thanks to dynamic panel threshold models, see Appendix B.1.

the efficiency of the financial institutions, as described in Table 2, may play a crucial role in the relationship between the real exchange rate and the international reserves.

Recall that we found a significant threshold effect for the Europe and Central Asia region (ECS). Indeed, the buffer effect is stronger when the international-reserves-to-GDP ratio is above 17.28% in this region. In Table 8 and Figures 7 and 8, we investigate the potential source of the regional threshold that we found in Figure 4. In columns (4) and (5) of Table 8, the results show that for low levels of the financial-market index (*FD*) and, more convincingly, for the financial-market depth index (*FMD*), the buffer effect is stronger when the financial market is underdeveloped. To ensure completeness, we can recall that The *FMD* index summarizes the information contained in the following variables: stock market capitalization to GDP, the stocks traded to GDP, international debt securities of government to GDP, total debt securities of financial corporations to GDP, total debt securities of non-financial corporations to GDP, as mentioned in Table 2.

These last results could indicate different regions of the world economy may face different underlying factors explaining the strength of the buffer effect. Thus, combining thresholds regressions with financial indicators and regional grouping helped us discover interesting evidence about the heterogeneity of the buffer effect in the different regions of the world economy and for several dimensions of financial development. However, another important factor could influence the relationship between the real exchange rate and international reserves, the financial openness. Indeed, the buffer effect could be different for a country with unsound financial institutions and large financial openness relative to a country with a low degree of financial openness.

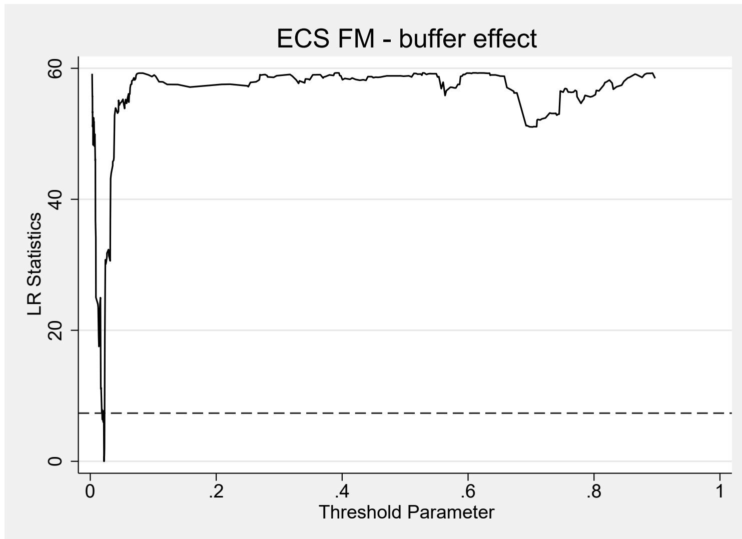
Figure 6: Construction of the confidence interval in the threshold model – FI



Notes: The estimation for the threshold value is the point where LR statistic is equal to zero. When the LR curve crosses the horizontal line for the first time, the lower limit of the CI is obtained. When the LR curve crosses the horizontal line for the second time, the upper limit of the CI is obtained. Source: Authors' estimations.

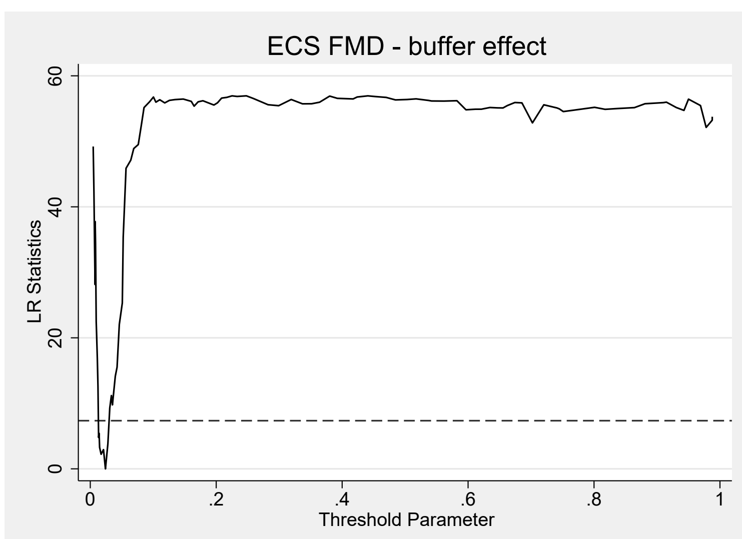
Consequently, we estimate the buffering effect of international reserves for different levels of financial openness as a complementary empirical exercise. As Table 9 shows, we find a U-shape

Figure 7: Construction of the confidence interval in the threshold model – FM, region ECS



Notes: The estimation for the threshold value is the point where LR statistic is equal to zero. When the LR curve crosses the horizontal line for the first time, the lower limit of the CI is obtained. When the LR curve crosses the horizontal line for the second time, the upper limit of the CI is obtained. Source: Authors' estimations.

Figure 8: Construction of the confidence interval in the threshold model – FMD, region ECS



Notes: The estimation for the threshold value is the point where LR statistic is equal to zero. When the LR curve crosses the horizontal line for the first time, the lower limit of the CI is obtained. When the LR curve crosses the horizontal line for the second time, the upper limit of the CI is obtained. Source: Authors' estimations.

Table 9: Panel threshold regression and financial openness

	(1)
Variables	*KAOPEN* *rer*
Estimated threshold 1	-0.1144**
95% confidence interval	[-0.1333; -0.1097]
Estimated threshold 2	0.2058**
95% confidence interval	[0.1921; 0.2073]
gdppk	0.7404***
	(0.0570)
govexp	0.1441***
	(0.0225)
$etot \times L.res.I$ ($L2.KAOPEN \leq \gamma_1$)	-0.0046***
	(0.0017)
$etot \times L.res.I$ ($\gamma_1 < L2.KAOPEN \leq \gamma_2$)	-0.0235***
	(0.0024)
$etot \times L.res.I$ ($L2.KAOPEN > \gamma_2$)	-0.0042*
	(0.0022)
Constant	0.8047**
	(0.2659)
Observations	1,764
Observation below threshold 1	870
Observation above threshold 2	825
Number of countries	98
RMSE	0.116

Note: Bootstrapped standard errors in parentheses where 10,000 replications have been used. Fixed effects are included, but not shown. ***, **, * indicate statistical significance at the 1%, 5% and 10% levels, respectively. L, $L2$, are the first and second lag operators, respectively. Source: Authors' estimates.

relationship between the buffer effect and the level of financial openness. Indeed, the buffer effect is stronger for intermediate levels of financial openness. In our estimates, we find two significant thresholds for the *KAOPEN* index. Before the first threshold γ_1, the coefficient for the buffer effect is around -0.007 and statistically significant at the 1% level. Between the first threshold γ_1 and the second threshold γ_2, the coefficient for the buffer effect is around -0.021 and statistically significant at the 1% level. The buffer effect for intermediate levels of financial openness is close to the baseline nonlinear regressions in Table 3. After the second threshold γ_2, the coefficient for the buffer effect is around -0.004 and no longer significant at the 1% level.

These results can be interpreted in the following way. After a positive terms-of-trade shock, the consequences in terms of real exchange rate appreciations could be more limited for observations (countries and periods) with a low level of financial development (inferior to γ_1). Thus, we detect a weak buffer effect in the first regime (low level of financial openness). Additionally, the consequences of a positive terms-of-trade shock in terms of real exchange rate appreciations could be more important for observations (countries and periods) with an intermediate level of financial openness and, probably, with a low level of financial development. Thus, we detect a strong buffer effect in the second regime (superior to γ_1 and inferior to γ_2). This last result provides further empirical support for our second type of conjecture, where countries may use international reserves as a shield against the consequences of terms-of-trade shocks. Finally, the consequence of a positive terms-of-trade shock could be more limited for observations (countries and periods) with a high level of financial openness. We do not detect the buffer effect at the 1% percent level in the third regime (superior to γ_2). In this last case, a high level of financial openness is typically associated with a high level of financial development. Thus, countries can deal with the consequences of a terms-of-trade shock on their exchange rate with their financial systems.

4.3. Overview of the results

In this subsection, we give an overview of the results found in our research, as we run several sub-sample analyses and use several types of econometric models. An interesting feature of our results is that the coefficient on the interaction term between international reserves, *lres*, and effective terms of trade, *etot*, is negative and significant in almost all the regressions, including the robustness analyses. As mentionned before, this coefficient measures the buffer effect of international reserves on the real exchange rate after terms-of-trade shocks. These results indicate that countries with a higher level of reserves will experience less exchange rate appreciation after a terms-of-trade shocks.

In most regressions, the buffer-effect coefficient fluctuates around the baseline value of -0.019 (see Table 3). Interestingly, the baseline value is not sensitive to the use of lagged values for the explanatory variables. The buffer-effect coefficients are close to the baseline value in the regional regressions, except for the East Asia (EAS) region where the coefficent is equal to -0.111 and significant at the one percent level (see Table 4). For the regressions with a low level of financial development in Table 5, the buffer-effect coefficients fluctuate around the baseline value and are generally higher in absolute value (around -0.022) and significant at the one percent level.

In the threshold regressions in Table 7 with the level of reserves as the threshold variable, the coefficient of interest is around the baseline value and the buffer effect is especially strong in the EAS region for high levels of reserves. In the threshold regressions with the level of financial institution development in Table 6, the value of the buffer effect is similar to the baseline for the Financial Market and the Financial Market Development indices, especially in the Europe and Central Asia (ECS) region for low level of financial market development. When considering financial openness in Table 9, we obtain a value similar to the baseline for intermediate level of financial openness.

In the Appendix, we complement our previous empirical analyses with several models that consider endogeneity of the covariates in a dynamic panel threshold model (Table B.1) and the endogeneity of the threshold variable (Table B.2). The results are very similiar to the threshold regressions in Table 8 for the buffer effect and the estimated threshold is very close (0.46 versus 0.48) for the financial institution index. With the panel Local Projections in Table B.1, we provide empirical evidence showing that the buffer effect stems only from the interaction term (and not from one of the variables in the interaction). Besides, in Figure B.2, we construct variables for the variation of *lres* and *etot* uncorrelated with the real exchange rate. At horizon $h = 0$, the buffer effect coefficient is very close to the baseline. We show that a unit shock on the interaction term has an impact on the real exchange rate for up to four years. The buffer effect of international reserves is not only a short-run phenomenon.

We also consider the presence of cross-sectional correlation in Table B.3, as these countries can be affected by common shocks and real exchange rates can exhibit co-movements at the macro level. We explore the presence of the buffer effect before and after the global financial crisis in Table B.4. The threshold effect of financial institutions is confirmed after the global financial crisis. In Table B.5, we analyse different country groupings. Unsurprisinlgy, the buffer effect is stronger for commodity exporters. In Table B.6, we consider the role of macroprudential policy. We found that the buffer effect is still confirmed when we introduce macroprudential measures. In Table B.7, we consider the effect of inward FDI and the economic size.

5. Conclusion

In an era during which financial integration is high, our paper aimed at examining the buffer effect of international reserves. After positive terms-of-trade shocks, countries can experience negative consequences in terms of real-exchange-rate appreciation and volatility. The buffer effect describes the fact that holding international reserves may help stabilize the real exchange rate after terms-of-trade shocks. We provide empirical evidence that indicates the buffer effect of international reserves is confirmed for a large macroeconomic sample of 110 countries observed between 2001 and 2020. Thanks to nonlinear regressions and panel threshold regressions, we provide empirical evidence showing the buffer effect can be observed above a threshold of international reserves in different country groups. In Europe and Central Asia, the buffer effect is observed only above a threshold of 17% for the international reserves.

Relying on new financial-development indexes developed by the IMF, we expand the literature by providing empirical support indicating the buffer effect is only observed in countries and periods where the development of financial institutions is low. Indeed, countries with a low development of their financial institutions may use the international reserves as a shield to deal with the negative consequences of terms-of-trade shocks on the real exchange rate. Thus, several countries could use international reserves as a substitute for sound financial institutions. In many emerging and developing economies, the development of sound financial institutions may be viewed as an alternative policy. We also find the buffer effect is more powerful in countries with intermediate levels of financial development. On the whole, this evidence may provide a further understanding of the consequences of international reserves holding.

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Appendix A. Data and country list

Appendix A.1. Source and definition

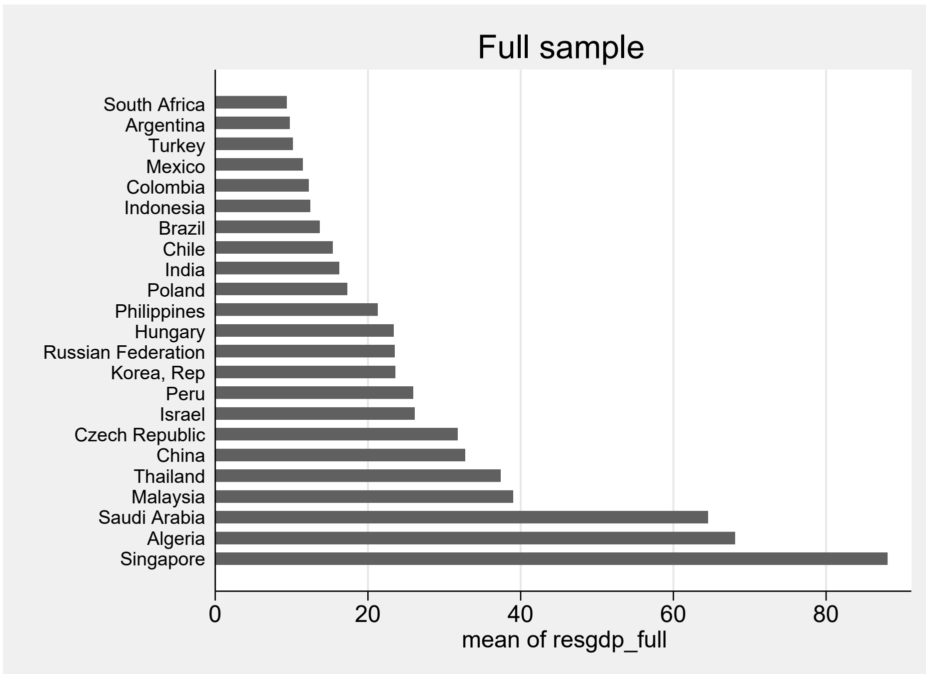
Country list: Albania, Algeria, Angola, Argentina, Armenia, Australia, Austria, Azerbaijan, Bangladesh, Belarus, Belgium, Bolivia, Botswana, Brazil, Bulgaria, Burundi, Cambodia, Canada, Chile, China, Colombia, Congo Dem Rep, Costa Rica, Croatia, Cyprus, Czech Republic, Denmark, Dominican Republic, Ecuador, Egypt Arab Rep, El Salvador, Estonia, Ethiopia, Finland, France, The Gambia, Georgia, Germany, Ghana, Greece, Guatemala, Guinea, Honduras, Hungary, India, Indonesia, Iraq, Ireland, Israel, Italy, Jamaica, Japan, Kazakhstan, Kenya, Korea Rep, Kuwait, Kyrgyz Republic, Lao PDR, Latvia, Lebanon, Lithuania, Luxembourg, Madagascar, Malawi, Malaysia, Mauritania, Mauritius, Mexico, Moldova, Mongolia, Morocco, Mozambique, Namibia, Nepal, Netherlands, New Zealand, Nicaragua, Nigeria, North Macedonia, Norway, Oman, Pakistan, Panama, Paraguay, Peru, Philippines, Poland, Portugal, Romania, Russian Federation, Rwanda, Saudi Arabia, Sierra Leone, Singapore, Slovak Republic, Slovenia, South Africa, Spain, Sri Lanka, Sweden, Thailand, Trinidad and Tobago, Tunisia, Turkey, Ukraine, United Kingdom, United States, Uruguay, Vietnam, Zambia.

Table A.1: Data sources and definitions

Variable	Definition	Source
rer	Real effective exchange rate (increase amounts to appreciation)	BRUEGEL, Darvas (2021)
to	Trade Openness as in Aizenman and Riera-Crichton (2008)	World Bank, WDI
tot	Term of Trade as in Aizenman and Riera-Crichton (2008)	World Bank, WDI
etot	Effective terms of trade as in Aizenman and Riera-Crichton (2008)	World Bank, WDI
res	International Reserves to GDP as in Aizenman and Riera-Crichton (2008)	World Bank, WDI
gdppk	GDP per capita	World Bank, WDI
govexp	Government expenditure as percent of GDP	World Bank, WDI
KAOPEN	Financial-Openness Index	Chinn and Ito (2006)
FD	Financial-Development Index	IMF, Svirydzenka (2016)
FI	Financial-Institution Index	IMF, Svirydzenka (2016)
FM	Financial-Market Index	IMF, Svirydzenka (2016)
FMD	Financial-Market Depth Index	IMF, Svirydzenka (2016)

Appendix A.2. Large holders of international reserves

Figure A.1: Large holders of international reserves as percent of GDP (full sample)



Notes: We select a sample of emerging and developing economies as in Arslan and Cantú (2019). The mean value of international reserves holding are represented. Source: Authors' calculations.

Appendix B. Additional robustness checks

Appendix B.1. Endogeneity

In this appendix, we follow Kremer et al. (2013) to investigate the possibility of threshold effects in the model. The dynamic version of the model in Equation (8) is estimated in three steps:

1. In the first step, we estimate a reduced form of the endogenous variable, $rer_{i,t-1}$, as a function of the instruments on a set of regressors restricted to 1 lag because instruments can overfit instrumented variables as shown by Roodman (2009). The endogenous variable, $rer_{i,t-1}$, is then replaced in the structural equation by the predicted values, $\widehat{rer}_{i,t-1}$.
2. In the second step, the dynamic verision of Equation (8) is estimated through least squares for a fixed threshold γ where *rer* is replaced by its predicted values from the first-step regression. We can denote the resulting sum of squares as $S(\gamma)$. This step is repeated for a strict subset of the support of the threshold variable, $FI_{i,t-1}$.
3. In the third step, the estimator of threshold value is selected as the one with the smallest sum of squared residuals, namely, $\hat{\gamma} = \underset{\gamma}{\operatorname{argmin}}\, S_n(\gamma)$. In accordance with Hansen (1999) and Caner and Hansen (2004), the critical values for determining the 95% confidence interval of the threshold value is given by

$$\Gamma = \{\gamma \, : \, LR(\gamma) \geq C(\alpha)\}$$

   where $C(\alpha)$ is the 95^{th} percentile of the asymptotic distribution of the likelihood ratio statistic $LR(\gamma)$. Once $\hat{\gamma}$ is determined, the slope of the coefficients can be estimated by the GMM for the previously used instruments and the previously estimated threshold $\hat{\gamma}$.

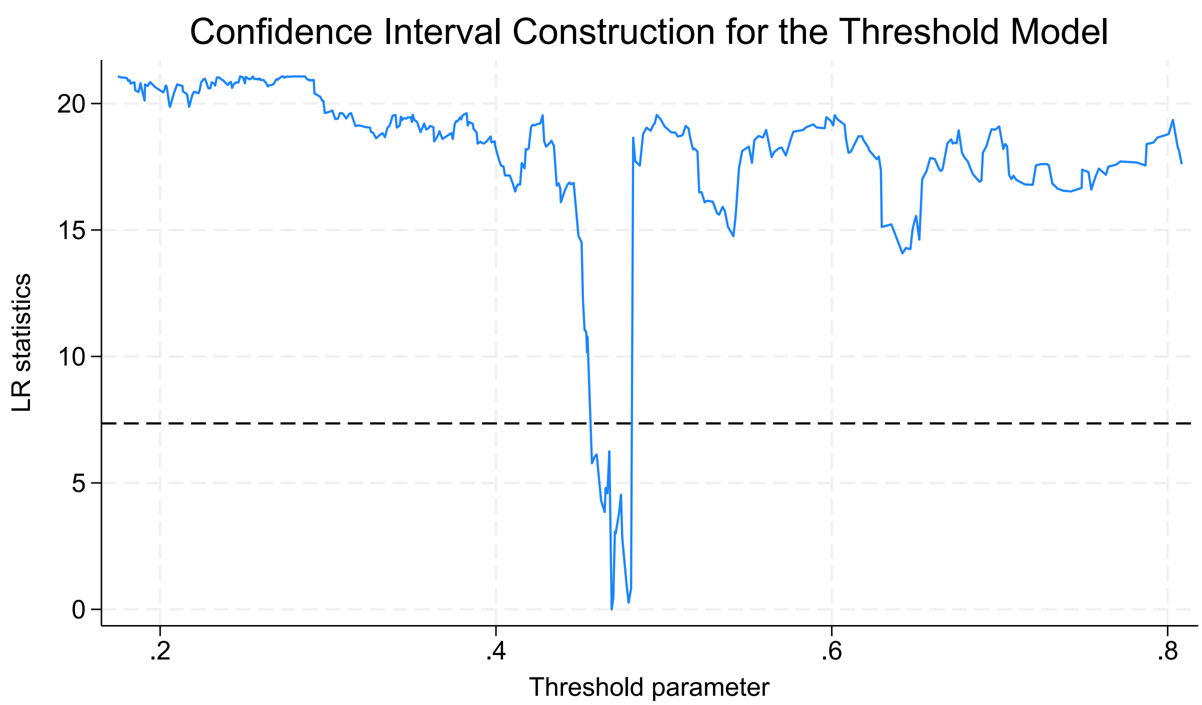
In Table B.1 and Figure B.2, the existence of the buffer effect below a threshold of financial institution development around 0.46 confirms the results obtained in Table 8. We also consider the approach of Seo and Shin (2016); Seo et al. (2019) as the threshold variables can be endogenous. In Table B.2, the results are qualitatively similar.

Table B.1: Dynamic threshold panel model (Kremer et al., 2013)

Variables	(1) $rer_{i,t}$
Estimated threshold for $FI_{i,t-2}$	0.4689
95% Confidence Interval	[0.4589; 0.4789]
Buffer effect	
$\beta_1 etot \times L.res$	-0.0104***
	(0.0042)
$\beta_2 etot \times L.res$	-0.0059
	(0.0097)
Impact of covariates	
$rer_{i,t-1}$	0.7779***
	(0.0520)
$gdppk_{i,t}$	0.0109
	(0.2589)
$govexp_{i,t}$	-0.1289**
	(0.0631)
Constant	1.3312
	(1.1149)
Observations	1800
Number of cross-sections	100
Number of instruments	51
Sargan test	$\chi^2(35)$ = 55.5557
	p-value = 0.1158
Bootstrap linearity test (p-value)	0.000

Notes: Robust standard errors in parentheses. The symbols ***, ** correspond to statistical significance at 1 and 5 percent, respectively. The non-significant time dummies has been removed with a general-to-specific approach. All differences are forward-orthogonal deviations. We use the lags of log reserves between $t-6$ and $t-9$ as instruments. The Sargan test provide support that instruments are valid. The p-value of the Bootstrap linearity test indicates that linearity is strongly rejected, 50 replications have been used.
Source: authors' calculations.

Figure B.2: Construction of the confidence interval in the dynamic model – FI



Notes: The estimation for the threshold value is the point where LR statistic is equal to zero. When the LR curve crosses the horizontal line for the first time, the lower limit of the CI is obtained. When the LR curve crosses the horizontal line for the second time, the upper limit of the CI is obtained. Source: Authors' estimations.

Table B.2: Dynamic threshold panel model (Seo and Shin, 2016)

Variables	(1) $rer_{i,t}$
	Lower regime
$rer_{i,t-1}$	0.0499
	(0.0753)
$gdppk_{i,t}$	0.6663***
	(0.1402)
$govexp_{i,t}$	0.0728**
	(0.0320)
$IRR_{i,t}$	-0.0249***
	(0.0041)
$kaopen_{i,t}$	-0.0069
	(0.0095)
Buffer effect in the lower regime	
$etot \times res$	-0.0248***
	(0.0037)
	Upper regime
Constant	0.9853
	(1.4825)
$rer_{i,t-1}$	0.9437****
	(0.1474)
$gdppk_{i,t}$	-0.8787***
	(0.2840)
$govexp_{i,t}$	-0.4535***
	(0.1181)
$IRR_{i,t}$	0.0398***
	(0.0113)
$kaopen_{i,t}$	-0.0436
	(0.0395)
Buffer effect in the upper regime	
$etot \times res$	0.0161**
	(0.0072)
Estimated threshold for $FI_{i,t}$	0.4641
95% Confidence Interval	[0.3948; 0.5333]
Observations	700
Number of cross-sections	100
Number of instruments	55
Bootstrap linearity test (p-value)	0.02

Notes: Standard errors in parentheses. The symbols ***, ** correspond to statistical significance at 1 and 5 percent, respectively. All differences are first differences. We use three-years averages for the variables to smooth business cycle fluctuations. The instruments are the lagged (three-years) averages values of the real effective exchange rates. The results are qualitatively similar without the exchange rate regime and the financial openness variables. The p-value of the bootstrap linearity test indicates that linearity is rejected, 50 replications have been used. Source: authors' calculations.

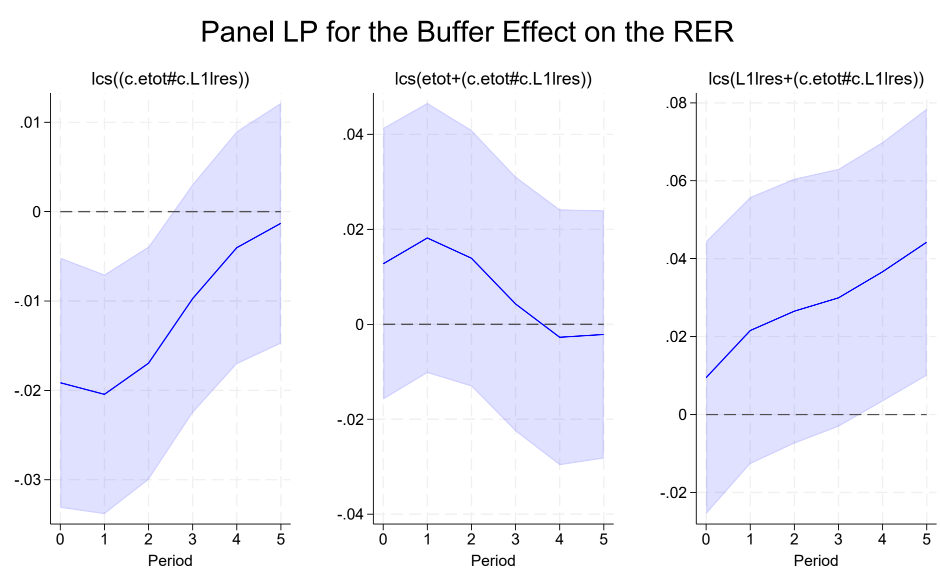
Appendix B.2. Local Projections

In this Appendix, we will use the Local Projections (hereafter LP) approach (Jorda, 2005) to provide further empirical evidence on the buffer effect of international reserves. Relying on the Stata package written by Ugarte (2023), we use panel LP to complement our baseline results in Table 3. The LP present several advantages as the estimation by single equation OLS at each horizon, a simple inference for impulse response coefficients, the effects are local to each horizon (i.e. no cross-period restrictions), the estimation of very non-linear and flexible models is straighforward, and the approach can be easily scaled to panel data. With respect to our research question, all features of the LP approach will help us to provide dynamic evidence on the buffer effect. We can formulate the LP approach as follows:

$$y_{i,t+h} = b_h S_{i,t} + \gamma_h y_{i,t-1} + \alpha \mathbf{z}_{i,t-1} + v_{i,t+h}$$
$$\text{IRF}(h) = \hat{b}_h$$

with y, is the explained variable, h, the horizon, S, the impulse variable, $\mathbf{z}$ is a vector of control variables, IRF, stands for impulse response function, and v, is the error term. In our case, the explained variable will be the real exchange rate, *rer*, and the impulse variable will be the interaction term between international reserves, *lres*, and effective terms of trade, *etot*, which captures the buffer effect. The control variables will be the same as in the baseline of Table 3, including the *de facto* exchange rate regime.

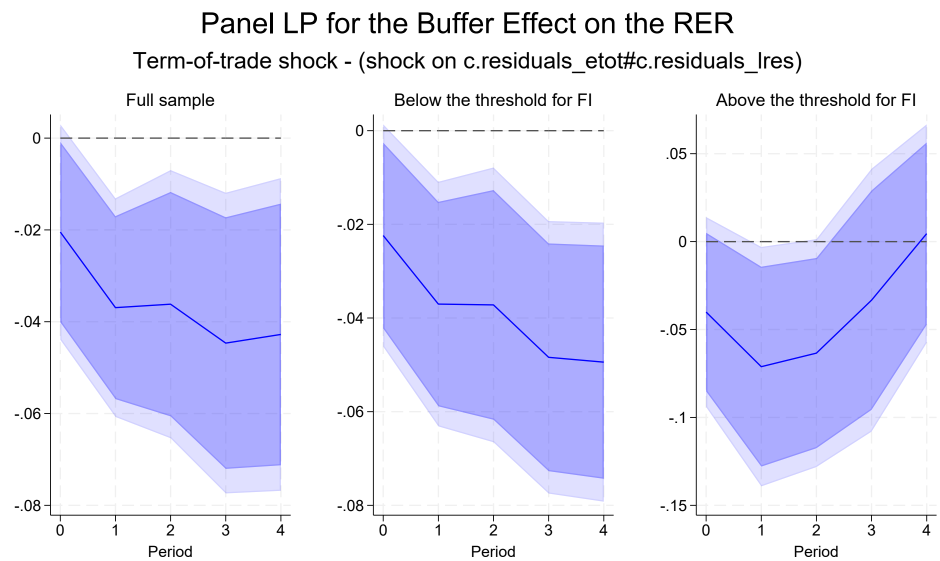
Figure B.1: Panel LP for the buffer effect on the RER



Notes: In the left panel, the unit shock is only on the interaction. In the center panel, the unit shock is on the interaction term and the effective terms of trade variable, simultaneously. In the right panel, the unit shock is on the interaction term and the international reserves variable, simultaneously. Robust standard errors. 95% confidence intervals in light blue. Source: Authors' estimations.

As a first step, we need to check whether the buffer effect comes only from the interaction term or form each variable *lres* and *etot*. As we can see in Figure B.1, we confirm that the buffer effect comes from the interaction term, and not from the variables *lres* or *etot*. At horizon $h = 0$, the coefficient is very close to the baseline of -0.019 for the buffer effect coefficient.

Figure B.2: Panel LP for the buffer effect on the RER



Notes: In the left panel, the unit shock is on the full sample. In the center panel, we use the data below the previously identifed threshold for the financial institution development. In the right panel, we use the data above the previously identifed threshold for the financial institution development. Robust standard errors. 90%, 95% confidence intervals in dark and light blue, respectively. Source: Authors' estimations.

In a second step, we construct two residual variables for *lres* and *etot* by running OLS regressions with country-fixed effects. We regress the variation of these variables on the real exchange rate, *rer*. This provides us with the variation of international reserves and effective terms of trade that are not linked to the real exchange rate. These variables are uncorrelated to the real exchange rate by construction. We will use the interaction of these residual variables as the impulse variable (i.e. the shock). Thus, the shock on the interaction term will be uncorrelated with real exchange movements. This will provide evidence robust to one possible source of endogeneity, namely, the reverse causality coming from the exchange rate on international reserves and effective terms of trade. As you can see in Figure B.2, the results of the IRF for the buffer effect are very close to the baseline results of Table 3 at horizon $h = 0$ (-0.020) providing further evidence about the robustness of our results. Furthermore, one of the main conjectures of this paper is confirmed: the buffer effect is stronger for countries / periods with low financial institution development.

Appendix B.3. Cross-sectional correlation

In this Appendix, we follow Sul (2019) to consider the influence of cross-sectional correlation on our estimates. In Table B.3, the introduction of year fixed effects or of the cross-sectional mean of the real exchange rate are sufficient to eliminate the potentially strong cross-sectional correlations, as witnessed by the large p-value obtained in the CD test of Pesaran (2014). These two approaches can be valid in the case of homogeneous factor loadings for the common factors. In case of heterogeneous factor loadings, we use the "iterative interactive fixed-effect" model introduced by Bai (2009). This factor-augmented panel regression produces consistent estimates in the presence of heterogeneous factors loading.

Table B.3: Factor augmented panel regressions

	(1)	(2)	(3)
	mean rer	*year effects*	*iterative interactive*
Variable	*rer*	*rer*	*rer*
gdppk	0.6946***	0.6957***	0.9216***
	(0.0657)	(0.0644)	(0.0824)
govexp	0.0722**	0.0759**	0.0956***
	(0.0296)	(0.0307)	(0.0184)
etot	0.0094	0.0105	0.0207**
	(0.0100)	(0.0104)	(0.0097)
L.res	0.0046	0.0038	-0.0010
	(0.0077)	(0.0078)	(0.0069)
etot $\times$ *L.res*	-0.0103***	-0.0107***	-0.0141***
	(0.0035)	(0.0036)	(0.0035)
Constant	-2.8202***	1.0916***	0.1177
	(0.3857)	(0.3341)	(0.3871)
Observations	1,900	1,900	1,919
CD test (p-value)	-0.851 (0.395)	0.538 (0.590)	-0.835 (0.404)
RMSE	0.106	0.106	0.0810

Note: The cross-sectional mean of lreer and the year fixed effects are significant at the 1% level. We use bootstrapped standard errors for (1) and (2) where 10,000 replications were used. The null hypothesis in the CD test is cross-sectional independence / weak cross-sectional dependence, and the alternative is strong cross-sectional dependence. In the CD test, a low p-value indicates some (strong) correlation between countries. ***, **, * indicate statistical significance at the 1%, 5% and 10% level, respectively. L stands for the lag operator. Source: Authors' estimates.

Appendix B.4. Before and after the GFC (Baseline and Financial institutions thresholds)

In this appendix, we want to test the buffer effect before and after the GFC. In the Table B.4, we find that the buffer effect is stronger after the GFC. The threshold effect for the financial-institution indicator is also confirmed after 2008.

Table B.4: Before and after the Global Financial Crisis

Variables	(1) FI – after GFC *rer*	(2) before GFC *rer*	(3) after GFC *rer*
gdppk	0.6241***	0.9524***	0.5712***
	(0.0778)	(0.1460)	(0.1549)
govexp	0.0578**	0.0245	0.0605
	(0.0272)	(0.0426)	(0.0447)
etot		0.0074	0.0288**
		(0.0260)	(0.0133)
L1.res		0.0068	0.0052
		(0.0174)	(0.0110)
$etot \times L.res$		-0.0162*	-0.0153***
		(0.0098)	(0.0051)
$etot \times L1.res.I\ (L2.FI \leq \gamma)$	-0.0083***		
	(0.0016)		
$etot \times L1.res.I\ (L2.FI > \gamma)$	0.0098***		
	(0.0029)		
Constant	1.6149***	0.0593	1.8404**
	(0.3674)	(0.6918)	(0.8013)
Observations	1200	700	1,200
RMSE	0.0894	0.0884	0.0909

Note: Bootstrapped standard errors in parentheses where 10,000 replications have been used. Fixed effects are included, but not shown. ***, **, * indicate statistical significance at the 1%, 5% and 10% level, respectively. $L1$ and $L2$ stand for the lag operator. Source: Authors' estimates.

Appendix B.5. Country groups (OECD, non-OECD, ECS without EZ, commodities after 2008)

In this Appendix, we test the baseline regression for different groups of countries. In Table B.5, we note that OECD countries[28] may be subject to the constraints imposed by the Trilemma: greater capital mobility hinders the potency of real-exchange-rate effects associated with countercyclical management of IR. For the same reasons, we remove eurozone countries from the ECS country group. Following Aslam et al. (2016), a country is classified as a commodity exporter (using data available for 1962–2014) if (1) commodities constitute at least 35% of its total exports and (2) net exports of commodities are at least 5% of its gross trade (exports plus imports) on average. We focus on commodity-exporting emerging markets and developing economies, because a fair share of commodity countries manage sovereign wealth funds, Norway being a prime example. For these countries, the buffer effect may be accomplished via sovereign wealth funds management. Finally, we find the buffer effect is stronger in non-OECD countries, in ECS without the eurozone and in non-OECD commodities countries after the GFC in 2008.

Table B.5: Baseline regressions for different country groups

Variables	(1) OECD *rer*	(2) Non-OECD *rer*	(3) ECS without EZ *rer*	(4) Commodities *rer*
gdppk	-0.1259*	0.8299***	0.8939***	0.9816**
	(0.0650)	(0.0963)	(0.1305)	(0.3946)
govexp	-0.0424	0.1413***	0.0466	0.1563
	(0.0711)	(0.0296)	(0.0703)	(0.1091)
etot	-0.0584***	0.0359**	0.1171**	0.1606**
	(0.0113)	(0.0153)	(0.0574)	(0.0654)
L.res	-0.0939***	0.0497***	0.1012***	0.0170
	(0.0119)	(0.0113)	(0.0227)	(0.0491)
etot × *L.res*	0.0482***	-0.0198***	-0.0446**	-0.0404**
	(0.0075)	(0.0053)	(0.0180)	(0.0206)
Constant	5.5121***	0.4632	-0.1427	-0.3075
	(0.4758)	(0.4553)	(0.6197)	(1.8690)
Observations	532	1,368	418	204
R-squared	0.4612	0.4780	0.4634	0.6715
RMSE	0.0741	0.129	0.104	0.117

Note: Bootstrapped standard errors in parentheses where 10,000 replications have been used. Fixed effects are included, but not shown. ***, **, * indicate statistical significance at the 1%, 5% and 10% respectively. *L* stands for the lag operator. Source: Authors' estimates.

[28] We select members who have been in the OECD for at least 20 years to match our sample period.

Appendix B.6. Controls for macroprudential measures

In this Appendix, we consider the following series to control for the influence of macroprudential measures (Cerutti et al., 2017) on the buffer effect, and we test the baseline regression in Table 2 of the paper. Indeed, macroprudential measures may reduce the need for international reserves. We took the value of the last quarter for the cumulative macroprudential indicators to match quarterly data with yearly data. The data are available only for 56 countries out of 110. We have the following variables for macroprudential measures: *cum_sscb_res*, cumulative change in sector specific capital buffer: real estate credit; *cum_sscb_cons*, cumulative change in sector specific capital buffer: consumer credit; *cum_sscb_oth*, cumulative change in sector specific capital buffer: other sectors; *cum_cap_req*, cumulative change in capital requirements; *cum_concrat*, cumulative change in concentration limit; *cum_ibex*, cumulative change in interbank exposure limit; *cum_ltv_cap*, cumulative change in the loan-to-value cap; *cum_rr_foreign*, cumulative change in reserve requirements on foreign currency-denominated accounts; *cum_rr_local*, cumulative change in reserve requirements on local currency-denominated accounts; *cum_sscb*, cumulative change in the aggregate sector-specific capital buffer instrument; *cum_PruC*, sum of the cumulative version of the nine instruments.

Our results in Table B.6 are robust to the inclusion of macroprudential measures. Interestingly, the cumulative change in the loan-to-value cap is the only macroprudential indicator that significantly weakens the buffer effect.

Table B.6: Baseline regressions with macroprudential indicators

Variable	(1) *rer*	(2) *rer*	(3) *rer*	(4) *rer*	(5) *rer*	(6) *rer*	(7) *rer*	(8) *rer*	(9) *rer*	(10) *rer*
gdppk	0.6141***	0.6599***	0.6026***	0.6676***	0.3618***	0.6566***	0.7986***	0.6529***	0.6674***	0.6671***
	(0.0644)	(0.0743)	(0.0673)	(0.0726)	(0.1108)	(0.0718)	(0.0722)	(0.0713)	(0.0632)	(0.0708)
govexp	0.1163***	0.1345***	0.1352***	0.1366***	0.2537***	0.0517	0.1071**	0.1368***	0.1364***	0.1344***
	(0.0341)	(0.0343)	(0.0321)	(0.0418)	(0.0649)	(0.0498)	(0.0467)	(0.0356)	(0.0421)	(0.0393)
etot	0.0707***	0.0671***	0.0695***	0.0649***	0.0213	0.0761***	0.0537***	0.0663***	0.0651***	0.0671***
	(0.0148)	(0.0173)	(0.0148)	(0.0151)	(0.0236)	(0.0159)	(0.0159)	(0.0159)	(0.0162)	(0.0173)
L.res	-0.0108	-0.0043	-0.0056	-0.0023	-0.0399**	-0.0126	-0.0067	-0.0030	-0.0022	-0.0044
	(0.0104)	(0.0099)	(0.0094)	(0.0093)	(0.0162)	(0.0094)	(0.0090)	(0.0091)	(0.0101)	(0.0085)
etot × *L.res*	-0.0222***	-0.0233***	-0.0229***	-0.0233***	-0.0143*	-0.0247***	-0.0201***	-0.0230***	-0.0234***	-0.0236***
	(0.0051)	(0.0056)	(0.0054)	(0.0052)	(0.0084)	(0.0055)	(0.0053)	(0.0059)	(0.0054)	(0.0055)
cum_PruC	0.0072***									
	(0.0011)									
cum_sscb		0.0051								
		(0.0036)								
cum_rr_local			0.0099***							
			(0.0029)							
cum_rr_foreign				-0.0010						
				(0.0034)						
cum_ltv_cap					0.0107**					
					(0.0043)					
cum_concrat						0.0337***				
						(0.0071)				
cum_cap_req							0.0304***			
							(0.0044)			
cum_sscb_oth								0.0137		
								(0.0106)		
cum_sscb_cons									-0.0051	
									(0.0126)	
cum_sscb_res										0.0068
										(0.0048)
Constant	1.2630***	0.9821***	1.2734***	0.9354**	2.2131***	1.2331***	0.3586	1.0077***	0.9367***	0.9477**
	(0.3306)	(0.3706)	(0.3451)	(0.3770)	(0.6064)	(0.3902)	(0.3754)	(0.3720)	(0.3307)	(0.3752)
Observations	952	952	952	952	427	947	951	952	952	952
R-squared	0.4645	0.4397	0.4473	0.4385	0.5796	0.4611	0.4710	0.4399	0.4386	0.4396
RMSE	0.0981	0.100	0.0997	0.101	0.0844	0.0984	0.0975	0.100	0.100	0.100

Note: Bootstrapped standard errors in parentheses where 10,000 replications have been used. Fixed effects are included, but not shown. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% level, respectively. L stands for the lag operator. Source: Authors' estimates.

Appendix B.7. Trade, financial openness, and economic size (Trade openness, FDI inward stocks, GDP size)

In this Appendix, we want to explore the interaction between trade and financial barriers in Table B.7. We do find that economic size reduces the buffer effect. Also, the stock of inward FDI is associated with a reduction of the buffer effect.

Table B.7: Baseline regressions, openness, and economic size

Variables	(1) FDI inward stocks *rer*	(2) Size (real GDP) *rer*	(3) Trade openness *rer*
gdppk	0.6842***	0.1705**	0.5965***
	(0.0643)	(0.0677)	(0.0670)
govexp	0.1537***	0.0809***	0.1456***
	(0.0262)	(0.0251)	(0.0252)
etot × *L.res*	-0.0086***	-0.0447***	-0.0201
	(0.0015)	(0.0164)	(0.0153)
fdistocks	0.0001***		
	(0.0000)		
etot × *L.res* × *fdistocks*	-0.0000**		
	(0.0000)		
lconsgdp		0.2801***	
		(0.0164)	
etot × *L.res* × *lconsgdp*		0.0016**	
		(0.0007)	
L2lto			-0.1669***
			(0.0270)
etot × *L.res* × *L2lto*			0.0041
			(0.0041)
Constant	1.0448***	-3.4838***	2.0800***
	(0.3190)	(0.3919)	(0.3961)
Observations	1,760	1,900	1,900
R-squared	0.4549	0.5772	0.4852
RMSE	0.117	0.107	0.118

Note: Bootstrapped standard errors in parentheses where 10,000 replications have been used. Fixed effects are included, but not shown. ***, **, * indicate statistical significance at the 1%, 5% and 10% respectively. $L1$, $L2$, stands for the first and second lag operators, respectively. Source: Authors' estimates.