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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. 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.

## 1. Introduction

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

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Note that the holding of international reserves is not a free lunch for countries (Ben-Bassat and Gottlieb, 1992; Rodrik, 2006; Korinek and Serven, 2016), however, policy makers are interested in making use of this tool to cope with external finance shocks. Given the divergence of monetary and trade policies, several groups of countries have followed their own approaches to manage their macroeconomic indicators. Therefore, motivated by this broader scope, this study evaluates the relationship between the holding of international reserves, terms of trade shocks, and real exchange rates. Similarly to Aizenman and Riera-Crichton (2008), our study also assumes that these effects may be of first order magnitude for developing economies. In doing so, our study attempts to disentangle these relationships by clustering different country groups.

This study is different from the existing literature for two main reasons. First, although we draw on the global sample for the baseline analyzes, the characteristics of several country groups are disaggregated. Since Aizenman and Riera-Crichton (2008) and Aizenman et al. (2012) claimed that 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 that 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 examines the threshold under the other constraints to strengthen the understanding of different contexts.

We summarize our findings in two main points. First, 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. Second, 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. These two findings are consistent with the existing literature that debates reserve holdings and its substitutability with capital controls (Obstfeld et al., 2010; Aizenman and Riera-Crichton, 2008; Alberola et al., 2016; Steiner, 2017; Cezar and Monnet, 2023).

This study contributes to the existing literature by investigating the buffer effect in an era of financial integration. Whereas the previous study emphasizes the buffer effect, our study expands this debate and discusses the existence of a complementarity between the holdings of international reserves and the development of sound financial institutions. 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 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, 2008; Aizenman et al., 2012; Al-Abri, 2013; Coudert et al., 2015; Adler et al., 2018; Aizenman and Jinjarak, 2020).

Our empirical investigation aims to explore the relationship between the real exchange rate and international reserves. In this regard, Aizenman and Riera-Crichton (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 that 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 that financial depth significantly reduces the role of reserves as a shock absorber for developing economies (Aizenman and Riera-Crichton, 2008).

In our empirical investigation, we deepen these last results of Aizenman and Riera-Crichton (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 thanks to the aggregated and disaggregated financial indices 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 that 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 the negative consequences of terms-of-trade shocks.

The mitigation effect of terms-of-trade shocks may result from reducing real exchange rate adjustments due to capital flows; thus reducing the probability of a full-blown financial crisis, as explained by Aizenman and Riera-Crichton (2008). Dominguez (2010) explains that firms in emerging countries with underdeveloped financial markets tend to rely excessively on external financing, using a simple model of private-sector external underinsurance. Then, private firms will exhibit underinsurance against future capital shortfall. In these countries, governments may accumulate *ex ante* reserves to mitigate this exposure.

For most emerging countries, the rapid increase in sovereign spreads and the appreciation of the dollar at times of 'flight to safety' generates a doom loop. Proper management of international reserves *ex ante* (hoarding reserves in good time, i.e., booming exports, strong terms of trade and higher export revenues), and selling them at times of 'flight to safety' reduces the probability of bankruptcy of the corporate sector (a major concern in S. Korea and other countries in the Asian Crisis and similar pressure during the GFC affecting most emerging countries), as explained in Aizenman and Jinjarak (2020). Hence, both consumption smoothing, as well as investment smoothing, reduces corporate defaults and explains the *ex ante* GFC and the *ex post* GFC hoarding and selling of international reserves during the GFC. Dominguez et al. (2012) discuss these issues in more details, noting also that once reserves fall below the threshold, net new capital inflows abruptly end, leading to debt rollover problems and capital flight. These capital flow reversals can, in turn, increase the pace of reserve depletion.

Aizenman et al. (2012) focus on the commodity terms of trade shocks[1] for several Latin American countries. They recall that the buffer stock approach to international reserves goes back to the Bretton Woods era.[2] They use panel data over the period 1970-2009, with quarterly data, to understand (a) how the real exchange rate reacts to commodity terms of trade shocks in several Latin American countries and (b) the influence of international reserves in this adjustment. They use two versions for their error correction model. In the first case, the international reserve-to-GDP ratio is a long-term determinant of the real exchange rate. In the second, the international reserve-to-GDP ratio also affects the adjustment speed toward 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 will push toward a real depreciation, thus softening the real appreciation. In this sense, international reserves could influence the speed of adjustment.

In general, the mitigation effect of international reserves after terms of trade shocks and the reduction in the adjustment speed are confirmed for most countries. Interestingly, they also considered 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 the 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), 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.

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 used panel-cointegration techniques along with panel threshold regression techniques to examine the relationship between the real exchange rate and the terms of trade. Although 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 January 1994 to December 2012 for the short-run impact on terms of trade volatility. Interestingly, they find that 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 commodity and financial volatility is high (measured by the S&P GSCI and the VIX) and for advanced commodity exporters, namely Australia, Canada, and Norway.

Adler et al. (2018) find the holding of international reserves is a key tool to smooth 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 that 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, there are no statistical differences between countries with low and high levels of international reserves for boom episodes, where restrictions on reserve accumulation are absent. Unfortunately, they do not report the results for the real exchange rate. However, we can reasonably infer that these asymmetrical effects of international reserves are also present in the exchange-rate adjustment.

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,[3] 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 sizable benefits, especially for countries with highly volatile real exchange rate and large sovereign spread.

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

[2] The prevailing rule of thumb for an adequate level of international reserves was four months of imports.

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

**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 |
| *etot* | 2,200 | -0.028 | 1.305 | -6.817 | 9.552 |
| *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.

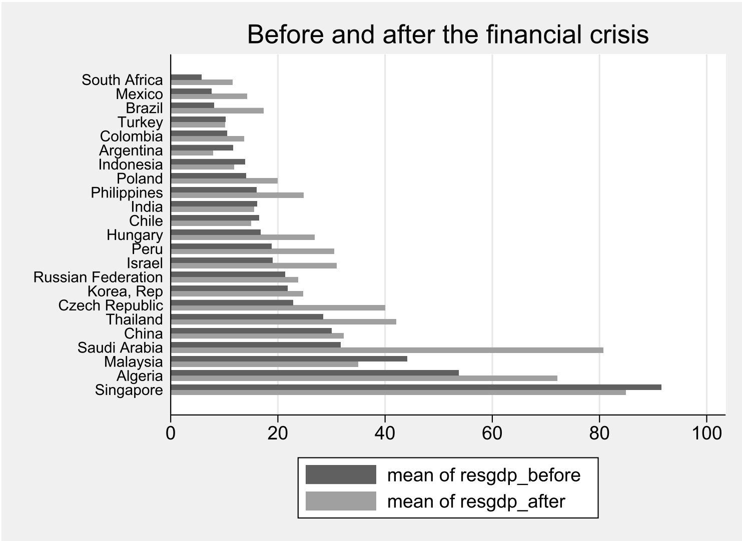


**Fig. 1.** Large holders of international reserves as percent of GDP (before and after the GFC). *Notes:* We select a sample of emerging and developing economies as in Arslan and Cantú (2019). We split the sample into two subperiods, 2001-2007 and 2010-2020, to observe the consequences of the GFC on reserves accumulation. Source: Authors' calculations.

## 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. Subsequently, the effective terms of trade, *etot* are obtained by multiplying *to* by *tot*. We present the descriptive statistics in Table 1.

In Fig. 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. We see that several emerging countries hold a large amount that represents a large 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. Most countries in this group have more reserves after the financial crisis. For Eastern European countries, we can note that the Czech Republic and Hungary have largely increased their holding of international reserves. Furthermore, for 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.

To provide a better understanding of the relationship between the real exchange rate and international reserves, we used two types 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 the intensity of capital controls. For clarity purposes, providing some explanations about the construction of these financial indexes may be useful.

First, we can briefly describe 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 et al. (2015), the effect of financial development on growth is non-linear and uneven across sectors. To capture several dimensions of financial development, Svirydzenka (2016) constructs a series of financial development indexes aimed at capturing the financial development of institutions and markets in terms of depth, access, and efficiency.[4]

In addition, we can list 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 level of activity of capital markets).

Second, we briefly mention that the *KAOPEN* index is a measure of financial openness (i.e., the openness of the capital account). Introduced by Chinn and Ito (2002), this index aims to measure the extensity of 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*).

### 3.2. Methodology

In addition to panel nonlinear regressions with interaction terms, we use panel threshold regressions introduced by Hansen (1999) in the empirical literature.[5] We consider the following model to investigate the possibility of nonlinearities and 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):

$$rer_{i,t} = \mu_i + \beta_1 etot_{i,t} + \beta_2 res_{i,t-1} + \beta_3 etot_{i,t} \times res_{i,t-1} + \alpha' \mathbf{x}_{i,t} + u_{i,t}, \tag{1}$$

where subscripts $i = 1, \dots, n$ represent the country and $t = 1, \dots, T$ index the time. $\mu_i$ is the country-specific fixed effect, and $u_{it}$ is the error term. The variables involved are presented in Table A.1 and in the previous section. Indeed, *rer* is the real effective exchange rate, *etot* stands for the effective terms of trade, and *res* represents the international reserves holding. The vector of control variables, $\mathbf{x}$, includes the GDP per capita and government expenditure as a percent of GDP.

From a mathematical point of view, threshold models are nested into nonlinear models. However, they provide a better understanding of the different regimes as they estimate thresholds. From an economic perspective, the threshold model can be interesting as it provides a more meaningful economic interpretation for the different regimes. That is especially true when we have an interaction between three different variables.[6] Interpreting this kind of interaction becomes increasingly complex. Thus, in our case, when we want to analyze whether the buffer effect is more powerful in countries with a less developed financial system, it seems more practical to use a threshold model and an interaction term rather than an interaction between three variables in our regressions.

$$rer_{i,t} = \mu_i + \theta_1 etot_{i,t} \times res_{i,t-1} I\left(k_{i,t-2} \le \gamma\right) + \theta_2 etot_{i,t} \times res_{i,t-1} I\left(k_{it-2} > \gamma\right) + \alpha' \mathbf{x}_{i,t} + u_{i,t}. \tag{2}$$

where $I(.)$ is an indicator function that indicates the regime defined by the threshold variable, and $\gamma$ is the value of the estimated threshold. With $k = \{FD, FI, FM, KAOPEN\}$, the different indicators of financial developments mentioned in the previous Section, *FD*, is the aggregated financial development index, *FI*, stands for the financial institution index, *FM*, is the financial market index and *KAOPEN* is the Chinn-Ito index.[7]

## 4. Empirical results

In our empirical approach, we use the regressions presented in the previous section to capture the ability of international reserves to deal with the consequence of the increase in terms of trade on the real effective exchange rate.[8] First, we will show the results for the panel nonlinear regressions. Then, 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. Finally, we test the intensity of the buffer effect for different levels of financial developments. In these cases, the threshold variable will be the financial development

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

[5] 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.

[6] When you have an interaction term between two variables, the marginal effects can be visualized in a 3-D plane, but when you have an interaction term between three variables, it is no longer possible to visualize the interaction, as we live in a world with three dimensions of space.

[7] The panel threshold variable has to be exogenous. Using two lags ensures that the financial institution indicator is not affected by reverse causality.

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

**Table 2**
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 or present values for all the explanatory variables. Source: Authors' estimates.

indexes. Our central hypothesis is that several countries could use international reserves as a substitute for sound financial institutions to deal with the consequences of terms-of-trade shocks on the real effective exchange rate. Lastly, we provide empirical evidence with the *KAOPEN* index of financial openness (Chinn and Ito, 2006) as the threshold variable to verify the complementarity between capital controls and international reserves, as underlined by Steiner (2017) and Cezar and Monnet (2023).

### 4.1. Nonlinear regressions

We test the baseline regressions (see Table 2), where the buffer effect is captured by the negative coefficient on the interaction term between lagged reserves and terms of trade.[9] 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 in 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.[10] Second, the positive coefficient in 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 nontradable goods.[11] 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 induces an increase in the 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 that the income effect is stronger than the substitution effect.[12] Fourth, the positive coefficient on the lagged reserves shows that holding reserves may lead to real appreciation.

Finally, the interaction coefficient is the main coefficient of interest in this study.[13] It captures the effect of the interaction between 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. 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

[9] 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 that 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.

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

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

[12] See De Gregorio and Wolf (1994) and Mendoza (1995), for example.

[13] The average marginal effect of *res* quickly becomes non-significantly different form zero, after the mean value confirming the buffer effect for various levels of *etot*. Similar results are found for the baseline estimate and the average marginal effects of international reserves are found with the *tot* variable.

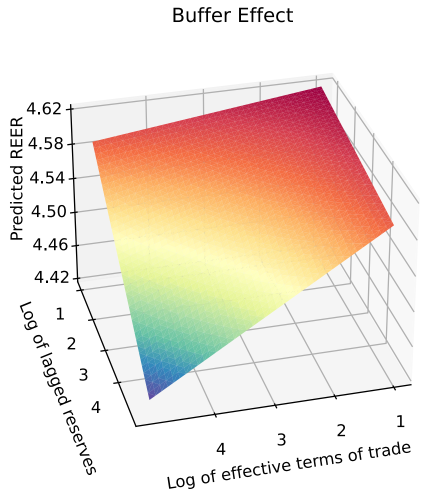


**Fig. 2.** 3-D plot for the buffer effect. *Note:* Blue areas indicate that the buffer effect (i.e., the mitigation of real exchange rate appreciation after a terms-of-trade shock) is stronger when the level of reserves is higher. We include year-fixed effects in the regressions. The results are similar without the year-fixed effects. Source: Authors' estimates. (For interpretation of the colors in the figure(s), the reader is referred to the web version of this article.)

the coefficient of the interaction term is negative and statistically significant. Indeed, we find that the buffer effect is statistically significant for this large macroeconomic panel over the last 20 years (see Table 2).[14]

In Fig. 2, we provide the 3-D plot to illustrate the interaction between the effective terms of trade and the lagged international reserves.[15] Visualizing the interaction between two continuous variables can 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. On the one hand, we can see in Fig. 2 that the effect of terms of trade shocks is stronger when countries have a low level of reserves (red areas). On the other hand, we can see in Fig. 2 that the effect of trade shock terms is weaker when countries have a high level of reserves (blue area).

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. In fact, 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. 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 international reserves as a shield to deal with the negative consequences of terms of trade shocks on the real exchange rate.

To quantify regional heterogeneity, we tested baseline regressions for various country groups based on the World Bank's classification in Table 3. On the whole, the R-squared and RMSE are quite close to those of the baseline regression. The coefficients have expected signs for GDP per capita, government consumption, effective terms of trade, and 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 did 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 found the highest value for the buffer effect in Asia. This regional heterogeneity of the buffer effect may be due to different levels of developments in the financial markets and financial institutions. Furthermore, we can argue that the level of financial openness can also greatly influence the buffer effect.

In Appendix B.6, 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 did not find any buffer effect for the OECD countries, we did find a similar buffer effect (-0.0198) for the non-OECD countries to those in the baseline regression in Table 2. Because the ECS countries group includes the euro area, controlling for the presence of these countries in this group may be worthwhile. Indeed, the policies of the European Central Bank provide buffers for most of the EZ countries, especially to deal with intrazone capital flights through the TARGET 2 system (Cheung et al., 2020). Instead of running down reserves like Mexico in 1994, EZ peripheral countries run up of TARGET2 liabilities vis-à-vis the Eurosystem, and Germany is accumulating corresponding claims.

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.

[14] In Appendix B.4, we provide empirical evidence showing that our results are robust when common factors (with homogeneous or heterogeneous factor loadings) are considered. The results are very similar when we lag all variables.

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

**Table 3**
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) |
| $etot \times L.res$ | -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.

**Table 4**
Panel threshold regressions and financial development.

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| | *FD* | *FI* | *FM* | *FM - ECS* | *FMD – ECS* |
| *Variables* | *rer* | *rer* | *rer* | *rer* | *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.

## 4.2. *Threshold regressions*

Having established this first set of results, which provides some support for our main set of conjectures, we are keen 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 above, countries can use international reserves as a substitute for a well-developed financial system to protect themselves from the negative consequences of commodity shocks.

As Table 4 and Fig. 3 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 to or equal to the threshold (around 0.48) in column (2). This central result provides some empirical support for our main conjectures.[16] In fact, countries with low development of their financial institutions may use international reserves as a shield to deal with the negative consequences of terms of trade shocks on the real exchange rate. For completeness, we recall that

[16] We also found that the results are robust to endogeneity thanks to dynamic panel threshold models and to local projections, see Appendix B.2 and Appendix B.3.

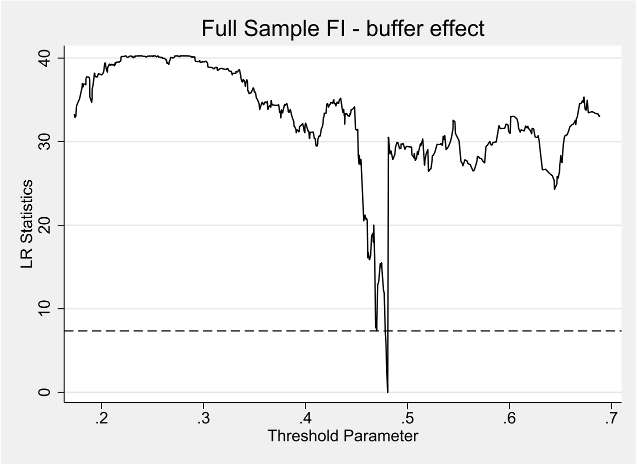


**Fig. 3.** Construction of the confidence interval in the threshold model – FI. *Notes:* The estimation for the threshold value is the point where LR statistic is equal to zero. When the LR curve crosses the horizontal line for the first time, the lower limit of the CI is obtained. When the LR curve crosses the horizontal line for the second time, the upper limit of the CI is obtained. Source: Authors' estimations.

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

In Table 4, we investigate the potential source of the regional threshold. In columns (4) and (5) of Table 4, 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.

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.

Consequently, we estimate the buffering effect of international reserves for different levels of financial openness. As Table 5 shows, we find a U-shape 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 $\gamma_1$, the coefficient for the buffer effect is around -0.007 and statistically significant at the 1% level. Between the first threshold $\gamma_1$ and the second threshold $\gamma_2$, 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 2. After the second threshold $\gamma_2$, the buffer effect is 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 $\gamma_1$). 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 $\gamma_1$ and inferior to $\gamma_2$). This last result provides further empirical support for our main conjectures, 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 $\gamma_2$). 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 *ex ante* international reserves, *res*, and effective terms of trade, *etot*, is negative and significant in almost all the regressions, including the robustness analyses. As mentioned before, this coefficient measures the buffer effect of international reserves on the real exchange

**Table 5**
Panel threshold regression and financial openness.

| *Variables* | (1) *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) |
| $etot \times L.res.I\left(L2.KAOPEN \leq \gamma_1\right)$ | -0.0046*** |
| | (0.0017) |
| $etot \times L.res.I\left(\gamma_1 < L2.KAOPEN \leq \gamma_2\right)$ | -0.0235*** |
| | (0.0024) |
| $etot \times L.res.I\left(L2.KAOPEN > \gamma_2\right)$ | -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.

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 2). 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 coefficient is equal to -0.111 and significant at the one percent level (see Table 3).

In the threshold regressions with the level of financial institution development in Fig. 3, 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 5, we obtain a value similar to the baseline for intermediate level of financial openness.

In the Appendix, the threshold regressions in Table B.2 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. We complement our previous empirical analyses with several models that consider endogeneity of the covariates in a dynamic panel threshold model (Table B.3) and the endogeneity of the threshold variable (Table B.4). The results are very similar to the threshold regressions in Table 4 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 Fig. B.2, 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 Fig. B.3, we construct variables for the variation of *res* 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.5, 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.6. The threshold effect of financial institutions is confirmed after the global financial crisis. In Table B.7, we analyze different country groupings. Unsurprisinlgy, the buffer effect is stronger for commodity exporters. In the online appendix, we consider the role of macroprudential policy.[17]

## 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

[17] The online appendix is available here: https://www.nber.org/papers/w30891.

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 that the buffer effect can be observed in different country groups.

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.

## CRediT authorship contribution statement

**Joshua Aizenman:** Conceptualization, Methodology, Project administration, Supervision, Writing – original draft, Validation, Writing – review & editing, Investigation, Formal analysis. **Sy-Hoa Ho:** Writing – review & editing, Investigation, Visualization. **Luu Duc Toan Huynh:** Visualization, Writing – review & editing, Investigation. **Jamel Saadaoui:** Data curation, Formal analysis, Investigation, Methodology, Writing – original draft, Writing – review & editing, Conceptualization, Project administration, Visualization. **Gazi Salah Uddin:** Data curation, Methodology, Visualization, Writing – review & editing, Writing – original draft.

## Appendix A. Data and country list

### A.1. Source and definition

**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 | World Bank, WDI |
| *tot* | Term of Trade | World Bank, WDI |
| *etot* | Effective terms of trade | World Bank, WDI |
| *res* | International Reserves to GDP | 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) |

**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.

## Appendix B. Additional robustness checks

### B.1. Threshold of international reserves

An important condition in the panel threshold model of Hansen (1999) is that the threshold must be exogenous for valid inference.[18] We test the persistence of international reserves to ensure that our threshold variable is sufficiently exogenous. We run an AR(1) panel regression with country-fixed effects. As we can see in Table B.1, international reserves are persistent and share a different frequency fluctuation with the terms of trade. The picture is similar when we look at the individual coefficients for each

[18] In Appendix B.2, we use two panel threshold estimators that relax this hypothesis.

**Table B.1**
Panel AR(1) regression for the international reserves.

| *Variables* | (1) *res* |
|---|---|
| *L.res* | 0.787*** |
| | (0.0138) |
| *Constant* | 0.550*** |
| | (0.0348) |
| | |
| Observations | 1,900 |
| Number of countries | 100 |
| R-squared | 0.644 |

Note: Standard errors are in parentheses. Fixed effects are included but not shown. *** indicate statistical significance at the 1% level. $L$ stands for the lag operator.
Source: Authors' estimates.

country. The AR(1) coefficient is greater than 0.6 for more than 80% of countries.[19] This could be explained by "fear of losing reserves" as explained by Aizenman and Hutchison (2012). To check whether international reserves react to terms of trade, we ran a panel regression with country-fixed effects between these two variables. The p-value of the coefficient for the terms-of-trade variable is above 20%, showing that international reserves are not very responsive to terms-of-trade. Together, these results may indicate that the threshold variable is exogenous. Thus, we test the following Equation:

$$rer_{i,t} = \mu_i + \delta_1 etot_{i,t} I\left(res_{i,t-1} \le \gamma\right) + \delta_2 etot_{i,t} I\left(res_{i,t-1} > \gamma\right) + \alpha' \mathbf{x}_{i,t} + u_{i,t}, \tag{B.1}$$

In Table B.2, we move to the threshold regressions for the whole panel and for some of the regional panels. Indeed, holding foreign reserves may mitigate the magnitude of exchange rate adjustments triggered by capital flow movements,[20] 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. In this case, the coefficient for the buffer effect is negative after the threshold. However, we could also expect that some regions of the world economy, especially those with large amounts of commodities, accumulate too many international reserves. In this case, we could observe the buffer effect up to a certain level of reserves. In fact, we do not observe any buffer effect after this threshold. Here, the coefficient for the buffer effect is negative before the threshold and nonsignificant after the threshold.

In Table B.2, 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. The threshold value corresponds to 17.28% for international reserves. These results mean that for observations (countries and periods) above 17.28% for international reserves,[21] 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 Fig. B.1 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 B.2, 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.

### *B.2. 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 (2) 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, $\widehat{rer}_{i,t-1}$.

[19] The intervention channel may also matter in the short run. In this case, identification would be more difficult as the level of reserves may influence currency interventions and the speed of exchange rate variation. However, since we used a sample with a time dimension of 20 years, we expect that the main channel is the management of *ex ante* reserves, and, hence, the balance sheet channel. These pieces of evidence about the "fear of losing reserves" support the balance sheet channel as the main causal channel in our study.

[20] Devereux and Wu (2022) and Ahmed et al. (2023) find that holdings of foreign reserves are associated with an exchange rate that is less sensitive to global shocks.

[21] 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.

**Table B.2**
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.

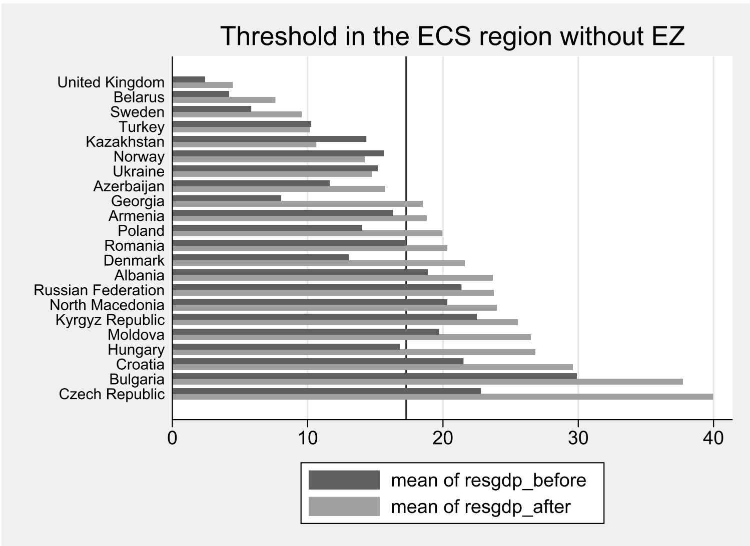


**Fig. B.1.** Threshold effect in the ECS region. *Notes:* We use a selection of emerging and developing ECS countries to compare the value of the threshold (17.28% of GDP) found in this region with the evolution of the holding of international reserves (mean value) before and after the GFC. Source: Authors' estimates.

2. In the second step, the dynamic version of Equation (B.1) is estimated using least squares for a fixed threshold $\gamma$ 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, $\hat{\gamma} = \underset{\gamma}{\operatorname{argmin}}\, S_n(\gamma)$. According to Hansen (1999) and Caner and Hansen (2004), the critical values for determining the 95% confidence interval of the threshold value are 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 $\hat{\gamma}$.

**Table B.3**
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%, respectively. The non-significant time dummies have 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 provides 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.

In Table B.3, the existence of the buffer effect below a threshold of financial institution development around 0.46 confirms the results obtained in Table 4. We also consider the approach of Seo and Shin (2016); Seo et al. (2019) as the threshold variables can be endogenous. In Table B.4, the results are qualitatively similar.

### *B.3. Local projections*

In this Appendix, we will use the local projections approach (hereafter LP) (Jorda, 2005) to provide further empirical evidence on the buffer effect of international reserves. Thanks to the Stata package written by Ugarte (2023), we use panel LP to complement our baseline results in Table 2. The LP approach presents 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 nonlinear and flexible models is straightforward, and the approach can be easily scaled to panel data. Regarding 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:

$$y_{i,t+h} = b_h S_{i,t} + \gamma_h y_{i,t-1} + \alpha' \mathbf{z}_{i,t-1} + v_{i,t+h}$$
$$\mathrm{IRF}(h) = \hat{b}_h$$

with $y$, is the explained variable, $h$, the horizon, $S$, the impulse variable, $\mathbf{z}$ is a vector of control variables, IRF, stands for the 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 2, including the *de facto* exchange rate regime.

As a first step, we need to check whether the buffer effect comes only from the interaction term or from each variable *res* and *etot*. As we can see in Fig. B.2, 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.

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. 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 not correlated with the real exchange rate by construction. We will use the interaction of these residual variables as the impulse variable (i.e., the shock). Thus, each shock in

**Table B.4**
Dynamic threshold panel model (Seo and Shin, 2016).

| Variables | (1)<br>$rer_{i,t}$ |
|---|---|
| | *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 are in parentheses. The symbols ***, ** correspond to statistical significance at 1 and 5%, respectively. All differences are first differences. We use three-year averages for the variables to smooth out the fluctuations of the business cycle. Instruments are the lagged (three-year) average 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.

the interaction term will be uncorrelated with real exchange movements. This will provide robust evidence for one possible source of endogeneity, namely the reverse causality that arises from the exchange rate on international reserves and effective terms of trade. As you can see in Fig. B.3, the IRF results for the buffer effect are very close to the baseline results of Table 2 at the horizon $h = 0$ (-0.020), providing further evidence of 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.

### B.4. Cross-sectional correlation

In this Appendix, we follow Sul (2019) to consider the influence of cross-sectional correlation on our estimates. In Table B.5, the introduction of year-fixed effects or the cross-sectional mean of the real exchange rate is sufficient to eliminate 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 common factors. In the 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.

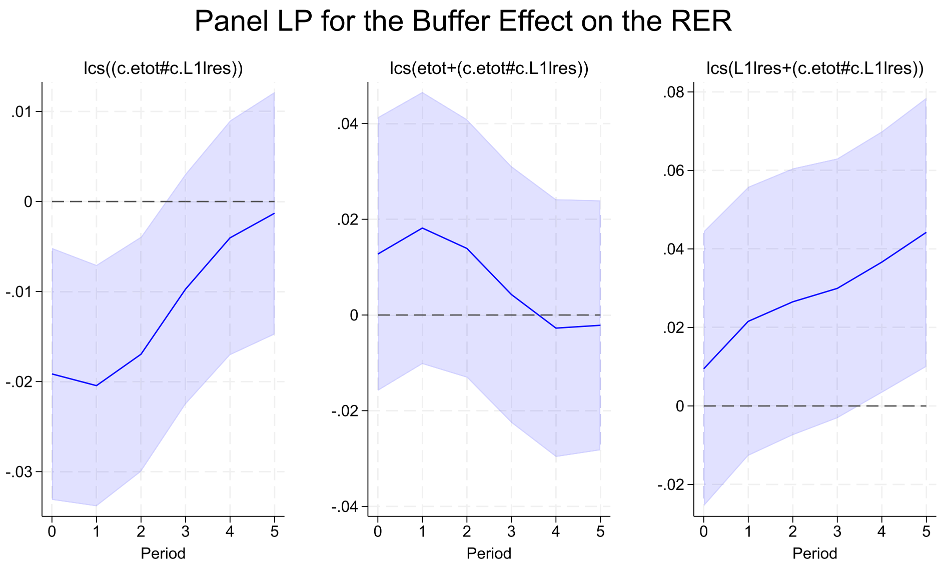


**Fig. B.2.** Panel LP for the buffer effect on the RER. *Notes:* In the left panel, the unit shock is only on the interaction. In the center panel, the unit shock is on the interaction term and the effective terms of trade variable, simultaneously. In the right panel, the unit shock is on the interaction term and the international reserves variable, simultaneously. Robust standard errors. 95% confidence intervals in light blue. Source: Authors' estimates.

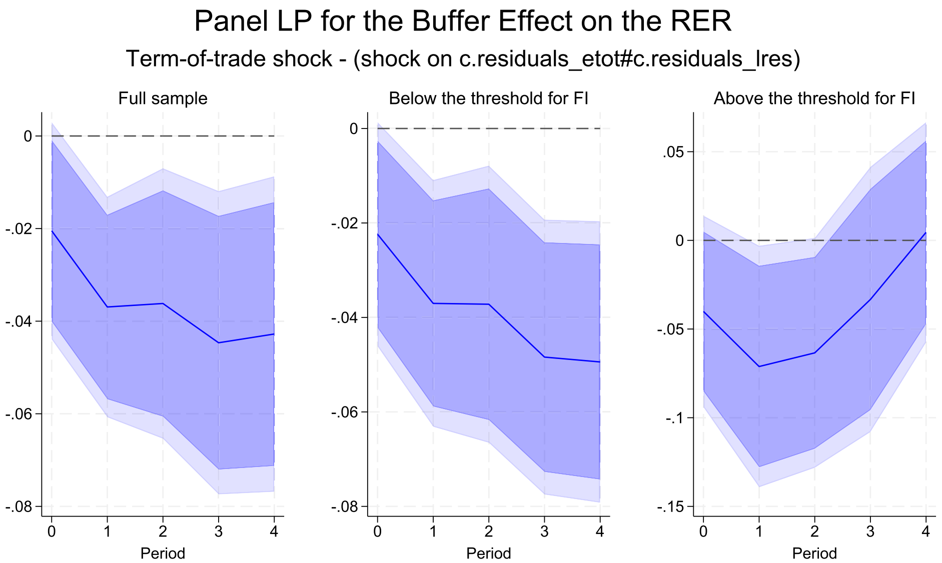


**Fig. B.3.** Panel LP for the buffer effect on the RER. *Notes:* On the left panel, the unit shock is on the full sample. In the center panel, we use the data below the previously identified threshold for financial institution development. In the right panel, we use the data above the previously identified threshold for financial institution development. Bootstrapped standard errors. 90%, 95% confidence intervals in dark blue and light blue, respectively. Source: Authors' estimates.

### *B.5. 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 Table B.6, we find that the buffer effect is stronger after the GFC. The threshold effect for the financial institution indicator is also confirmed after 2008. Our results hold, before and after the GFC, and indicate that the accumulation of *ex ante* reserves and, therefore, the balance sheet channel is the most important causal channel in our sample. These robustness checks allow us to control for the GFC period, where almost all emerging countries performed currency interventions, as documented by Dominguez et al. (2012).

**Table B.5**
Factor augmented panel regressions.

| | (1) | (2) | (3) |
|---|---|---|---|
| *Variable* | *mean rer*<br>*rer* | *year effects*<br>*rer* | *iterative interactive*<br>*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* × *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 used 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.

**Table B.6**
Before and after the Global Financial Crisis.

| *Variables* | (1)<br>FI – after the GFC<br>*rer* | (2)<br>before the GFC<br>*rer* | (3)<br>after the GFC<br>*rer* |
|---|---|---|---|
| [6pt] *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* × *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.

### *B.6. 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.7, we note that OECD countries[22] 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 the 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)

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

**Table B.7**
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.

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 achieved through the management of sovereign wealth funds. Finally, we find that 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.

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