Inelastic Financial Markets and Foreign Exchange Interventions*

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Abstract

We provide empirical support for models of inelastic international financial markets by leveraging the rebalancings of a local-currency government bond index as a well-identified currency demand shock. Our results are consistent with model predictions of inelastic financial markets where foreign exchange interventions emerge as an effective policy tool to manage exchange rates, without compromising monetary policy independence even in the presence of free capital mobility, relaxing the classical trilemma constraint. We show empirically that countries with a free-floating exchange rate regime are most effective at stabilizing exchange rates, consistent with the theoretical prediction that their volatile exchange rates generate further departure from the trilemma constraint.

JEL Codes: F31, F32, G11, G15, G23
Keywords: Exchange rates; inelastic financial markets; uncovered interest parity; foreign exchange interventions; benchmark investments; sovereign bonds.

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1 Introduction

Recent theories of exchange rate determination emphasize inelastic international financial markets, where exchange rates respond significantly to exogenous currency demand shocks (Gabaix and Maggiori, 2015; Itskhoki and Mukhin, 2021). Under inelastic financial markets, foreign exchange interventions act as financial flows similar to currency demand shocks, serving as an effective policy tool to manage exchange rates.[1] The key mechanism in these models stems from financial intermediaries' limited risk-bearing capacity when absorbing imbalances in financial flows, which leads to deviations from uncovered interest parity (UIP) and movements in exchange rates.

Quantifying the intermediaries' risk-bearing capacity in the foreign exchange markets is vital for policymakers. This capacity determines both exchange rate elasticity and the effectiveness of foreign exchange interventions. However, empirically validating these theoretical predictions remains difficult, requiring a well-identified demand shock that moves exchange rates independent of the macroeconomic fundamentals.

In this paper, we overcome this identification challenge by exploiting an exogenous currency demand shock from the mechanical rebalancings of J.P. Morgan's Government Bond Index-Emerging Markets (GBI-EM) Global Diversified. Our empirical results provide evidence supporting models of inelastic financial markets, where foreign exchange interventions effectively stabilize exchange rates while maintaining monetary policy independence. Our findings indicate substantial heterogeneity in the effectiveness of foreign exchange interventions across exchange rate regimes and volatility, consistent with the theoretical predictions on the intermediaries' risk-bearing capacity.

Our identification strategy relies on the exogenous currency demand shock created by the mechanical rebalancings of the GBI-EM Global Diversified index. This is the most widely tracked benchmark index by mutual funds that invest in local-currency

[1] The predictions from the inelastic financial markets model contradict that from the standard textbook models of international finance (Mundell, 1962; Gali and Monacelli, 2005), which assume financial markets are perfectly elastic and the uncovered interest parity (UIP) condition holds. Under the assumption of elastic markets, foreign exchange interventions, defined as central banks buying or selling international reserves on a large scale, would be a mere "re-shuffling" of assets between households and the central bank and thus have no traction on exchange rates.

government bonds in emerging markets. The monthly rebalancings cap the benchmark weight of each country in the index at 10%, with excess weight redistributed to smaller countries. This mechanical rule generates exogenous currency demand shocks: smaller countries *not* at the cap receive weights change not due to their economic fundamentals, but solely because larger countries hit the cap. The resulting rebalancing flows from index-tracking mutual funds provide clean identification of currency demand shocks that are "information-free" and independent of country fundamentals.

We demonstrate the exogeneity of our constructed currency demand shock through regression analysis of macroeconomic variables and policy instruments around rebalancing events. Despite the significant exchange rate response in both the short- and long-run that lasts for at least three months, macroeconomic variables (e.g., GDP, consumption, inflation, and net exports), central bank policy rates, and foreign exchange interventions data do not respond to the currency demand shock. This means that macroeconomic conditions are smooth around the index rebalancing events, consistent with the exogeneity assumptions.

The fact that exchange rates respond significantly to the currency demand shock is consistent with predictions from models of inelastic international financial markets. These models predict that financial intermediaries have limited risk-bearing capacity, making exchange rates responsive to currency demand shocks and foreign exchange interventions. As a result, interventions can effectively stabilize exchange rates while preserving monetary policy independence, even with free capital mobility. This ability to simultaneously achieve exchange rate stability and monetary autonomy creates a departure from the classical trilemma of Mundell (1962), which we refer to as the "relaxed trilemma." (Basu, Boz, Gopinath, Roch, and Unsal, 2023; Itskhoki and Mukhin, 2023a; Masao, Nakamura, and Steinsson, 2024)

Our empirical results suggest that the more inelastic the financial markets, the more effective the foreign exchange interventions. This pattern aligns well with theoretical predictions of Gabaix and Maggiori (2015) and Itskhoki and Mukhin (2021): the higher exchange rate volatility in free-floating regimes ("free floaters") reduces intermediaries' risk-bearing capacity, making financial markets more inelastic and creating a larger departure from the trilemma constraint. In contrast, pegged regimes ("peggers") operate under the classical trilemma (Mundell, 1962; Gali and Monacelli, 2005) where financial markets are perfectly elastic, causing interventions to have no traction on exchange rates.

We also find empirical evidence supporting the "relaxed trilemma." We show that regardless of the exchange rate regimes, capital control measures and monetary policy rates are uncorrelated with our exogenous currency demand shock. This is contradictory to the trilemma, under which the movements in exchange rates would have to be offset one on one by monetary policy rates for exchange rates to be fixed, under any given capital control taxes. Our findings are consistent with the model implications of inelastic financial markets where foreign exchange interventions work as an effective policy tool to stabilize exchange rates.

Our results suggest that the required size of foreign exchange intervention is less than half as large for free floaters as for managed floaters or peggers, indicating that interventions are more effective for the free floaters. This greater effectiveness is evidenced by the larger exchange rate response to currency demand shocks among free floaters. By computing the assets under management of mutual funds tracking the index, we convert this exchange rate response into implied flows from the rebalancings. We find that achieving a 1 percent exchange rate appreciation (or depreciation) requires the central bank to sell (or buy) foreign reserves equivalent to approximately 0.04 % of GDP through foreign exchange interventions for free floaters, while the numbers for managed floaters and peggers are 0.07 % and 0.13%, respectively.

Related Literature. Our results contribute to several strands of literature in international finance and macroeconomics and provide insights for central bank policymakers regarding foreign exchange interventions.

First, our paper advances the literature on demand estimation for exchange rates, building on recent papers including Menkhoff, Sarno, Schmeling, and Schrimpf (2016, 2017), Koijen, Moskowitz, Pedersen, and Vrugt (2018) Koijen and Yogo (2020), Jiang, Richmond, and Zhang (2022, 2024). We add to this literature by finding a plausible exogenous currency demand shock through leveraging the mechanical rebalancings of a local-currency government bond index as a quasi-natural experiment.

Compared with other work on exchange rates estimation using index rebalancings, we adopt a distinct approach from Hau, Massa, and Peress (2009) and Broner, Martin, Pandolfi, and Williams (2021) by analyzing repeated rebalancing events over a long time series, providing a powerful framework for analyzing persistent foreign exchange interventions.[2] Compared with Pandolfi and Williams (2019)'s initial study using the GBI-EM Global Diversified index, we reveal fundamentally new insights on how rebalancing shocks heterogeneously impact currencies across exchange rate regimes and volatilities, estimate medium-run elasticity over much longer horizons relevant for policymakers evaluating sustained intervention episodes, and interpret our empirical findings through an inelastic financial markets framework that supports model predictions about country-specific risk-bearing capacities.

Secondly, our paper provides empirical support for models of exchange rate determination in inelastic financial markets and contribute to the intermediary-based asset pricing literature. Recent theories such as Gabaix and Maggiori (2015) and Itskhoki and Mukhin (2021) predict that exchange rates respond to currency demand shocks due to limited risk-bearing capacity of intermediaries.[3] Our findings not only validate this core prediction, but also quantify currency-specific risk-bearing capacity that varies with exchange rate regimes and volatility. This adds to the intermediary-based asset pricing literature by documenting novel cross-sectional heterogeneity in currency risks that affects foreign exchange intervention effectiveness.[4]

[2]While several recent studies like Hau, Massa, and Peress (2009) and Broner, Martin, Pandolfi, and Williams (2021) have examined exchange rate responses to index changes, they focus on one-time events like announcements for index redefinitions or inclusions. In addition, while Camanho, Hau, and Rey (2021) estimate currency elasticity using fund-level equity portfolio flows using granular instrumental variables, our work achieves particularly sharp identification through a quasi-natural experiment of mechanical index rebalancings. Other related work using index rebalancings to estimate the demand curves for international assets include Raddatz, Schmukler and Williams (2017), Williams (2018), and Moretti, Pandolfi, Schmukler, Villegas-Bauer and Wiiliams (2024).

[3]The related papers, building on the "portfolio rebalancing model" by Kouri (1976), include Jeanne and Rose (2002), Gabaix and Maggiori (2015), Gourinchas, Ray, and Vayanos (2019), Greenwood, Hanson, Stein, and Sunderam (2020), Itskhoki and Mukhin (2021), Jiang, Krishnamurthy, and Lustig (2022), Kekre and Lenel (2024), and Bacchetta, Davis, and van Wincoop (2024).

[4]Our paper contributes to the intermediary-based asset pricing literature in foreign exchange markets (Du, Hébert, and Huber, 2023; An and Huber, 2024; Liao and Zhang, 2025) that focuses on time-varying risk-bearing capacity, while we document novel cross-sectional heterogeneity arising from exchange rate

Thirdly, our work speaks to the large literature in international asset pricing on exchange rates prediction and deviations from uncovered interest parity condition. Several papers including Hassan and Mano (2019) and Hassan, Mertens, and Wang (2024) have documented persistent differences in currency risks across countries.[5] We contribute to this strand of literature by relating currency risks with exchange rate regimes and volatilities to inform the effectiveness of intervention policies. Another strand of literature leverages taste shocks or expectation errors in forecasting exchange rates.[6] Our empirical strategy relies on an exogenous quantity shock from the mechanical index rebalancings for predicting exchange rates.

Lastly, we contribute to the empirical literature on foreign exchange interventions (Fatum and Hutchison, 2003; Blanchard, Adler, and de Carvalho Filho, 2015; Fratzscher, Menkhoff, Sarno, and Stohr, 2019; Adler, Lisack, and Mano, 2019) by providing clean identification of intervention effectiveness through quasi-natural experiments based on mechanical index rebalancings. Our results also provide empirical support for the "relaxed trilemma" in the theoretical framework on foreign exchange interventions by Basu, Boz, Gopinath, Roch, and Unsal (2023) and Itskhoki and Mukhin (2023a).[7]

2 Introducing the Currency Demand Shock

We leverage the mechanical rebalancing features of a local-currency government bond index for emerging countries to construct an exogenous currency demand shock. We document in detail the rebalancing rules of the index and introduce our measure for the

regimes and volatility.

[5]Several papers have provided empirical evidence on currency return differentials across countries relating to various economic features, including Lustig and Verdelhan (2007), Lustig, Roussanov, and Verdelhan (2011, 2014), Martin (2011), Hassan (2013), Richmond (2019), and Korsaye, Trojani and Vedolin (2023). Several other important papers have attempted to microfound the time variation in currency risks, including Verdelhan, (2010), Lewis (2011), Colacito and Croce (2011, 2013), Stathopoulos (2017), and Colacito et al. (2018).

[6]The related papers include Gourinchas and Tornell (2004), Lustig and Verdelhan (2007), Kremens and Martin (2019), Bacchetta and van Wincoop (2021), Jiang, Krishnamurthy and Lustig (2022), Kremens, Martin and Liliana (2023), Engel and Wu (2024), and Chahrour, Cormun, De Leo, Guerrón-Quintana, and Valchev (2024).

[7]Other related theoretical work on the foreign exchange policy framework in Jeanne (2012), Amador, Bianchi, Bocola, and Perri (2019), Cavallino (2019), Fanelli and Straub (2021), and Bacchetta, Benhima, and Berthold (2023).

currency demand shock as well as the implied capital flows from the shock.

2.1 Mechanical Rebalancings of the GBI-EM Global Diversified

Our empirical strategy relies on the mechanical rebalancings of the Government Bond Index-Emerging Markets (GBI-EM) Global Diversified published by J.P. Morgan. This is the largest local-currency government bond index for emerging countries in terms of its assets under management. At the time of writing, there are 19 emerging countries in the index; each country's weight equals the share of its market value of the local-currency sovereign bonds in the index. A larger country, such as Brazil, has a larger weight in the index than a smaller country, such as Hungary or Philippines.

The mechanical rebalancings by the GBI-EM Global Diversified index on the country weight cap are crucial for the identification in this paper. The country weight fluctuates daily as the market price of the sovereign bonds moves up or down. However, at the rebalancing date (which is the end of the last business day of each month), the index mechanically caps the country weight at 10% for all countries to limit concentration risk. Any excess weight above the 10% cap is redistributed to smaller countries that are below the cap, proportionally to their allocation so that all country weights add up to 100%. In addition, countries at the cap will follow the "10/10 rule": the country that meets the 10% cap or larger in the GBI-EM index will be staggered over for a 10-month period. This means that once Brazil's allocation in the index exceeds 10%, the rule fixes its allocation at 10% for each rebalancing date over the subsequent 10-month period.[8]

We argue that for countries *not* at the 10% country-weight cap, their change in weights in the GBI-EM Global Diversified index creates currency demand shocks that are uninformative to the macroeconomic fundamentals. For example, if Brazil's country weight is rebalanced from 12% to 10% and leads to an increase in Hungary's country weight, those benchmarked mutual funds have to sell local-currency sovereign bonds of Brazil and buy Hungarian Forint in order to purchase local-currency sovereign bonds of Hungary. In this rebalancing example, a smaller country experienced a positive currency

[8] It is for this reason that we see that the monthly after-rebalancing weights of Brazil are typically staggered at the 10% cap at the end of month, as reported in Table B.6 in the Appendix.

demand shock on its local-currency bonds independent of its own macroeconomic conditions and purely as a result of a larger country hitting the 10% cap. The rebalancings can continue recursively for multiple rounds until all the country weights are either at or below the 10% cap. Appendix A.1 gives more details on the rebalancing rules of the GBI-EM Global Diversified index.

2.2 Measuring the Currency Demand Shock

We introduce $\mu_{c,t}$ to capture the currency demand shock from the mechanical rebalancings of the GBI-EM Global Diversified index, for country c at the rebalancing date t. As shown in equation (1), we define $\omega_{c,t}^{\text{before}}$ and $\omega_{c,t}^{\text{after}}$ as the country weight before and after the rebalancing event, respectively, at the rebalancing date. Taking market price $P_{c,t}$ as given, J.P. Morgan adjusts the country weights (from $\omega_{c,t}^{\text{before}}$ to $\omega_{c,t}^{\text{after}}$) through changing the par value ($\hat{Q}_{c,t}$) of the local-currency sovereign bonds of the countries included in the index

$$\mu_{c,t} = \frac{\omega_{c,t}^{\text{after}} - \omega_{c,t}^{\text{before}}}{\omega_{c,t}^{\text{after}}}, \tag{1}$$

where $\omega_{c,t}^{\text{before}} = \frac{P_{c,t}\hat{Q}_{c,t-1}}{\sum_c P_{c,t-1}\hat{Q}_{c,t-1}} \times \frac{1}{(1+r_t)}$ and $\omega_{c,t}^{\text{after}} = \frac{P_{c,t}\hat{Q}_{c,t}}{\sum_c P_{c,t}\hat{Q}_{c,t}}$; $P_{c,t}$ is the aggregate market price of the local-currency sovereign bonds for country c at the rebalancing date; $\hat{Q}_{c,t-1}$ and $\hat{Q}_{c,t}$ are the aggregate par value of the local-currency sovereign bonds included in the index from the last rebalancing and the current rebalancing, respectively;[9] r_t is the monthly return of the GBI-EM Global Diversified index from the past rebalancing date $t-1$ to the current rebalancing t. The weight $\omega_{c,t}^{\text{before}}$ is also the buy-and-hold weight of the country and the weight of passive funds tracking the index.[10]

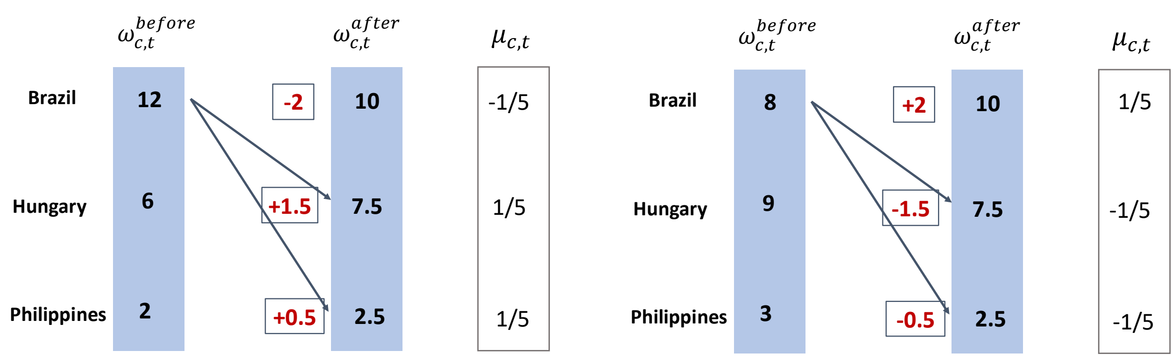
[9] It is important to distinguish the face amount of sovereign bonds included in the index ($\hat{Q}_{c,t}$) from the face amount of the actual issuance ($Q_{c,t}$) by the sovereign. Appendix A.1 explains the linear extrapolation rule where a portion of the country's actual sovereign bonds outstanding is included in the GBI-EM Global Diversified index.

[10] This is because $\sum_c P_{c,t-1}\hat{Q}_{c,t-1} \times (1+r_t)$ is the is the market value of the index at the rebalancing date t if rebalancing does not take place. Throughout our sample, we can safely assume that $\sum_c P_{c,t-1}\hat{Q}_{c,t-1} \times (1+r_t) \approx \sum_c P_{c,t}\hat{Q}_{c,t}$. The difference between the two can occur due to deletion of matured bonds or addition of new bonds to the index (another layer of rebalancing addressed in Appendix A). The average difference between $\sum_c P_{c,t-1}\hat{Q}_{c,t-1} \times (1+r_t)$ and $\sum_c P_{c,t}\hat{Q}_{c,t}$ in our sample is about 0.5%.

Intuitively, the currency demand shock, $\mu_{c,t}$, is "information-free" of market and macroeconomic conditions. As both $\omega_{c,t}^{\text{before}}$ and $\omega_{c,t}^{\text{after}}$ use the market price of the sovereign bonds at the rebalancing date t, their difference captures the change in country weight induced purely by the mechanical rebalancing of the index. We normalize $\mu_{c,t}$ by its own weight after rebalancing to account for differences in sovereign bond market sizes across countries. Table 2.1 illustrates this with two simplified rebalancing examples: one where smaller countries receive a positive weight increase (and thus a positive demand shock), and another where they receive a negative demand shock.

Table 2.1: Two Simplified Rebalancing Examples at the 10% Weight Cap

(a) Example I (b) Example II



Note: This table presents two simplified rebalancing examples that cap the country weight at 10%. For tractability, we consider a scenario with 11 countries where 8 countries are fixed at the 10% cap, focusing our analysis on the remaining three countries' weight adjustments.

Example I illustrates a standard cap-driven rebalancing. Brazil enters with a 12% weight (exceeding the cap), while Hungary (6%) and Philippines (2%) begin below the cap. The rebalancing mechanism reduces Brazil's weight to 10% and redistributes the excess 2% proportionally among Hungary and Philippines according to their initial weights. This redistribution yields a currency demand shock ($\mu_{c,t}$) of 1/5 for both Hungary (calculated as 1.5/7.5) and Philippines (calculated as 0.5/2.5).

Example II demonstrates the "10/10 rule" mechanism. Although Brazil's initial weight (8%) falls below the cap, its previous breach of the 10% threshold triggers the rule, requiring its weight to increase to 10%. This forced increase leads to proportional weight reductions in Hungary (9%) and Philippines (3%), resulting in negative currency demand shocks ($\mu_{c,t}$) of -1/5 for both countries (calculated as -1.5/7.5 and -0.5/2.5, respectively).

In both examples, while we report Brazil's $\mu_{c,t}$, we exclude it from our main empirical analysis, focusing instead on the induced shocks to the smaller countries.

Our main empirical analysis focuses on currency demand shocks from countries that do *not* meet the 10% cap at the rebalancing dates. These countries have to change their weights as a result of the bigger countries meeting the cap (either by exceeding the cap or staggered due to the “10/10 rule"). Therefore, their weight changes are independent of their macro-fundamentals; their corresponding currency demand shocks are “information-free" by construction, as the market prices and the magnitudes of shock are only revealed at the rebalancing dates. In the example in Table 2.1, we would use only the change in weights from Hungary and Phillippines for our identification.

Remark 1. *How do publicly known rebalancing events generate “information-free” currency demand shocks?*

The mechanical rebalancing rules of the GBI-EM Global Diversified index are publicly known. For example, when a large country's weight at the beginning of the month exceeds 10%, investors anticipate to sell positions in that country's local-currency government bonds. However, the currency demand shock $\mu_{c,t}$ remains “information-free" because the precise magnitude of weight changes for smaller countries depends on market prices at rebalancing. This uncertainty in magnitude, despite predictable direction, ensures that the shock carries no information about market conditions or fundamentals.

Table 2.2: Distribution of the Currency Demand Shock ($\mu_{c,t}$ in %)

$\mu_{c,t}$, *including* observations at the 10% cap

Obs.	Mean	Std.	Min.	Max.	Median	90%	10%
2,197	-0.152	4.480	-15.627	21.192	-0.031	4.387	-5.371

$\mu_{c,t}$, *excluding* observations at the 10% cap

Obs.	Mean	Std.	Min.	Max.	Median	90%	10%
1,565	-0.162	4.844	-15.627	21.192	-0.156	4.686	-5.643

Note: This table reports the summary statistics on the currency demand shock ($\mu_{c,t}$), in percentage points, implied by the monthly rebalancings of the GBI-EM Global Diversified index. In the top panel we report the distribution statistics including those at the 10% cap, and in the bottom panel we drop the observations at the 10% cap. A negative $\mu_{c,t}$ (< 0) implies that the country is rebalanced downwards, and vice versa for a positive $\mu_{c,t}$. In the main empirical analysis, we drop the countries at the 10% cap.

Table 2.2 shows the summary statistics of the currency demand shocks. Table B.5 and B.6 in the appendix report the time series of the currency demand shock and weight after the rebalancing for each country. There is significant heterogeneity across countries in meeting the 10% weight cap: Specifically, while Brazil, Mexico, and Poland each had the majority of its time series with weights capped at 10%, Indonesia, Malaysia, Russia, South Africa, Thailand, and Turkey also occasionally met the 10% cap and were staggered for a 10-month period because of the "10/10 rule" explained above. Smaller countries (Argentina, Chile, Colombia, Czech Republic, Hungary, Peru, the Philippines, Romania, and Uruguay) never met the 10% cap throughout the sample.

2.3 Flows Implied by the Currency Demand Shock

We show that mutual funds in the Emerging Portfolio Fund Research (EPFR) dataset with large positions that track the GBI-EM Global Diversified index closely follow its rebalancing rules, given by their high R-squared values against index returns.

We identify passive index-tracking funds in the EPFR dataset through two complementary approaches. First, we select all emerging market bond funds whose reported benchmark indices are the GBI-EM Global Diversified index[11]. Second, we identify funds through regressing the monthly returns of each fund in the EPFR dataset on the returns of the index and selecting those funds whose performance R-squared[12] (or "passivism") is at least 0.9. Taking the union of funds identified through these two approaches, we result in a final dataset comprising 2113 unique funds with a median performance R-squared of 0.92 (Table B.4a). To verify that funds with larger positions demonstrate high "passivism," we construct the weighted average return (by assets under management) of these mutual funds and regress the weighted return on the index returns. This results in an even higher R-squared of 0.97 (Table B.4b).

To convert the currency demand shocks to USD flows, we estimate the total assets under management of mutual funds tracking the GBI-EM Global Diversified index globally. Our estimation process involves two steps. First, using EPFR data, we measure the

[11] Details on how we selected mutual funds into the data are reported in Appendix A.2.

[12] We follow Amihud and Goyenko (2013) and Pandolfi and Williams (2019) and use the return regression to test the performance of mutual funds. The method regresses the fund-level monthly returns on the monthly returns of the GBI-EM Global Diversified: $r_{i,t} = \alpha + \beta r_{B,t}$, where $r_{i,t}$ is the monthly returns from fund i at time t, and $r_{B,t}$ is the monthly returns from the benchmark (i.e.,the J.P. Morgan GBI-EM Global Diversified index). We then collect the fitted R-squared from each return regression. A higher fitted R-squared indicates the fund tracks the benchmark index more closely (or higher "passivism").

assets under management of funds tracking the index from 2016 to 2022 (Figure B.3, a). Second, we scale up this measure using data from the Investment Company Institute (ICI) Global, which shows that EPFR data covers approximately 60% of the global mutual fund population in 2019 (Figure B.3, b). This implies that total global index-tracking assets of GBI-EM Global Diversified index add up to 200 billion USD in 2019.[13]

Remark 2. *Why do we use the assets under management of passive funds that track the index?*

We focus on passive mutual funds when computing flows from GBI-EM Global Diversified index rebalancing because they best capture the information-free currency demand shock. Unlike active funds that can trade anytime, passive funds must hold their positions until month-end rebalancings. Thus, only passive fund positions are directly linked to our constructed currency demand shock, $\mu_{c,t}$, in equation (1).

2.4 Data Sources

The main data source we use is the Index Composition and Statistics reports from J.P. Morgan, which include monthly information on benchmark weights for the GBI-EM Global Diversified index. Our sample comprises a panel of 18 countries from 2010 to 2021: Argentina, Brazil, Chile, Colombia, Czech Republic (Czechia), Hungary, Indonesia, Malaysia, Mexico, Peru, the Philippines, Poland, Romania, Russia, South Africa, Thailand, Turkey, and Uruguay.[14]

The second main data source we use is the EPFR data. We show that the currency demand shock correspond to changes in asset positions of the passive mutual funds tracking the GBI-EM Global Diversified index in the EPFR data. Moreover, we use the EPFR data to compute the flows in US dollars implied by the rebalancings by our currency demand shock.

[13] In 2019, our sample of EPFR funds tracking the GBI-EM Global Diversified index managed 120 billion USD. Since EPFR covers 60% of global mutual funds, we estimate total global assets tracking the index at 200 billion USD for that year.

[14] We exclude China from the current analysis because there are limited time series on this country in the dataset, as China entered the GBI-EM Global Diversified index only in 2020; we exclude Dominican Republic and Nigeria from the analysis because there are limited data on exchange rates for these countries from the Bank for International Settlements (BIS) statistics.

Finally, we combine J.P. Morgan reports and EPFR fund flows data with daily data of exchange rates and data on central bank policy rates from the Bank for International Settlements. We complement these data with the deposits yields data from Tullett Prebon Information (accessed via Haver)[15], and sovereign bond yields for each country from Du and Schreger (2016), with the dataset updated until 2021.

3 Currency Demand Shocks and Exchange Rates

We present four novel empirical facts on how the currency demand shock affects exchange rates and interest rates. We start by showing exchange rates respond to currency demand shocks in the pooled sample of countries following one rebalancing event (Fact 1), before presenting the core empirical results of the paper that the country-specific exchange rate response (Fact 2) correlates with exchange rate regimes and volatility. We also show that monetary policy rates do not contaminate our identification (Fact 3) and that the responses of exchange rates to the currency demand shocks are long-lasting (Fact 4). The section ends with further discussion on identification concerns.

Empirical Fact 1. *The currency demand shock moves exchange rates in the short run. A one standard deviation increase in the shock appreciates exchange rates by an average of 1.2% for the cross-country sample.*

Figure 1 reports the estimated coefficients of cumulative exchange rate changes on our currency demand shock as measured by $\mu_{c,t}$ in equation (1). The regression takes the following form

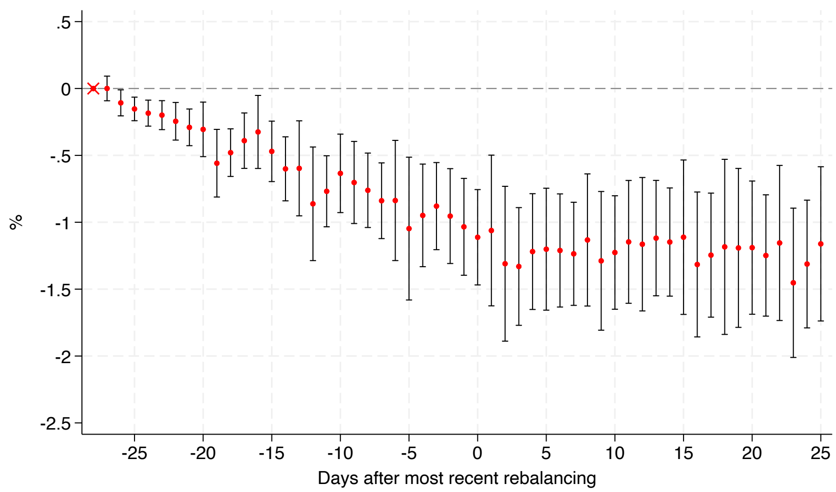
$$\Delta e_{c,t+d} = \beta_0 + \beta_\mu \ \mu_{c,t} + \phi \ X_{c,t} + \epsilon_{c,t}, \tag{2}$$

where $\mu_{c,t}$ is the currency demand shock defined in equation (1); β_0 is the constant; and $X_{c,t}$ is a set of dummies that control for country and date fixed effects, respectively. Standard errors are clustered at the date level. We include time fixed effects at the

[15]Tullett Prebon Information offers proprietary data on independent real-time price information from global OTC financial and commodity markets. Data is available for all countries in the sample, except for Peru and Uruguay. For these two countries, interest rates are sourced from their respective Central Bank interbank rates for the same maturity.

date level to account for the cyclicality of the global financial cycle (Rey, 2013) and the documented tighter balance sheet constraints for banks toward the quarter ends (Du, Tepper, and Verdelhan, 2018).

Figure 1:
Fact 1: Currency Demand Shock Moves Exchange Rates in the Short Run



Note: This figure presents the estimated regression coefficient of the change in exchange rates on the currency demand shock measured by $\mu_{c,t}$ in equation (1); $\mu_{c,t}$ is standardized by its mean and standard deviation in the regression. The change in exchange rates (local currencies per US dollar) is measured as the cumulative change starting from 28 days before the recent rebalancing at day 0. The regression is performed in a pooled OLS using time and country fixed effects, with standard errors clustered at the date level. The results are shown as point estimates (red dots) with 95% confidence intervals (black bars) for each regression.

Exchange rates are measured in local currencies per US dollar, and the exchange rate change $\Delta e_{c,t+d}$ is the cumulative change from the time interval from 28 days before the rebalancing date 0 until d days after ($d < 0$ for days before the rebalancing date 0; if $d > 0$, vice versa). Regression estimates for each d days after rebalancing are represented by the red dots in Figure 1, with the 95% confidence intervals represented as black bars. We standardize the currency demand shock by its mean and standard deviation according to the distribution in Table 2.2. As discussed, in our main empirical analysis we exclude all country-month observations that exceed the 10% weight cap from the regression to ensure the currency demand shock is information-free and independent of the macro-fundamentals.[16]

[16]Nevertheless, we report the regression results that include countries at the 10% weight cap in Figure B.1 in the appendix, which shows that the estimates are largely identical to those in Fact 1.

The pooled OLS regression shows that a one standard deviation increase in $\mu_{c,t}$ appreciates local-currency exchange rates significantly, by 1.2%, after one rebalancing event for the cross-country sample. The estimate of 1.2% is the average regression coefficient 0 to 5 days after a rebalancing event, when the exchange rate level does not appreciate further. A one standard deviation increase in $\mu_{c,t}$ is equivalent to a 4.844% (by Table 2.2) change in the country weight in the index, which is on average about 0.62 billion USD flows for the emerging countries in our sample.[17]

Remark 3. *Do the pre-rebalancing exchange rate adjustments invalidate the identification?*

The identification strategy remains valid because $\mu_{c,t}$ captures an exogenous shock unrelated to market conditions or macroeconomic fundamentals. While active funds may trade ahead of the rebalancing dates and lead to pre-rebalancing exchange rate adjustments, our estimated coefficient β_μ in equation (2) isolates the impact of mechanical flows from passive funds at rebalancing dates on exchange rates. Following the completion of rebalancing, we observe no additional exchange rate effects as there is no further flow pressure on exchange rates from this specific rebalancing event.

Remark 4. *Why do we use 28 days before rebalancing when defining $\Delta e_{c,t+d}$?*

We limit our analysis to 28 days before the rebalancing date because index rebalancing occurs monthly on the last business day. Extending the window beyond 28 days before rebalancing would introduce confounding effects from the previous month's rebalancing. The current window enables us to capture the gradual price adjustments that occur prior to the rebalancing date, driven by anticipatory trading from active funds.

Empirical Fact 2. *The country-specific exchange rate response to the currency demand shocks differs by exchange rate regime, with free floaters being much more responsive than peggers.*

[17] Details of the computation are in section 5. Note that our estimates appear more inelastic than Pandolfi and Williams (2019), who find that a one standard deviation of flows implied by rebalancings (FIR) – 0.875 percentage points or roughly 0.5 billion in market value – move exchange rates by 32 basis points. The difference arises from our estimation windows. While they focus on the 5-7 days before and after rebalancing dates, we estimate medium-run elasticity over longer horizons that is more suitable to draw policy implications on foreign exchange interventions.

We find heterogenous responses of exchange rates to the currency demand shock across countries. We repeat the exercise in Fact for each country and collect the estimated coefficients for the regression at rebalancing dates ($d = 0$), in the empirical specification

$$\Delta e_{c,t+d} = \beta_{0,c} + \beta_{\mu,c}\ \mu_{c,t} + \phi_c\ X_{c,t} + \epsilon_{c,t}, \tag{3}$$

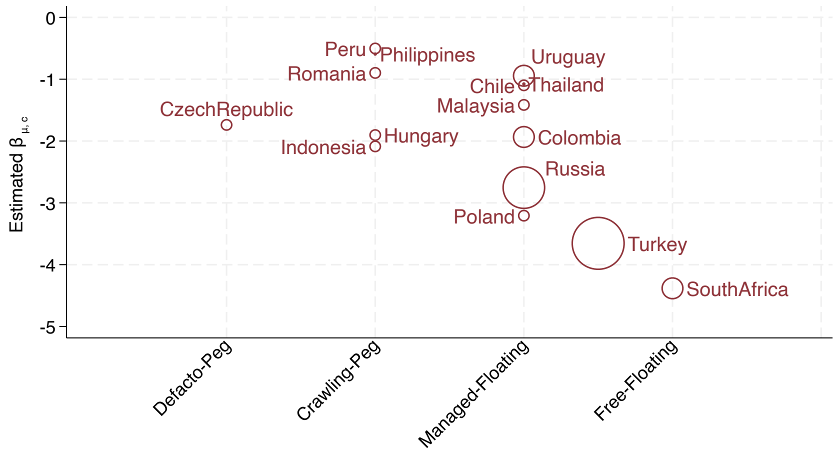
where we now estimate the country-specific exchange rate response $\beta_{\mu,c}$. All regressions include constants and a set of dummies $X_{c,t}$ that contain month and year fixed effects,[18] and standard errors are clustered at the year level. All countries respond to $\mu_{c,t}$ with less than 1% significance and predict the right sign. Specifically, a positive local-currency demand shock (an increase in $\mu_{c,t}$) appreciates local-currency exchange rates and decreases the price of US dollars in units of local currency. Table B.7 in the appendix gives the country-specific exchange rate response.

There is a clear relation between the country-specific exchange rate response and the exchange rate regimes, as illustrated by the downward trend in Figure 2. The y-axis is the estimated country-specific response of exchange rate change to the currency demand shock ($\mu_{c,t}$); the x-axis is the coarse exchange rate regimes ranging from de facto peg to free floating as classified by Ilzetzki, Reinhart and Rogoff (2021). The figure makes clear that free floaters (e.g., South Africa and Turkey) are much more responsive to $\mu_{c,t}$ compared with either managed floaters (e.g., Colombia, Malaysia, Poland, and Thailand) or peggers (e.g., Czech Republic, Romania, and Peru).

In addition, the country-specific exchange rate response correlates strongly with exchange rate volatility. Figure 2A plots the exchange rate volatility against exchange rate regimes (panel a) and against country-specific estimates (panel b). There is a clear correlation between the exchange rate volatility and the country-specific response. By contrast, we find no correlation between country-specific estimates and macroeconomic metrics such as GDP, M_2, consumption, and inflation (Table B.12 in appendix). Taken together with Figure 2 and 2A, country-specific exchange rate response increases with

[18] When month fixed effects are included, we need to drop both Brazil and Mexico from the regression because they have limited observations. Thus, we include only year fixed effects for these countries, and we report their estimates in Table B.7 in the appendix.

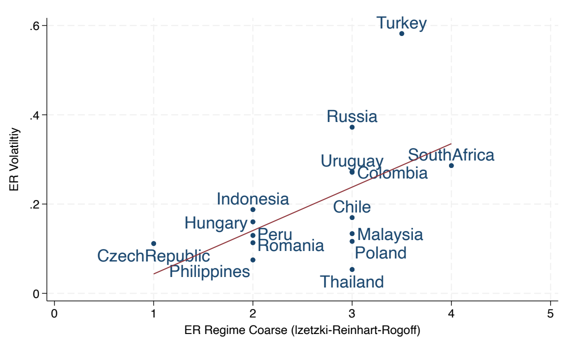
Figure 2:
Fact 2: Free Floaters Respond More to the Currency Demand Shock than Peggers



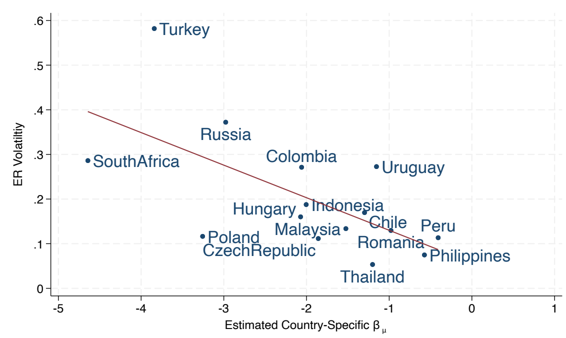
Note: This figure presents the relation between country-specific exchange rate response to the currency demand shock (measured by $\mu_{c,t}$) and the exchange rate regimes classified by Ilzetzki, Reinhart and Rogoff (2021). The y-axis is the estimated response of cumulative exchange rate change to $\mu_{c,t}$ at the rebalancing dates, with both month and year fixed effects, and standard errors are clustered at the year level. We do not plot the estimates for Brazil and Mexico because they cannot include month fixed effects owing to limited observations (these countries are dropped from the sample because they are often at the 10% cap). We also drop Argentina, whose exchange rate regime is classified beyond the standard free-floating regime. The x-axis is the exchange rate regimes ranging from de facto peg (left) to free floating (right). All regression estimates are significant at the 1% level. The circle size is proportional to the exchange rate volatility of the currency in our sample.

Figure 2A: Fact 2a
Exchange Rate Response Correlates with Exchange Rate Volatility

(a) Exchange Rate Vol. and Regime



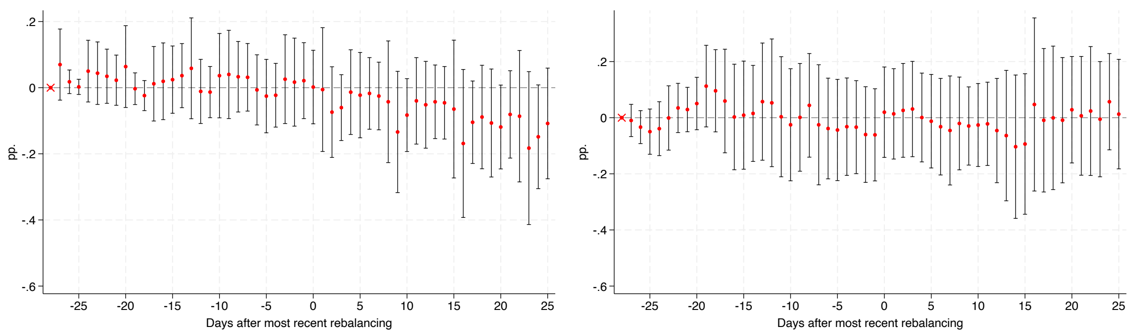
(b) Exchange Rate Vol. and $\beta_{\mu,c}$



Note: This figure presents the relation between (a) the country-specific exchange rate response to the currency demand shock (measured by $\mu_{c,t}$) and the exchange rate volatility and (b) the relation between the exchange rate regime and exchange rate volatility. The red line is the fitted regression for the x- and y-axis variables.

Figure 3:
Fact 3: Policy Rates and Yields Do not Respond to the Currency Demand Shock

(a) Policy Rates (b) Three-Month Deposit Yields



Note: Pooled regression coefficients of the change in monetary policy rates (in percentage points, left panel) and change in annualized three-month local-currency deposit yields (in percentage points, right panel) with 95% confidence intervals. Monetary policy rates and three-month deposit yields are provided at the daily frequency and are defined as the cumulative change from 28 days before the rebalancing date. The regressions in both figures include time and country fixed effects, with standard errors clustered at the date level.

the volatility of the exchange rates and as the exchange rate regimes move toward free floating.

Empirical Fact 3. *Policy rates and short yields do not respond to the currency demand shock.*

Another concern for identification is that central bank policy rates might respond to the rebalancings of the GBI-EM Global Diversified index. If the policy rates were to move, the macro-fundamentals and exchange rates would also respond, violating the exogenous nature of the currency demand. We show that this is not the case.

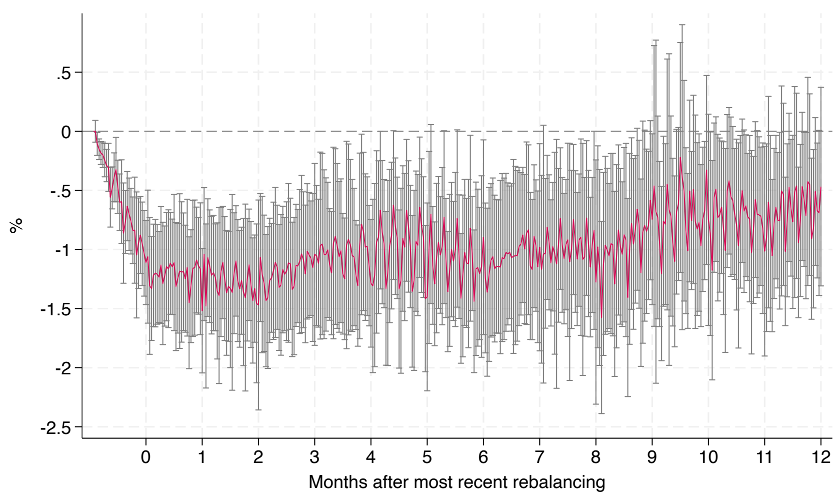
Central bank policy rates and short-term yields are not responsive to the exogenous currency demand shock.[19] The regression uses cumulative changes in central bank policy rates (since 28 days before the rebalancing event) on the currency demand shock and gives insignificant coefficients for the cross-country sample in Figure 3 (a). The results make clear that the central banks are not using monetary policy rates to offset the

[19]Our finding does not invalidate the results by Pandolfi and Williams (2019), who find that a one standard deviation in the flows implied by the rebalancings of the GBI-EM Global Diversified index leads to an increase in sovereign debt prices of 8 basis points for *long-term* government bonds. Their results make sense as the bonds in the index are long-term government bonds with at least 13 months to maturity by design. Different from their analysis, our regression uses short-term yields and policy rates.

exchange rate appreciation due to the rebalancings of the index. In addition, Figure 3 (b) shows that changes in short-term local-currency deposit yields have an insignificant response to the currency demand shock.[20] The regressions in both figures include time and country fixed effects, with standard errors clustered at the date level.

Empirical Fact 4. *The currency demand shock has a persistent and significant effect on exchange rates, lasting at least three months after a rebalancing event.*

Figure 4:
FACT 4: CURRENCY DEMAND SHOCK HAS PERSISTENT EFFECTS ON EXCHANGE RATES



Note: This figure plots the estimated coefficients of the change in the cumulative exchange rate on the currency demand shock measured by $\mu_{c,t}$ in the four-month horizon after a rebalancing event; $\mu_{c,t}$ is standardized by its mean and standard deviation in the regression. The dependent variable is the change in cumulative exchange rates starting from 28 days prior to the first rebalancing event. All regressions are performed in a pooled OLS using time and country fixed effects, with standard errors clustered at the date level. The results are shown as point estimates (red dots) with 95% confidence interval (blacks bars) for each regression.

We extend the regression window in equation (2) to *months* after a rebalancing event at time 0 while keeping the same definition on cumulative change of exchange rates $e_{c,t+d}$ (defined as the cumulative change since 28 days before the rebalancing date). Figure 4 gives away two key messages. First, the effects of rebalancings on exchange rates do not disappear immediately; instead, they remain significant for at least three months

[20]We also report in the appendix (Figure B.2) that local-currency government bond yields relative to USD do not respond significantly to $\mu_{c,t}$.

after a rebalancing event. Second, same as the short-run response in Fact 1, there's no additional impact from the currency demand shock ($\mu_{c,t}$) on exchange rates after the first rebalancing event at time 0, as we see that the response remains flat for the months after the first rebalancing. This is because once the first rebalancing is implemented, there's no additional flow pressure on exchange rates from this rebalancing event.[21]

Remark 5. *Why does the currency demand shock have persistent effects on exchange rates?*

One reason that the effects are long-lasting is due to the persistence of the shock. As reported in Table B.8[22], the country-specific index weights follows an AR(1) process with an average persistence of 0.792, with their residuals following a close-to white noise process.[23] In addition, our regression captures a level shift in exchange rates (starting 28 days before a rebalancing event) and there are thus *no* gains of excess returns for the arbitrageurs in the financial market even if the effects persist for three months.

Remark 6. *Can other local-currency emerging market sovereign bond indices also contribute to the observed exchange rate movements?*

We examine the rebalancing mechanisms of all major local-currency emerging market sovereign bond indices. While most indices employ different rebalancing schemes and timing than the GBI-EM Global Diversified index, the EMGBI-Capped follows a similar rebalancing pattern.[24] However, our analysis of the EPFR dataset reveals that funds

[21]We do not control for macro-fundamentals because our variables (such as GDP and net foreign asset positions) are much more slow-moving compared with exchange rates, and including them does not alter the baseline results. We also show formally in Table B.10 that the macro-fundamentals do not respond to the currency demand shock.

[22]In Table B.8, we test the persistence of GBI-EM Global Diversified weights ω_{ct} for each country; we also conduct Portmanteau white noise tests for the residuals of ω_{ct}, which match the definition of μ_{ct} as percentage changes in ω_{ct}.

[23]The fact that the currency demand shocks are persistent over time provides a powerful framework for us to draw implications on foreign exchange interventions, as discussed in the second half of the paper. Fratzscher, Menkhoff, Sarno, and Stohr (2019) document that 68.6 percent of intervention days are preceded by an intervention in the same direction from the same central bank on the day before (and 86.9 percent during the three days before), with the average length of all intervention episodes lasting 4.5 days (and average length for free floaters of 9.2 days). They also show that interventions do not only happen during crises, with 0.225 of their sampled days are covered in turbulent times.

[24]The EMGBI-Capped index is also called the Russell FTSE Emerging Markets Government Bond Index. It was introduced in 2018 as a rebranding of an older Citi Group WGBI index and is an emerging market local-currency government bond index that also has an end-of-month country weight cap at 10%.

tracking the EMGBI-Capped represent less than 10% of the assets tracking the GBI-EM Global Diversified index, making their impact negligible.

Remark 7. *Do co-movements of macro-fundamentals across countries invalidate identification?* One might suspect strong positive co-movements in sovereign bond prices among our sample countries. However, Table B.9 demonstrates significant heterogeneity in local-currency sovereign bond prices at rebalancing dates, even within the inflation-targeting Latin American countries. Moreover, given that index rebalancing occurs monthly throughout our decade-long sample, it is improbable that any pair of countries' prices consistently move in the same direction at each rebalancing event.

4 Inelastic Financial Markets and Implications for Foreign Exchange Interventions

This section connects our empirical facts to models of inelastic financial markets. We first argue that models with perfectly elastic financial markets cannot explain the observed empirical facts on our currency demand shocks and exchange rates. We show that a stylized inelastic financial markets model with endogenous deviations in the uncovered interest parity (UIP) condition can rationalize our empirical facts. We also discuss the implications for foreign exchange interventions and the "relaxed trilemma."

4.1 Empirical Evidence for Inelastic Financial Markets

There are two types of models with perfectly elastic markets. First, in classical macroeconomic models (Mundell, 1962; Obstfeld and Rogoff, 1995; Gali and Monacelli, 2005) where UIP holds, a currency demand shock plays no role in determining neither the path of exchange rates nor short-term interest rate differentials. Another class of macroeconomic models violate the UIP condition but not the assumption of perfectly elastic financial markets. Those models typically have capital control taxes and risk-premium shocks (Devereux and Engel 2002; Farhi and Werning, 2012) as exogenous deviations from the

UIP condition (i.e., *exogenous UIP shocks*) that move exchange rates but do not change the equilibrium allocation of assets.

We introduce the following modified UIP condition in Definition 1 that includes the exogenous UIP shocks. Let $i_{c,t}$ and $i^*_{c,t}$ be the returns of home- and foreign-currency bonds, respectively; $e_{c,t}$ is the exchange rate measured in the number of home currencies per US dollar (foreign); $\mathbb{E}_t \Delta e_{c,t+1}$ is the expected change in exchange rates from t to $t+1$.

Definition 1. *The modified UIP condition with exogenous shocks is given by*

$$i_{c,t} - i^*_{c,t} - \mathbb{E}_t \Delta e_{c,t+1} \;=\; \tau_{c,t} \;+\; \psi_{c,t} \quad , \tag{4}$$

where we denote capital control taxes by $\tau_{c,t}$, the exogenous risk-premium shock by $\psi_{c,t}$. Both $\tau_{c,t}$ and $\psi_{c,t}$ are exogenous UIP shocks.[25]

We show that a model where UIP holds or with exogenous UIP shocks cannot square with our stylized empirical facts. First, our empirical facts that that the currency demand shock moves exchange rates (Fact 1) but not interest rates (Fact 3) are direct invalidation of the UIP condition. Second, we show that both measures for capital taxes $\tau_{c,t}$ and risk-premium shock $\psi_{c,t}$ are immune to our exogenous currency demand shock $\mu_{c,t}$ in Table B.10. Intuitively, both capital control taxes and risk premium for macroeconomic conditions are slow-moving variables compared with the exogenous currency demand shocks and therefore cannot account of the high-frequency movements in exchange rates.

Only in models with inelastic markets where a currency demand shock generates *endogenous* UIP deviations ($\Lambda_{c,t}$) in Definition 2 can account for our observed empirical facts. In these models, a currency demand shock changes the positions of the arbitrageurs who conduct currency carry trade in a segmented financial market. As the arbitrageurs have limited risk-bearing capacity, a currency demand shock translates into movements in exchange rates, changes in the equilibrium allocation of assets, and gives rise to endogenous deviations from the UIP condition.

[25]Without loss of generality, we use net capital tax defined as the difference between home and foreign capital tax. We follow Farhi and Werning (2012) and introduce risk-premium shock as a combination of the export demand shocks and consumption shocks that changes the macro-fundamentals of the economy. An example of risk-premium shock is a change in the growth rate of world consumption that increases the world interest rate, without changing the equilibrium allocation of assets and exchange rates.

Definition 2. *The modified UIP condition with both exogenous and endogenous deviations is*

$$i_{c,t} - i^*_{c,t} - \mathbb{E}_t \Delta e_{c,t+1} = \underbrace{\tau_{c,t} + \psi_{c,t}}_{exogenous} + \underbrace{\Lambda_{c,t}}_{endogenous}, \tag{5}$$

where we denote capital control taxes by $\tau_{c,t}$, the exogenous risk-premium shock by $\psi_{c,t}$, and $\Lambda_{c,t}$ is the component of endogenous UIP deviation. Both $\tau_{c,t}$ and $\psi_{c,t}$ are exogenous UIP shocks, and the endogenous risk-premium is given by $\Lambda_{c,t}$.

4.2 Interventions in Inelastic Foreign Exchange Markets

Our empirical facts point to a model with endogenous UIP deviations in inelastic financial markets. In this section, we present a stylized model following Gabaix and Maggiori (2015) and Itskhoki and Mukhin (2021) where a currency demand shock gives rise to endogenous UIP deviations. We then discuss implications for foreign exchange intervention policies in these markets.

4.2.1 Key Model Equations for Inelastic Markets

Consider a small open economy, denoted by c. There are four types of agents in a partially segmented financial market where both home and foreign households can hold only government bonds of their own currency. Households demand home-currency bonds $b_{c,t}$, which are shaped by the macroeconomic fundamentals in the economy. There are also three types of agents who can trade both home- and foreign-currency bonds in the international financial market, namely, noise traders (with positions $n_{c,t}$ and $n^*_{c,t}$ for home- and foreign-currency bonds), arbitrageurs ($d_{c,t}$ and $d^*_{c,t}$), and the government ($f_{c,t}$ and $f^*_{c,t}$), and we assume without loss of generality that they all reside in the home country. We describe the problem of each of these agents in detail in Appendix C.1. Our model notations follow closely Itskhoki and Mukhin (2021, 2023a).

The arbitrageurs' holdings and market clearing condition for the home-currency bonds are as follows

$$i_{c,t} - i^*_{c,t} - \mathbb{E}_t \Delta e_{c,t+1} \; - (\tau_{c,t} \; + \; \psi_{c,t}) \; = \; \underbrace{\lambda_{c,t} \, d_{c,t}}_{=\Lambda_{c,t}} \tag{6}$$

$$b_{c,t} \; + n_{c,t} \; + \; d_{c,t} + \; f_{c,t} = 0, \tag{7}$$

where the endogenous UIP component $\Lambda_{c,t} = \lambda_{c,t} d_{c,t}$, and $\lambda_{c,t} = \omega \; \sigma^2_{e_{c,t}}$ governs the arbitrageurs' risk-bearing ability; parameter ω is the arbitrageurs' risk-aversion coefficient, and $\sigma^2_{e_{c,t}}$is the equilibrium volatility of exchange rates.

An exogenous local-currency demand shock (an increase in $\mu_{c,t}$)[26]shifts noise traders' positions $n_{c,t}$ and affects arbitrageurs' holdings through the market clearing condition (7). In other words, the exogenous currency demand shock traces out the slope of the demand curve and arbitrageurs' risk-bearing capacity $\lambda_{c,t}$ in equation (6). The larger the $\lambda_{c,t}$, the lower the risk-bearing capacity, the more inelastic the financial markets, and the steeper the currency demand curve.[27] The risk-bearing capacity is modeled as a function of equilibrium volatility of exchange rates due to our empirical evidence on exchange rate responses strongly correlate with exchange rate volatility (Fact 2).

4.2.2 Policy Function of Foreign Exchange Interventions

We show empirically that foreign exchange interventions data do not respond to our currency demand shock $\mu_{c,t}$. We characterize the policy function of foreign exchange interventions to (counterfactually) stabilize exchange rates by offsetting noise trader shocks.

We follow Itskhoki and Mukhin (2023a) and assume the policy function of foreign exchange (FX) interventions from the government follows a "Taylor-Rule" formula

$$f_{c,t} = -\alpha^f_c \; n_{c,t} + \; \epsilon^f_{c,t} \; , \tag{8}$$

where $\alpha^f_c \in [0, 1]$ is the FX intervention intensity in offsetting the noise trader shocks

[26]Appendix A.3 discusses how to connect the currency demand shocks from our empirical data to the shocks in noise trade positions in the model.

[27]In the limit that $\lambda_{c,t} \to \infty$, the international bonds market is completely segmented, with financial autarky. On the other extreme, when $\lambda_{c,t} = 0$, the arbitrageurs are able to take infinite positions and absorb any nonzero excess returns in the currency carry trade. In the case when $\lambda_{c,t} \in (0, \infty)$, the model endogenously generates UIP deviations given by the arbitrageurs' limited risk-taking capacity.

in $n_{c,t}$; and $\epsilon^f_{c,t}$ is the discretionary component of FX intervention that's independent of the noise trader shocks.

We find no correlation between FX intervention data and our exogenous currency demand shock $\mu_{c,t}$ for nearly all countries in our sample. To estimate the intervention intensity α^f_c in the policy function, we use monthly FX interventions data from Adler, Lisack, Mano (2019) and regress that on our currency demand shock.[28] As shown in Table B.11, the lack of correlation suggests that the central banks are *not* actively using FX interventions in the spot market to offset the noise trader shocks from the exogenous currency demand. This also means that α^f_c is zero in equation (8) and the FX interventions we observe in the data come from the discretionary component $\epsilon_{c,f}$.

A central bank trying to stabilize exchange rates would need to use FX interventions to exactly offset the noise trader shocks, at the same magnitude and persistence, to ensure that $\partial e_{c,t}/\partial f_{c,t} = -\partial e_{c,t}/\partial n_{c,t}$. Thus, by estimating the response of exchange rates to the noise trader shocks, we obtain the counterfactual required size of FX interventions to stabilize exchange rates. The statement also requires all variables on the right-hand side of equation (A.7) (eg., interest rates, capital control taxes, macro-fundamentals) to be immune to the currency demand shock that moves noise traders' positions, which we have already demonstrated in the previous section.

To arrive at the closed-form of the policy function of foreign exchange intervention, we provide solutions from two model examples in appendix C — a partial equilibrium model under the Taylor rule (Engel and West, 2005) and a general equilibrium model with a fully specified goods market and the country's intertemporal budget constraint (Itskhoki and Mukhin, 2021). Through solving these models, one can then match the estimated regression coefficient in the empirical specification in equation (3) with the impulse response function of exchange rates in response to the noise trader shocks.

[28] We use estimated spot foreign exchange interventions data over GDP at monthly frequency from Adler, Lisack, Mano (2019) updated to 2021. The estimated interventions data use changes in central bank's reserves as proxies for FX interventions in the spot market. Thailand is the only country in our sample whose proxied foreign exchange intervention data have a significant response to the currency demand shock.

4.3 Discussion on the Relaxed Trilemma

We discuss the implications of foreign exchange interventions and the trilemma condition under inelastic financial markets. Under inelastic financial markets, foreign exchange intervention serves as an effective policy tool to stabilize exchange rates without compromising monetary policy independence, regardless of the capital controls, breaking down the classical trilemma of Mundell (1962). We refer to this condition as the "relaxed trilemma," using the language of Itskhoki and Mukhin (2023a).

4.3.1 The Relaxed Trilemma and Related Literature

Definition 3. *The relaxed trilemma constraint states that it is possible to have all three of the following conditions simultaneously: an independent monetary policy (inward focused on domestic inflation and output gap), free capital mobility (absence of capital control taxes), and a managed exchange rate. By contrast, under the classical trilemma constraint it is possible to have only two of the three conditions simultaneously.*

Definition 3 contradicts the classical trilemma constraint (Mundell, 1962), which states that it is *not* possible to have all three conditions in definition 3. Two types of models are subject to the classical trilemma constraint: the models where the UIP condition holds and the models with exogenous UIP shocks. If UIP holds, there is free capital mobility by construction and the economy faces a direct trade-off between an independent monetary policy and a fixed exchange rate, as seen in equation (6). If the UIP deviations come from exogenous shocks, monetary policy rates would have to move one on one with exchange rates unless capital control taxes ($\tau_{c,t}$) and exogenous risk premium ($\psi_{c,t}$) can both be used as policy instruments to offset exchange rates; however, this is clearly not feasible. Thus, under the trilemma constraint, exchange rate stabilization comes at the cost of compromising monetary policy independence.

By contrast, models with endogenous UIP deviations can have all three conditions in the trilemma met, because these models have FX interventions as an additional policy instrument to stabilize exchange rates. Under inelastic financial markets, FX interventions shift the arbitrageurs' positions, which then lead to endogenous deviations in UIP and move exchange rates. Therefore, the central bank can now stabilize exchange rates through foreign exchange interventions while the monetary policy is entirely domestically focused to close the output gap.[29] Even with perfectly mobile capital flows, the economy no longer has to compromise monetary policy independence to stabilize exchange rates, relaxing the classical trilemma constraint.

Taken together, FX interventions are ineffective in models where UIP holds or with exogenous UIP shocks only, as these models are subject to the classical trilemma tradeoff (*trilemma-type models*). By comparison, FX interventions become an effective in models with inelastic financial markets and endogenous UIP deviations (*non-trilemma–type models*). We summarize this concept in Definition 4 and present the related papers in classical trilemma models, models with exogenous UIP shocks, and models with endogenous UIP deviations in Table 4.1.

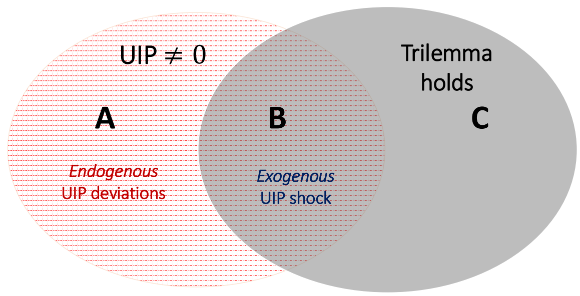
Definition 4. *Trilemma-type models are UIP models that bind under the classical trilemma constraint; non-trilemma–type models are UIP models that hold under the relaxed trilemma constraint. Foreign exchange interventions are only effective in non-trilemma-type models.*

4.3.2 Empirical Evidence for the Relaxed Trilemma

Our empirical facts show that there are significant exchange rate responses to the exogenous currency demand shock of almost all currencies but no response of the policy rates. Under trilemma-type models, the movements in exchange rates must be offset one on one by monetary policy rates for exchange rates to be fixed, for any given capital control taxes. Our evidence provides empirical support for non-trilemma–type models and implies that countries under managed exchange rate regimes (namely, de facto peg, crawling peg, and managed floaters) have used instruments other than monetary policies to manage their exchange rates in response to the currency demand shocks from index rebalancings.

[29]The central bank's objective is to minimize the international risk-sharing wedge and domestic output gap. Appendix C provides detailed illustration on central bank's problem.

Table 4.1: Trilemma Constraint and UIP



Model	Financial market	FXI effective?	Related Papers
endogenous UIP deviation	(imperfectly) inelastic	yes	Gabaix and Maggiori (2015), Itskhoki and Mukhin (2021), Fanelli and Straub (2021), Basu et al. (2023), Masao, Nakamura, and Steinsson (2024)
exogenous UIP shock	perfectly elastic	no	Devereux and Engel (2002), Farhi and Werning (2012)[a]
classic trilemma (UIP = 0)	perfectly elastic	no	Mundell (1962), Dornbusch (1976), Obstfeld and Rogoff (1995), Gali and Monacelli (2005)

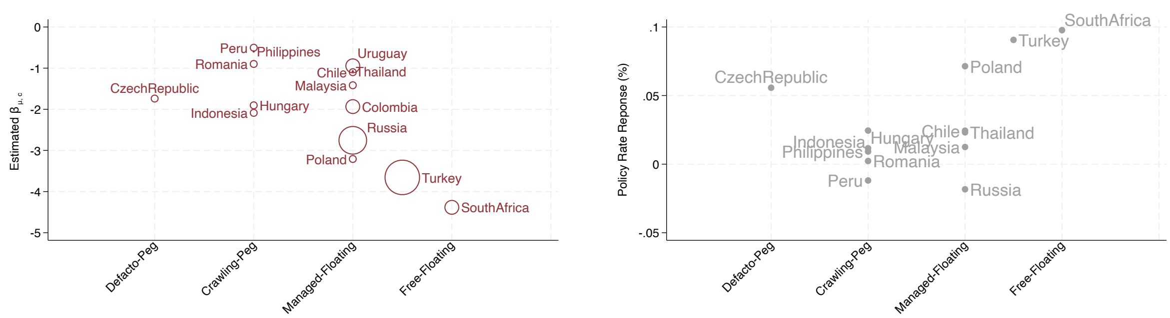
Note: This diagram presents the relation between models where UIP fails (left circle) and models where the trilemma constraint holds (right circle). Region A refers to models under the relaxed trilemma (classic trilemma fails) and UIP fails under endogenous UIP deviation; region B refers to models where UIP fails due to exogenous UIP shocks but the trilemma holds; region C represents the classic trilemma models where UIP holds. One can also think of region A as the *non-trilemma-type models* in Definition 4 where foreign exchange interventions (FXI) are effective. Both region B and C would belong to the *trilemma-type models* where foreign exchange interventions are ineffective as they are subject to the trilemma constraint and their financial markets are elastic. The references for each type of models are listed. Our empirical results are consistent with models in region A (*non-trilemma-type models*).

[a]We do not add the "convenience yield" models such as Jiang, Krishnamurthy, and Lustig (2018) into the table above. The convenience yields on US treasury are modeled as exogenous UIP shocks in their partial equilibrium of asset markets, which are also assumed to be imperfectly inelastic.

Table [4.2](#) puts both the response of exchange rates and policy rates to the currency demand shock side by side and summarizes this result. While there's a clear downward trend between exchange rate response and its regime, policy rates response is insignificant regardless of the exchange rate regime. We view this as the most direct piece of evidence supporting the relaxed trilemma constraint.

Table 4.2: Empirical Support for Relaxed Trilemma

(a) Exchange Rates (b) Policy Rates



Note: Scatterplot of country-specific exchange rates (left) and policy rates (right) response to the currency demand shock $\mu_{c,t}$ against the exchange rates regime (from strict to relaxed) as classified by Ilzetzki, Reinhart and Rogoff (2021). Exchange rates responses are significant for all countries; policy response are insignificant for all countries, regardless of the exchange rate regimes. Regressions are estimated at the rebalancing dates and the dependent variables in both figures are defined as the cumulative change in exchange rates (or policy rates) since 28 days before the rebalancing dates.

5 Identifying the Size of Foreign Exchange Interventions

In this section, we identify the required size of foreign exchange intervention to stabilize exchange rates and discuss its effectiveness across different exchange rate regimes. We find that free floaters are twice more effective at stabilizing exchange rates on average compared with managed floaters or peggers, with free floaters requiring lower amount of reserves to stabilize exchange rates.

5.1 Converting the Estimates to the Size of the Interventions

We use the cross-country estimates from Fact 1 to compute the implied capital flows from the rebalancings. This allows us to report the average counterfactual required size of foreign exchange interventions to stabilize exchange rates.

The caveat in this exercise is that our currency demand shock is measured in change in country weights, whereas the required intervention is in capital flows. Fact 1 shows that a one standard deviation of $\mu_{c,t}$ (4.844% change in country weight) moves exchange rates by 1.2% after one rebalancing for pooled the regression. We use this number to compute the *average* implied capital flows across countries, together with the average country weight in the index (6.36%) and the total mutual funds positions in the EPFR tracking the index in 2019 (120 billion USD), as well as the fact that EPFR represents about 60% of the global mutual fund population in the Investment Company Institute (ICI) Fact Book (Table B.3).

We form the following equation to convert our estimate into the USD flows required to stabilize exchange rates by 1%

$$\frac{1}{\beta_\mu} \times \text{std.}(\mu_{c,t}) \times \text{average country weight } \times \frac{\text{EPFR mutual fund positions}}{\text{Share of EPFR funds in ICI}}.$$

The results of our pooled OLS regression imply that the average required foreign exchange (FX) interventions to move exchange rates by 1% is $\frac{1}{1.2} \times 4.844\% \times 6.36\% \times \frac{120}{0.6} =$ 0.51 billion USD, or about 0.1% of the average annual GDP in 2019 (the average annual nominal GDP in 2019 is 500 billion USD, reported in Table B.1). In other words, the estimated currency elasticity for the cross-country sample is about 0.5.

Our estimates on cross-country average currency elasticity are largely consistent with the literature, but more on the lower (inelastic) side. Adler, Lisack, and Mano (2019) use FX intervention episodes and that an intervention with magnitude of 1% of GDP results in exchange rate depreciation of 1.7% to 2% (or the size of 0.5% of GDP to move exchange rates by 1%). Their results are larger than ours likely because their sample consists of both advanced and emerging countries (with the former having much smaller exchange rate volatility, and thus more elastic markets), while our sample consists of only emerging markets. For the same reason, most papers on currency demand estimation (Evans and Lyons, 2002; Hau, Massa, and Peress, 2009; Camanho, Hau, and Rey, 2021) that mainly focus on advanced economies (or a mix of advanced and emerging economies) also have estimates larger than ours.

5.2 Size of Interventions for Different Currency Regimes

Our key contribution is documenting the heterogeneity in the required size of foreign exchange interventions across different exchange rate regimes. Using country-specific estimates from Fact 2, we calculate the required size of FX interventions to stabilize exchange rates fore each country. This calculation incorporates each country's specific response to currency demand shocks and its individual average weight in the GBI-EM Global Diversified index. Table 5.1 reports the counterfactual required intervention size as a share of GDP for each country. The required size of intervention as a share of broad money supply (M2) is reported in Table B.2.

We find that to stabilize exchange rates, free floaters require an intervention that is half as large (as a share of GDP) as that required by managed floaters or peggers; thus, free floaters are more effective at using FX interventions. For example, the required FX intervention to move exchange rates by 1% is about 0.038% of GDP for free floaters (0.028% for Turkey and 0.048% for South Africa), while the group average required intervention is 0.069% of GDP for managed floaters and 0.128% of GDP for peggers.

Our results on FX interventions being more effective for free floaters are consistent with the event studies of FX interventions. Using confidential intervention data from 33 countries, Fratzscher, Gloede, Menkhoff, Sarno, and Stohr (2019) determine the *success* of interventions (defined as the consistency in the movement of exchange rates during the intervention and its intended direction) across different regimes and find that interventions are most effective for free floaters, with a success rate of 0.53 through pure purchase or sale of foreign exchange reserves. By comparison, the success rate for broad band, narrow band, and other exchange rates regimes are significantly lower.

Our empirical results are consistent with the model mechanism in section 4. The risk-bearing capacity of the financiers governs the elasticity of the exchange rate response to the currency demand shock. A more stable or managed exchange rate would therefore imply smaller exchange rate volatility and thus a more elastic market. In the limit of exchange rates being fully pegged, we are back to the elastic financial market model under the trilemma constraint where exchange rates are immune to currency demand

Table 5.1: FX Interventions Required To Induce A 1% Exchange Rate Change

Country	ER regime (code)	ER Vol.	FXI	FXI / GDP
Czech Republic	de facto peg (1)	0.112	0.214	0.084%
Peru	crawling peg (2)	0.113	0.548	0.238%
Hungary	crawling peg (2)	0.160	0.464	0.287%
Romania	crawling peg (2)	0.131	0.260	0.104%
Indonesia	crawling peg (2)	0.187	0.453	0.040%
Philippines	crawling peg (2)	0.074	0.065	0.017%
Thailand	managed floating (3)	0.053	0.669	0.119%
Malaysia	managed floating (3)	0.134	0.545	0.148%
Colombia	managed floating (3)	0.271	0.268	0.083%
Chile	managed floating (3)	0.169	0.107	0.041%
Poland	managed floating (3)	0.116	0.285	0.047%
Uruguay	managed floating (3)	0.273	0.016	0.028%
Russia	managed floating (3)	0.372	0.236	0.013%
Turkey	managed floating/free falling (3.5)	0.582	0.203	0.028%
South Africa	free floating (4)	0.286	0.194	0.048%
Group average				
	peg / crawling peg	0.130	0.334	0.128%
	managed floating	0.198	0.304	0.069%
	free floating / falling	0.434	0.198	0.038%
Whole sample average			0.302	0.088%

Note: This table reports the country-specific required size of foreign exchange intervention (FXI) to stabilize exchange rates by 1%, in billions of US dollars (column 4) and as a share (%) of each country's 2019 nominal GDP (column 5). The exchange rate volatility (ER Vol.), measured as the standard deviation of the log exchange rate level by country, is reported in column 3. In column 2, we sort countries by their coarse exchange rate (ER) regimes (as classified by Iltzetki, Rogoff and Reinhart 2021) from de facto peg to free floating. For countries having multiple exchange rate regime codes during our sample period (2010–2021), as for Mexico and Turkey, we take the average regime code across time.

The required size of intervention is computed using the country-specific exchange rate response to the currency demand shock in the 0–5 day horizon after the rebalancing date. All estimates are significant at the 1% level. A table with the country's GDP, market value in the GBI-EM Global Diversified index, and broad money supply (M2) can be found in the appendix (Table B.1).

shocks. In other words, FX interventions are more effective for floaters precisely because they have larger exchange rate volatility (Fact 2) and a more inelastic financial market, and are thus further away from the trilemma constraint.

Remark 8. *What types of FX interventions can our quasi-natural experiment best speak to?*

The exogenous currency demand shock from our quasi-natural experiment would be most analogous to a sterilized FX intervention in the spot exchange market. Similar to the open market operations in the spot exchange market, the index rebalancings create currency demand shocks that move exchange rates as the mutual fund investors buy or sell their positions of local-currency government bonds. The fact that we find the monetary policy rates are not moving with respect to the currency demand shock makes the experiment most suitable for understanding the effects of sterilized intervention.

6 Conclusion

In this paper, we use a well-identified currency demand shock on noise traders that gives rise to endogenous uncovered interest parity deviations under an inelastic financial market. This finding provides direct support for models with inelastic financial markets and the relaxed trilemma constraint. We assess the effectiveness of foreign exchange interventions for an emerging-market central bank for stabilizing exchange rates under the inelastic financial market hypothesis. When markets are inelastic, foreign exchange intervention works as an additional policy tool to move exchange rates without compromising monetary policy independence, providing evidence relaxing the classical trilemma constraint. Our results contribute to various strands of literature including those on foreign exchange intervention and asset demand estimation, and are informative for policymakers at emerging market central banks.

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Online Appendix

A Data Description and Background

A.1 More on the GBI-EM Global Diversified Index

The GBI-EM Index Family

Published by J.P. Morgan in 2005, the GBI-EM Global Diversified index is the largest local-currency government bond index for emerging countries. It is also the most popular index among the GBI-EM family of six local-currency emerging market government bond indices: three basic versions (i.e., GBI-EM Broad, GBI-EM Global, and GBI-EM Narrow) and a diversified version for each. Each diversified version is created from the corresponding basic version by maintaining the same set of countries but with more equalized country weights for the purpose of reducing market concentration risks. Among all basic versions, GBI-EM broad has the broadest coverage of countries, followed by GBI-EM Global, and then GBI-EM Narrow. The three basic versions are compared in Table A.1.

Table A.1: Three Versions of the J.P. Morgan GBI-EM Indices

	GBI-EM Broad	GBI-EM Global	GBI-EM Narrow
Explicit capital control	✓		
Tax/Regulatory constraints	✓	✓	
Direct access by foreigners	✓	✓	✓
No. countries as of 2021	21	19	16
Country criteria	GNI per capita below the IIC[a] for 3 consecutive years		
Instrument criteria	Fixed/Zero coupon; Maturity $>$ 13 months Minimum face amount $>$ US \$1 bn.		

Source: J.P. Morgan Market Reports

[a] Index income ceiling for emerging countries

Apart from having different restrictions on capital controls and tax regulations for different versions of the GBI-EM index, all versions have the same control on income capita and credit ratings. A country is chosen to enter (and remain in) the GBI-EM

Global diversified index if the country's gross national income (GNI) per capita is *below* the J.P. Morgan-defined index income ceiling (IIC) for three consecutive years. A country is chosen to exit the index if the country's GNI per capita is above the IIC for three consecutive years and if the country's long-term local-currency sovereign credit rating (the available ratings from S&P, Moody's, and Fitch) is A-/A3/A- (inclusive) or above for three consecutive years. In addition, the government bonds included in the index have to be in local currency and have month-to-maturity of over 13 months as the threshold.

Three Layers of Rebalancings

The monthly rebalancings of the GBI-EM Global Diversified index have three layers, which are done in order on the last weekday of the month. The first layer uses a diversification methodology that includes in the index only a portion of a country's current face amount outstanding. This value — called the adjusted face amount — is based on the respective country's relative size in the index and the average size of all countries in the index. The adjusted face amount is then used to compute the market value of each country in the index. The second layer focuses on the bond maturity threshold that drops from the index the bonds with fewer than 13 months to maturity. As the third and last layer of control, the index rebalancing caps the weight of each country at 10%, computed using the adjusted face amount.

How is the Adjusted Face Amount Computed?

In the first layer of rebalancing, the adjusted face amount in the diversified version is created from its corresponding basic version with different weighting strategies for countries. Specifically, the following formula is used to construct the diversified country face amount (FA_c^D) for country c

$$FA_c^D = \begin{cases} ICA \times 2 & \text{if} \quad FA_{\max} \\ ICA + \frac{ICA}{FA_{\max} - ICA}\left(FA_c - ICA\right) & \text{if} \quad FA_c > ICA \\ FA_c & \text{if} \quad FA_c \leq ICA, \end{cases} \tag{A.1}$$

where FA_c is the face amount of country c. Additionally, ICA is the average face amount of the countries (or currencies) in the index

$$ICA = \frac{\sum_c \text{Country face amount}}{\text{No. countries in the index}}. \tag{A.2}$$

The computation of country-level weights in the index relies on diversified face amounts (FA^D). For each country, its market value is calculated by multiplying its FA^D by the dirty price (which includes both price and accrued interest). The country's weight is then determined by dividing its market value by the total market value of the index. When comparing diversified and non-diversified versions of the GBI-EM, small countries (those with $FA_c \leq ICA$) have identical market values in both versions but receive higher weights in the diversified version because the diversified index's total market value is smaller. Larger countries have lower market values in the diversified version due to the adjustment on the country-level face amount from A.1 and the 10% country weight cap.

How Often Are the Weights Adjusted?

Both the diversified and non-diversified versions update their weights daily based on changes in bonds' dirty prices, which affect market values. However, the diversified version has two additional monthly adjustments that occur only at month-end on the rebalancing dates: the rebalancing of countries' diversified face amounts and the application of the 10% weight cap. These month-end adjustments create weight changes beyond those caused by daily price movements. Between monthly rebalancing dates, the diversified face amounts remain fixed, so any weight changes in the diversified version during this period reflect only changes in market returns (dirty prices).

How are Bonds Deleted from the Index?

The only bonds removed from the index are those that no longer meet the maturity threshold during rebalancing. The face amount rebalancing process, which distinguishes the diversified from non-diversified versions, adjusts bond quantities but preserves the same set of bonds in both versions. For bonds from countries above the ICA threshold, their face amounts are reduced in the diversified version, while bonds from countries below the ICA remain unchanged. This means identical bonds should have the same yields and returns in both versions, but may differ in market value. For instance, a Chinese bond in January 2022 would show the same yield and returns in both the GBI-EM Broad and its diversified version, but have a smaller market value in the diversified version due to its reduced face amount. Conversely, a Philippines bond would have identical metrics across both versions, including market value, since its face amount falls below the ICA and thus remains unchanged in the diversified version.

A.2 Estimating the Aggregate Flows from EPFR

To estimate the total assets under management (AUM) of mutual funds tracking the GBI-EM Global Diversified index, we first identify relevant funds in the EPFR dataset and then scale up their assets using the EPFR database's share of the global mutual fund market. Table B.3 presents our empirical findings: panel (a) shows the AUM of EPFR mutual funds that passively track the index, while panel (b) shows EPFR's share of the Investment Company Institute (ICI) database, which we use to scale up the AUM to estimate the total industry-wide assets tracking the index.

To identify mutual funds in the EPFR dataset that track the index, we first filter the dataset for emerging market bond funds that specifically benchmark against variations of the J.P. Morgan GBI-EM Global Diversified index. These variations include the standard GBI-EM Global Diversified, its composite version, ESG version, and regional versions (Europe or LATUM, Asia). We exclude funds that benchmark against other indices in the GBI-EM family (such as "GBI-EM Broad") or the investment grade version of the GBI-EM Global Diversified.

To refine our selection, we conduct a regression analysis of monthly returns for each bond fund in the EPFR dataset against the returns of the GBI-EM Global Diversified index and select funds with a performance R-squared of at least 0.9. Our final dataset combines funds that explicitly benchmark against the GBI-EM Global Diversified index variations and funds that meet this R-squared threshold, resulting in 2,113 unique funds.

To measure the degree of passive management in our mutual fund dataset, we apply the R-squared performance method developed by Amihud and Goyenko (2013). This method involves regressing fund-level monthly returns against the GBI-EM Global Diversified index returns. We conduct these regressions using a 12-month rolling window from January 2016 to January 2022, rather than the entire time series, to account for potential time variation in fund passivism. Table B.4a presents the histogram of the estimated R-squared values for the mutual fund performance. Our analysis shows that the mutual funds in our dataset have a median R-squared of 0.9.

To further verify fund passivism, we create a hypothetical fund by computing the AUM-weighted average returns of all mutual funds identified as closely tracking the GBI-EM Global Diversified index from the EPFR dataset. As shown in Table B.4b, this constructed fund's monthly returns closely mirror those of the GBI-EM Global Diversified index, with an OLS regression yielding an R-squared of 0.97. Table B.4a and B.4b collectively demonstrate that these funds must adjust their positions at rebalancing dates to match the index returns, following the previously discussed rebalancing scheme. The evolution of these funds' total assets under management from 2016 to 2022 is illustrated in Table B.3 panel (a).

The final step in computing the aggregate flows of the mutual funds tracking the index is to estimate the population share of the EPFR data in the ICI dataset. The Investment Company Fact Book reports the global mutual fund data population. We aggregate equity, bonds, and money market end-of-month assets for both industrialized and emerging markets from the EPFR data and divide the number we obtain by the global mutual fund data population from the Investment Company Fact Book. The result is the population presentation of the EPFR data in the global mutual fund industry, as reported in panel (b) of Table B.3.

A.3 Converting the Currency Demand Shocks into Noise Trader Shocks

We show how to connect the flows implied by rebalancings (FIR) with the noise traders' positions. Given that we do not observe the entire variation in the noise trader shocks, we decompose noise traders' positions $n_{c,t}$ into two components: The first is the buy-and-hold portfolio of benchmark investments that are subject to mechanical rebalancings ($\tilde{n}_{c,t}$). The second is the part of the noise traders' positions unexplained by rebalancings

($\tilde{e}_{c,t}$). The two components are *additive* and *orthogonal* to each other

$$n_{c,t} = \tilde{n}_{c,t} + \tilde{e}_{c,t}, \qquad \text{where} \quad \tilde{n}_{c,t} \perp \tilde{e}_{c,t}.$$

The holdings of benchmark investments ($\tilde{n}_t$) are subject to noise trader shocks ($\tilde{\psi}_t$) when rebalancing happens. These shocks are orthogonal to macroeconomic fundamentals, as illustrated in the model. The position $\tilde{n}_t$ at time t is

$$\tilde{n}_{c,t} = \begin{cases} \left(\frac{\tilde{n}_{c,t-1}}{R_{c,t-1}}\right) R_{c,t} & \text{o.w} \\ \tilde{\psi}_t R_{c,t} & \text{if} \quad t \text{ is the rebalancing date.} \end{cases} \tag{A.3}$$

At the rebalancing date,

$$\begin{aligned} \tilde{n}_{c,t} = \tilde{\psi}_{c,t} R_{c,t} &= \underbrace{\tilde{\psi}_{c,t} R_{c,t} - \left(\frac{\tilde{n}_{c,t-1}}{R_{c,t-1}}\right) R_{c,t}}_{\text{flows implied by rebalancings}} + \underbrace{\left(\frac{\tilde{n}_{c,t-1}}{R_{c,t-1}}\right) R_{c,t}}_{\text{market value buy-and-hold}} \\ &= \quad \text{FIR}_{c,t} \qquad\qquad + \qquad \text{market value}^{BH}_{c,t}, \end{aligned}$$

where market value$^{BH}_{c,t}$ is the buy-and-hold market value that equates the face amount of previous rebalancing $t-1$ times the market price at time t. We can therefore rewrite the noise trader shocks $n_{c,t}$ as

$$n_{c,t} = \text{FIR}_{c,t} + \quad \text{market value}^{BH}_{c,t} + \quad \tilde{e}_{c,t}, \tag{A.4}$$

where $\tilde{e}_{c,t} \perp \text{FIR}_{c,t}$; that is, the components of the noise trader shocks unexplained by rebalancings are orthogonal to the flows implied by the rebalancings of the index.

B Additional Figures and Tables

Table B.1: Country Statistics for Computing the Required Size of Intervention

Country	2019 mkt. value	2019 GDP	2019 broad money (M2)
Argentina	3.83	360.57	
Brazil	92.81	1833.49	1761.21
Chile	29.72	262.98	221.51
Colombia	62.86	321.81	157.54
Czech Republic	38.09	256.02	211.28
Hungary	40.20	161.72	94.05
Indonesia	92.06	1138.96	441.46
Malaysia	55.98	369.14	454.31
Mexico	92.81	1297.19	490.13
Peru	33.07	229.93	112.80
Philippines	2.63	384.63	294.62
Poland	82.88	602.6	412. 15
Romania	24.33	249.67	
Russia	74.20	1764.64	1042.48
South Africa	80.65	400.25	268.35
Thailand	82.92	560.20	691.15
Turkey	36.91	725.20	426.82
Uruguay	1.82	57.82	26.33
Sample average*	49.22	499.04	436.78

Note: The sample average (last row) excludes Argentina, Brazil, and Mexico during aggregation, because their estimates are not used to infer the required size of intervention in the main text. Column 2 gives the average market value of the local-currency government bonds of each country in the GBI-EM Global Diversified in 2019. Column 3 gives the annual nominal GDP of 2019. Column 4 gives the annual broad money supply (M2) in 2019 from the International Financial Statistics (IFS) of the International Monetary Fund, with the data for Argentina and Romania missing. All values are in billions of US dollars.

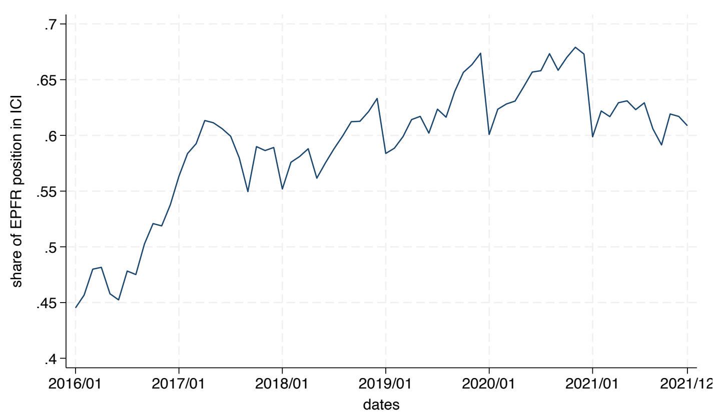
Table B.2:
Interventions Required to Induce 1% Exchange Rate Change (as a Share of M2)

Country	**ER regime (code)**	**ER Vol.**	**FXI**	**FXI / M2**
Czech Republic	de facto peg (1)	0.112	0.214	0.101%
Peru	crawling peg (2)	0.113	0.548	0.486%
Hungary	crawling peg (2)	0.160	0.464	0.493%
Indonesia	crawling peg (2)	0.187	0.453	0.103%
Philippines	crawling peg (2)	0.074	0.065	0.022%
Thailand	managed floating (3)	0.053	0.669	0.097%
Malaysia	managed floating (3)	0.134	0.545	0.120%
Colombia	managed floating (3)	0.271	0.268	0.170%
Chile	managed floating (3)	0.169	0.107	0.048%
Poland	managed floating (3)	0.116	0.285	0.069%
Uruguay	managed floating (3)	0.273	0.016	0.061%
Russia	managed floating (3)	0.372	0.236	0.023%
Turkey	managed floating / free falling (3.5)	0.582	0.203	0.048%
South Africa	free floating (4)	0.286	0.194	0.072%
Group average				
	peg / crawling peg	0.130	0.334	0.241%
	managed floating	0.198	0.304	0.084%
	free floating / falling	0.434	0.198	0.060%
Whole sample average			0.302	0.137%

Note: This table reports the country-specific required size of foreign exchange intervention (FXI) to stabilize exchange rates by 1%, in billions of US dollars (column 4) and as a share (%) of each country's 2019 broad money (M2) supply (column 5). The exchange rate volatility, measured as the standard deviation of the log exchange rate level by country, is reported in column 3. We sort countries by their coarse exchange rate regimes (column 2, as classified by Iltzetki, Rogoff and Reinhart 2021) from de facto peg to free floating. For countries having multiple exchange rate regime codes during our sample period (2010–2021), as for Mexico and Turkey, we take the average regime code across time. The required size of intervention is computed using the country-specific cumulative exchange rate response to the currency demand shock on the rebalancing dates. All estimates are significant at the 1% level.

Table B.3: AUM of the GBI-EM Index in EPFR Data and its Share in ICI

(a) AUM of Funds Tracking the GBI-EM Global Diversified Index in EPFR

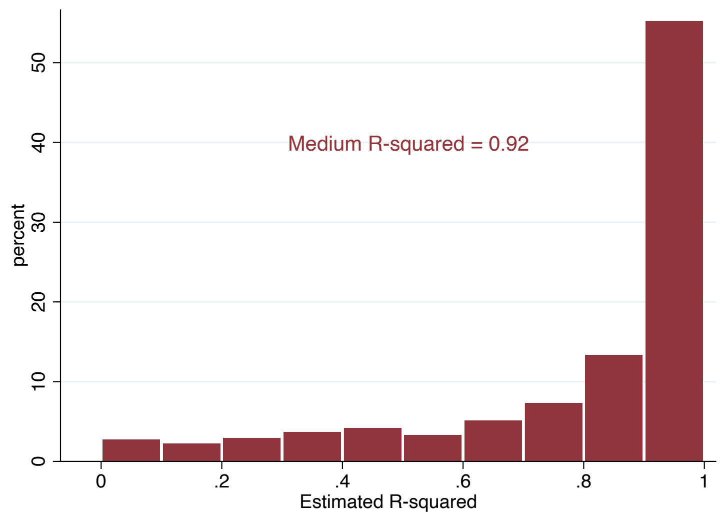


(b) EPFR Mutual Funds Population Share in ICI

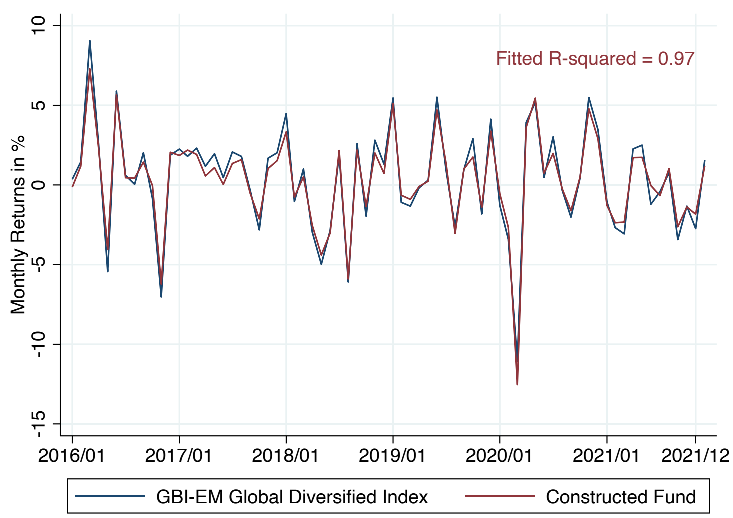
Note: This figure reports the total assets under management (AUM) of bond funds that track the GBI-EM Global Diversified index in the EPFR dataset (panel a) and the share of total EPFR data representation for the entire mutual fund industry (panel b). The bond funds aggregated in panel (a) are in billions of USD and are selected from mutual funds whose benchmark indices track the J.P. Morgan GBI-EM Global Diversified or their performance R-squared is at least 0.9. Observations are in monthly frequency from January 2016 to December 2021.

For the share of mutual fund representation in panel (b), we aggregate equity, bonds, and money market end-of month assets for both industrialized and emerging markets from the EPFR data and divide that by the global mutual fund data population from the Investment Company Fact Book. The result is the population presentation of the EPFR data in the global mutual fund industry.

Table B.4: Return Performance of Mutual Funds in the Data



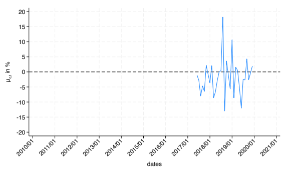
(a) Histogram of Fund Performance R^2



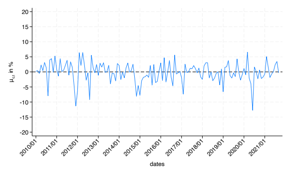
(b) Weighted (by Positions) Average Returns

Note: The left panel presents the histogram of estimated R-squared of 12-month rolling window regressions of monthly fund returns on the returns of the GBI-EM Global Diversified index; the median R-squared is 0.92. The right panel plots the returns of the GBI-EM Global Diversified index and the returns of weighted (by assets under management) mutual funds tracking the index; the performance R-squared is 0.97.

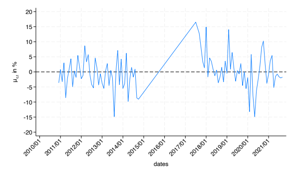
Table B.5: Currency Demand Shock ($\mu_{c,t}$ in %) in Time Series



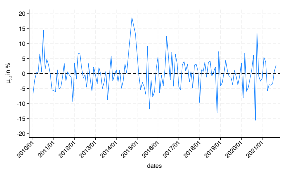
(a) Argentina



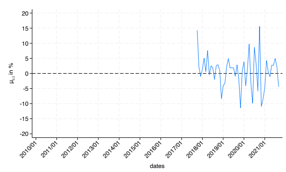
(b) Brazil



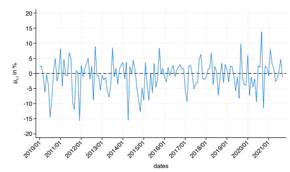
(c) Chile



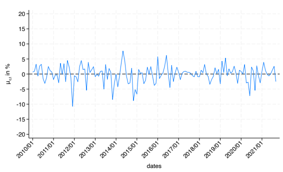
(d) Colombia



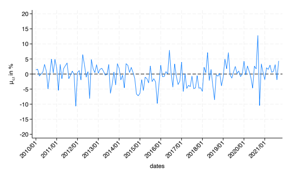
(e) Czech Republic



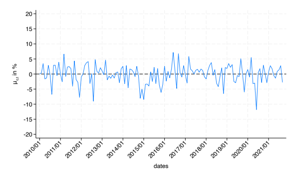
(f) Hungary



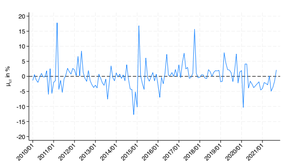
(g) Indonesia



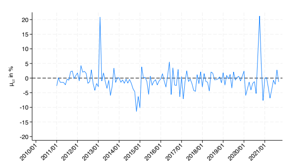
(h) Malaysia



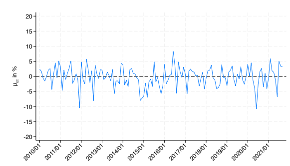
(i) Mexico



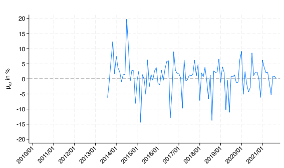
(j) Peru



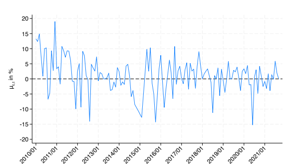
(k) Philippines



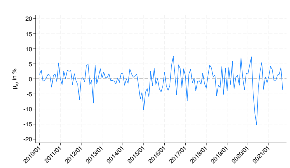
(l) Poland



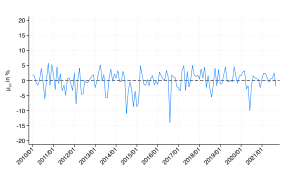
(m) Romania



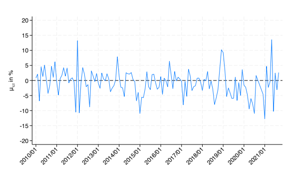
(n) Russia



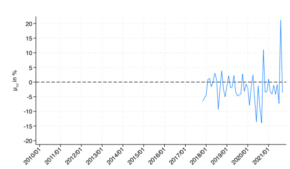
(o) South Africa



(p) Thailand



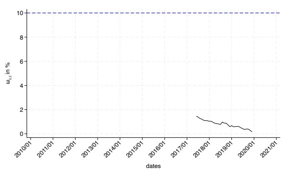
(q) Turkey



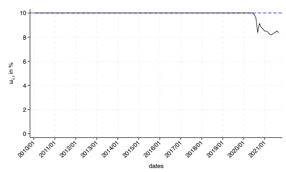
(r) Uruguay

Note: The figure depicts the monthly currency demand shock ($\mu_{c,t}$, measured in percentage points) for each country in the GBI-EM Global Diversified index between 2010 and 2021. Missing values in a given month means the country is not included in the GBI-EM Global Diversified index in that month.

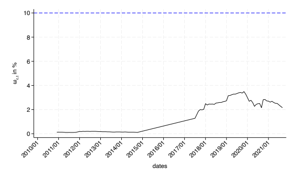
Table B.6: Weights after Rebalancing ($\omega_{c,t}$, in %) in Time Series



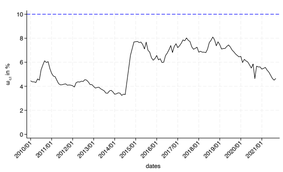
(a) Argentina



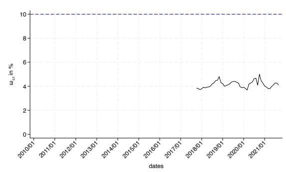
(b) Brazil



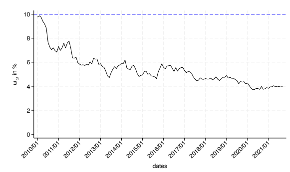
(c) Chile



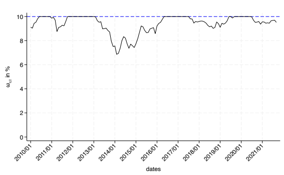
(d) Colombia



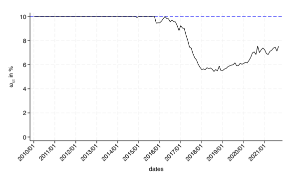
(e) Czech Republic



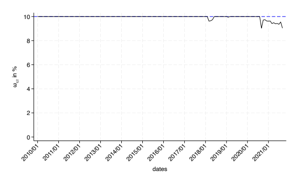
(f) Hungary



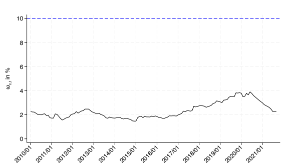
(g) Indonesia



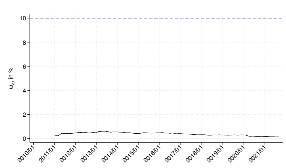
(h) Malaysia



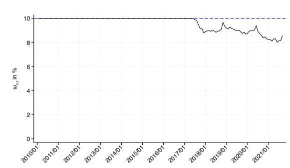
(i) Mexico



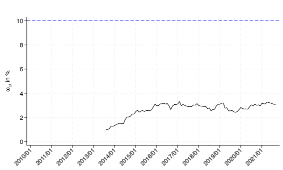
(j) Peru



(k) Philippines



(l) Poland



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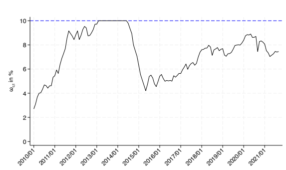
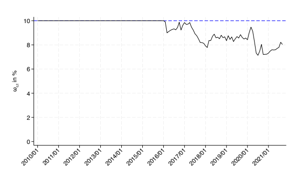
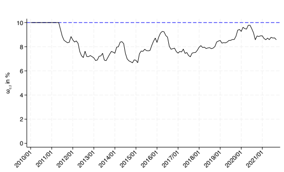


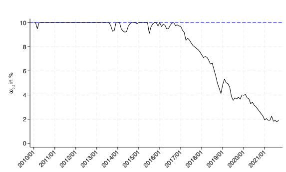
(n) Russia



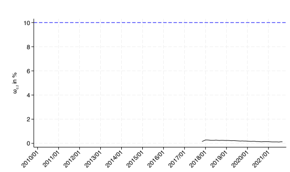
(o) South Africa



(p) Thailand



(q) Turkey



(r) Uruguay

Note: This figure shows the monthly weights after rebalancing ($\omega_{c,t}$, measured in percentage points) on rebalancing dates for each country between 2010 and 2021 (black curves). The dotted blue line indicates the 10% country weight cap that is mechanically imposed at each rebalancing date. Missing values in a given month means the country is not included in the index in that month. Note that while the weights of the country can go above 10% at any given time during the month (as the market price fluctuates), they have to be capped at 10% at the rebalancing dates (the last business day of each month).

Table B.7: Country-Specific Exchange Rate Response to $\mu_{c,t}$

	(1) Chile	(2) Colombia	(3) Czech Republic	(4) Hungary	(5) Indonesia	(6) Malaysia	(7) Peru
$\mu_{c,t}$	-1.099**	-1.935***	-1.739**	-1.903***	-2.087***	-1.416	-0.503***
	(0.367)	(0.430)	(0.385)	(0.397)	(0.424)	(0.745)	(0.123)
Const.	0.242***	0.439***	0.134*	0.162***	0.305***	0.115**	0.169***
	(0.0039)	(0.0088)	(0.0515)	(0.042)	(0.013)	(0.038)	(0.0081)
Obs.	78	136	43	131	82	65	103
R^2	0.3751	0.4887	0.6882	0.4065	0.5497	0.4371	0.3328
Adj. R^2	0.141	0.378	0.496	0.272	0.371	0.217	0.128

	(1) Philippines	(2) Poland	(3) Romania	(4) Russia	(5) South Africa	(6) Thailand	(7) Turkey	(8) Uruguay
$\mu_{c,t}$	-0.593**	-3.207**	-0.898***	-2.754***	-4.383***	-1.077**	-3.655***	-0.944**
	(0.233)	(1.052)	(0.148)	(0.540)	(0.417)	(0.402)	(0.540)	(0.293)
Const.	0.0902***	-0.0365	0.379***	0.855***	0.112***	-0.0147	-0.0596	0.726***
	(0.0225)	(0.0274)	(0.0309)	(0.171)	(0.0139)	(0.0132)	(0.150)	(0.095)
Ob.	118	46	91	111	62	116	71	33
R^2	0.275	0.611	0.418	0.596	0.809	0.338	0.686	0.580
Adj. R^2	0.097	0.396	0.252	0.490	0.729	0.181	0.561	0.253

Standard errors in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This figure shows the regression coefficient of country-level cumulative exchange rate change (in % or $100 \times \Delta \log(.)$) in response to $\mu_{c,t}$ at the rebalancing date. Exchange rate change is defined as the change from 28 days before rebalancing to the current rebalancing date.

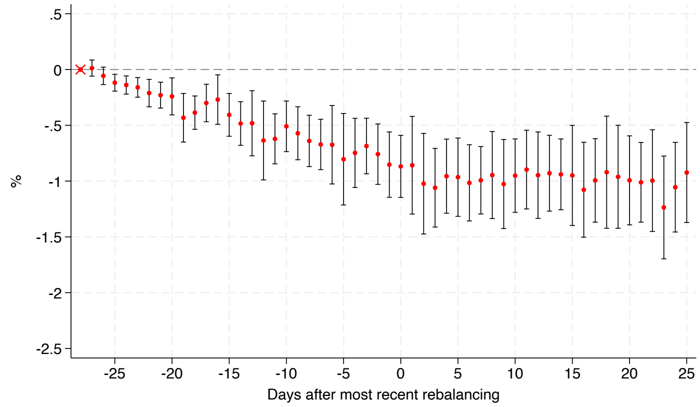
Table B.8: Autocorrelation Tests for country-specific time-series of GBI-EM Global Div. Weights ω_{ct}

	(1) Argentina	(2) Brazil	(3) Chile	(4) Colombia	(5) Czech Republic	(6) Hungary	(7) Indonesia	(8) Malaysia	(9) Mexico
Auto-corr. Coef.	0.730	0.440	0.9713	0.945	0.650	0.889	0.936	0.869	0.8720
Portmanteau test									
Test-statistics	17.10	3.53	93.42	124.09	124.09	21.59	30.86	126.28	70.95
P-value	0.00	0.06	0.00	0.00	0.00	0.00	0.00	0.00	0.00

	(1) Peru	(2) Philippines	(3) Poland	(4) Romania	(5) Russia	(6) South Africa	(7) Thailand	(8) Turkey	(9) Uruguay
Auto-corr. Coef.	0.045	0.554	0.9640	0.929	0.834	0.924	0.906	0.869	0.928
Portmanteau test									
Test-statistics	.040	10.465	131.98	111.29	37.66	86.31	101.0	53.63	109.29
P-value	0.85	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00

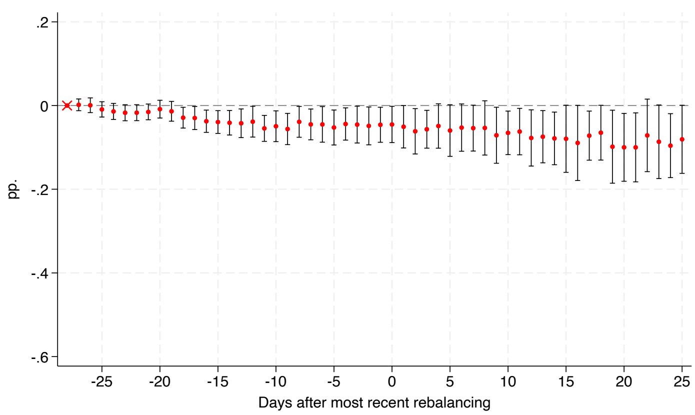
Note: This panel of figures the country-specific autocorrelation tests for the weights in GBI-EM Global Diversified index ω_{ct} after rebalancing. As in the main empirical specification, we drop those observations at the weight cap of 10% in the monthly rebalancing events. We report the estimated auto-correlation for the fitting country-specific $\mu_{c,t}$ with AR(1) and the Portmanteau white noise test on the residuals after fitting. The null hypothesis of the Portmanteau test is that the error terms are white noise. All Portmanteau white noise tests give significant coefficient except for Peru. The average auto-correlation coefficient of all countries with significant coefficients is 0.792.

Figure B.1: CUMULATIVE EXCHANGE RATE CHANGE ON $\mu_{c,t}$ (INCLUDING AT 10% CAP)



Note: This figure presents the estimated regression coefficient of the exchange rate change on the currency demand shock measured by $\mu_{c,t}$, which is standardized by its mean and standard deviation in the regression. Different from Fact 1 in the main text, this regression includes observations at the 10% threshold. Exchange rate change (local currency per USD) is defined as the cumulative change starting from 28 days before the recent rebalancing at day 0. The regression is performed in a pooled OLS using time and country fixed effects, with standard errors clustered at the date level. The results are reported in point estimates (red dots) with 95% confidence intervals (black bars).

Figure B.2: THREE-MONTH LOCAL-CURRENCY YIELDS RELATIVE TO USD YIELDS ON $\mu_{c,t}$



Note: This figure presents the estimated regression coefficient of the three-month local government bond yields relative to USD on the currency demand shock measured by $\mu_{c,t}$, which is standardized by its mean and standard deviation in the regression. The three-month yields are annualized local-currency government bond yields relative to synthetic USD $(i_{c,t} - i^*_{c,t})$ in percentage points. The regression is performed in a pooled OLS using time and country fixed effects, with standard errors clustered at the date level. The results are reported in point estimates (red dots) with 95% confidence intervals (black bars).

Table B.9: Correlation Matrix of Sovereign Bond Prices across Countries

	Argentina	Brazil	Chile	Colombia	Czech Republic	Hungary	Indonesia	Malaysia	Mexico	Peru	Phillippines	Poland	Romania	Russia	South Africa	Thailand	Turkey	Uruguay
Argentina	1	-0.8375	-0.8129	0.2588	0.5674	0.3118	0.6365	-0.6978	0.2861	-0.1942	0.1479	-0.7477	0.5894	0.1125	0.0567	-0.3082	0.6161	0.6515
Brazil	-0.8375	1	0.8259	0.136	0.0007	-0.16	0.1876	0.6409	-0.241	0.6052	0.3611	-0.0826	-0.6697	0.6915	-0.1826	0.3843	-0.2603	-0.1387
Chile	-0.8129	0.8259	1	-0.0893	-0.0991	0.2982	0.0205	0.7709	0.2172	0.5313	0.5934	0.5881	-0.0891	0.6214	-0.2868	0.6411	0.0394	-0.0867
Colombia	0.2588	0.136	-0.0893	1	0.0854	-0.6275	0.7322	0.3045	0.4828	0.6615	0.1472	-0.1345	-0.2691	0.4563	0.8199	-0.3833	0.5485	0.2267
Czech Republic	0.5674	0.0007	-0.0991	0.0854	1	0.5746	0.5731	-0.0654	0.2933	0.0762	0.1848	-0.0742	0.7641	0.0734	0.0455	0.0894	0.5688	0.3162
Hungary	0.3118	-0.16	0.2982	-0.6275	0.5746	1	-0.6249	-0.306	-0.117	-0.3587	0.2331	0.5308	0.6694	-0.4758	-0.5827	0.3045	-0.1248	-0.1467
Indonesia	0.6365	0.1876	0.0205	0.7322	0.5731	-0.6249	1	0.494	0.4165	0.605	0.1923	-0.2212	-0.058	0.4223	0.6708	-0.1097	0.4343	0.7759
Malaysia	-0.6978	0.6409	0.7709	0.3045	-0.0654	-0.306	0.494	1	0.3731	0.6203	0.5879	0.1779	-0.2167	0.6014	0.075	0.3943	0.1972	0.2707
Mexico	0.2861	-0.241	0.2172	0.4828	0.2933	-0.117	0.4165	0.3731	1	0.3078	0.5318	0.5108	0.5711	-0.09	0.5411	-0.057	0.7291	0.5352
Peru	-0.1942	0.6052	0.5313	0.6615	0.0762	-0.3587	0.605	0.6203	0.3078	1	0.5845	0.1084	-0.5196	0.6308	0.4414	-0.1035	0.3259	0.3149
Phillippines	0.1479	0.3611	0.5934	0.1472	0.1848	0.2331	0.1923	0.5879	0.5318	0.5845	1	0.6392	0.1009	0.3058	0.043	0.3197	0.4332	0.6155
Poland	-0.7477	-0.0826	0.5881	-0.1345	-0.0742	0.5308	-0.2212	0.1779	0.5108	0.1084	0.6392	1	0.4257	-0.3521	-0.1133	0.0005	0.2222	0.2845
Romania	0.5894	-0.6697	-0.0891	-0.2691	0.7641	0.6694	-0.058	-0.2167	0.5711	-0.5196	0.1009	0.4257	1	-0.7123	0.2355	0.1476	0.6135	0.7802
Russia	0.1125	0.6915	0.6214	0.4563	0.0734	-0.4758	0.4223	0.6014	-0.09	0.6308	0.3058	-0.3521	-0.7123	1	0.0026	0.1857	-0.0248	0.288
South Africa	0.0567	-0.1826	-0.2868	0.8199	0.0455	-0.5827	0.6708	0.075	0.5411	0.4414	0.043	-0.1133	0.2355	0.0026	1	-0.423	0.5578	0.0321
Thailand	-0.3082	0.3843	0.6411	-0.3833	0.0894	0.3045	-0.1097	0.3943	-0.057	-0.1035	0.3197	0.0005	0.1476	0.1857	-0.423	1	-0.0057	0.0783
Turkey	0.6161	-0.2603	0.0394	0.5485	0.5688	-0.1248	0.4343	0.1972	0.7291	0.3259	0.4332	0.2222	0.6135	-0.0248	0.5578	-0.423	1	0.2783
Uruguay	0.6515	-0.1387	-0.0867	0.2267	0.3162	-0.1467	0.7759	0.2707	0.5352	0.3149	0.6155	0.2845	0.7802	0.288	0.0321	0.0783	0.2783	1

Note: This table reports the correlation coefficient in aggregate local-currency sovereign bond prices at the rebalancing date ($P_{c,t}$ in equation (1)) across countries. Each entry in the matrix is the time-series correlation in prices between the two countries over the sample period from 2010 to 2021 at monthly frequency. The cells highlighted in red indicate positive correlations, and those highlighted in blue indicate negative correlations, with the darker shade implying a stronger correlation in magnitude. Diagonal entries are shown in the darkest shade of red because their correlation coefficients equal 1.

Table B.10: Capital Controls and Macro-fundamentals are not Responsive to $\mu_{c,t}$

	(1) Capital controls	(2) GDP	(3) Consumption	(4) NFA	(5) Net exports	(6) Inflation
$\mu_{c,t}$	-0.0375	1.27	0.430	-1.040	0.285	-1.462
	(0.0289)	(0.357)	(6.540)	(0.880)	(2.984)	(0.9602)
Constant	0.523***	1.439***	2.214***	3.947***	0.466***	4.075***
	(0.00002)	(0.112)	(0.0533)	(0.00178)	(0.00543)	(0.0022)
Observations	1793	1247	1368	1992	1835	1956
R^2	0.9781	0.9199	0.9264	0.8637	0.1875	0.6451
Adjusted R^2	0.978	0.919	0.925	0.862	0.175	0.6401

Standard errors in parentheses
* $p < 0.05$, ** $p < 0.02$, *** $p < 0.01$

Note: This table shows the OLS regression results of the following independent variables on the currency demand shock ($\mu_{c,t}$): capital control measures (Fernandez et. al., 2016), nominal GDP, consumption, net foreign asset positions (NFA), net exports, and inflation. Capital controls, GDP (billions of local currency), and inflation are in annual frequency. NFA, consumption and net exports are in billions of local currency and of quarterly frequency. Inflation level is computed from the consumer price index that treats year 2010 as the base year and is of quarterly frequency. All regressions include country and year fixed effects, with standard errors clustered at the year level.

Table B.11: Foreign Exchange Interventions are not Responsive to $\mu_{c,t}$

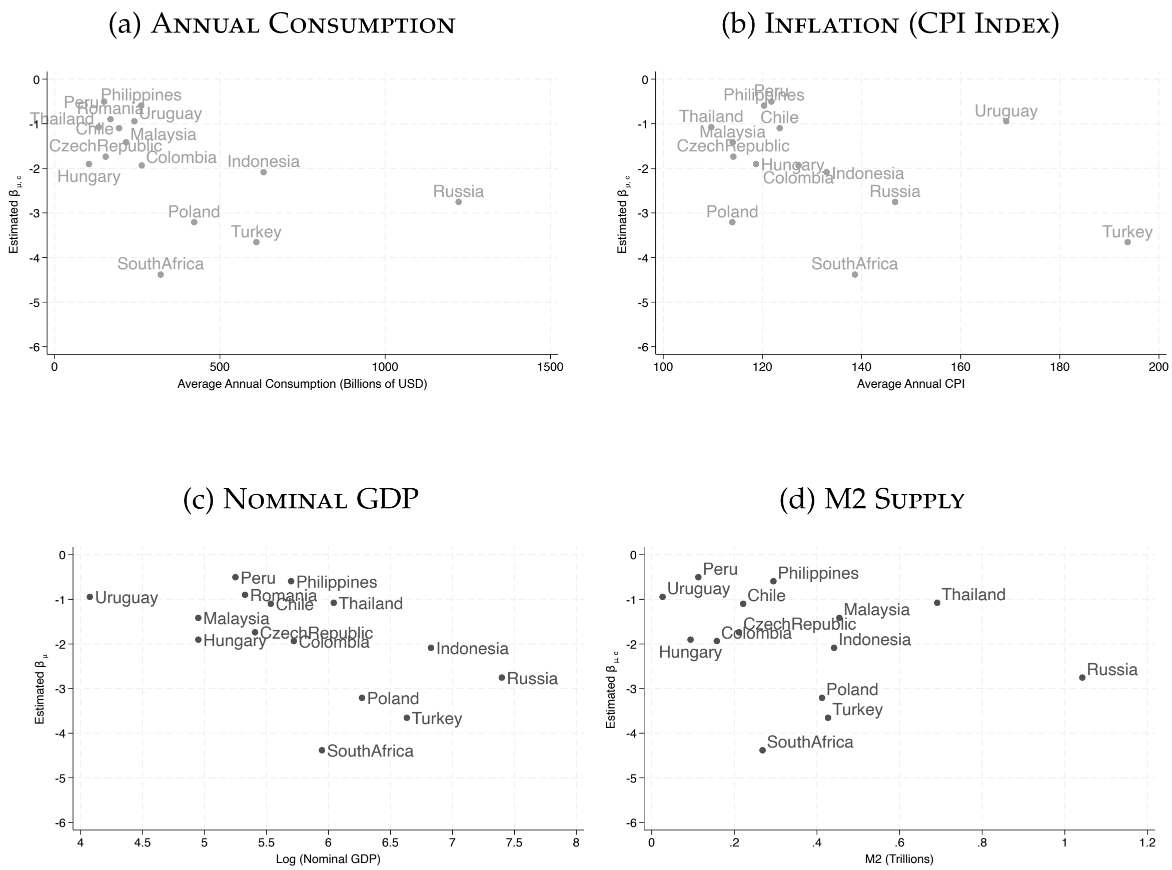
	(1) Chile	(2) Colombia	(3) Czech Republic	(4) Hungary	(5) Indonesia	(6) Malaysia	(7) Peru
$\mu_{c,t}$	0.561	-0.418	13.58	3.262	-1.456	3.00	-2.455
	(0.324)	(0.253)	(9.501)	(2.192)	(1.280)	(1.970)	(1.508)
Const.	0.044***	0.047***	0.123	0.0384	0.035***	0.006	0.096***
	(0.00008)	(0.0002)	(0.057)	(0.019)	(0.002)	(0.009)	(0.002)
Obs.	90	127	42	132	81	64	130
R^2	0.762	0.186	0.471	0.194	0.478	0.218 y	0.3816
Adj. R^2	0.693	0.013	0.167	0.032	0.292	-0.071	0.254

	(1) Philippines	(2) Poland	(3) Romania	(4) Russia	(5) South Africa	(6) Thailand	(7) Turkey	(8) Uruguay
$\mu_{c,t}$	-0.194	6.642	2.071	0.272	0.241	4.067**	0.294	3.086
	(0.462)	(8.425)	(1.479)	(0.955)	(0.394)	(1.393)	(0.651)	(1.983)
Const.	0.145***	0.100*	0.107***	0.153***	-0.016***	0.103***	-0.033**	0.126
	(0.003)	(0.028)	(0.011)	(0.010)	(0.0003)	(0.004)	(0.010)	(0.045)
Obs.	120	45	89	111	62	118	72	39
R^2	0.4952	0.3026	0.2030	0.2832	0.4942	0.3603	0.3820	0.4659
Adj. R^2	0.381	-0.058	-0.016	0.114	0.299	0.220	0.156	0.118

Standard errors in parentheses
* $p < 0.05$, ** $p < 0.02$, *** $p < 0.01$

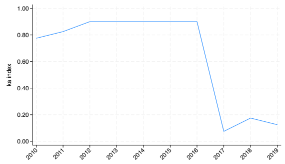
Note: This table shows the OLS regression results of the foreign exchange intervention data on the currency demand shock ($\mu_{c,t}$). Intervention data are the estimated spot foreign exchange interventions data over GDP at monthly frequency from Adler et al. (2021), which use changes in the balance sheet of central bank reserves rather than publicly available official data. We use the estimated intervention data from Adler et al. (2021) because they have a larger coverage of countries and are meant be capture the covert interventions from the central banks that are not reported otherwise. The regression includes year and month fixed effects, with standard errors clustered at the year level.

Table B.12: Exchange Rate Response and Macro/Financial Metrics

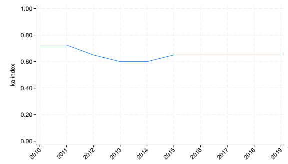


Note: Scatterplot of country-specific exchange rates response to the currency demand shock $\mu_{c,t}$ against different macroeconomic and financial metrics. Panel (a) is average annual consumption in billions of USD; (b) is average annual inflation measured by the CPI index; (c) is Nominal GDP; and (d) is M2 supply.

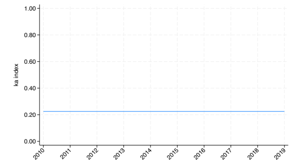
Table B.13: Capital Controls Overall Restriction Index



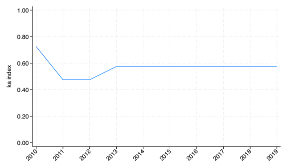
(a) Argentina



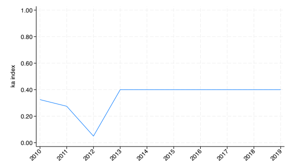
(b) Brazil



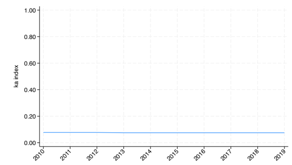
(c) Chile



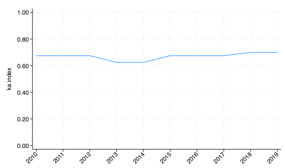
(d) Colombia



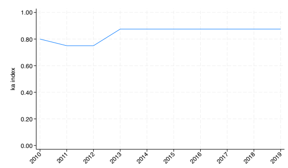
(e) Czech Republic



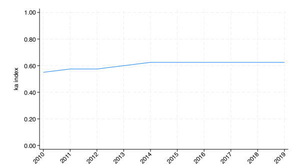
(f) Hungary



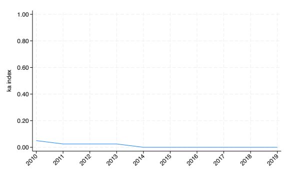
(g) Indonesia



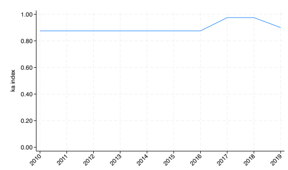
(h) Malaysia



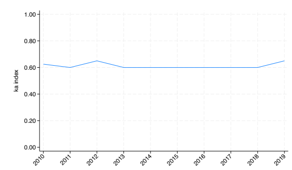
(i) Mexico



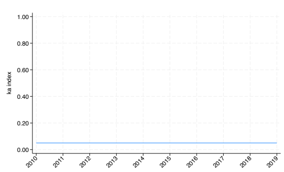
(j) Peru



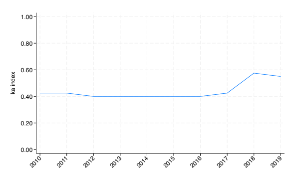
(k) Philippines



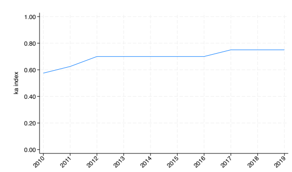
(l) Poland



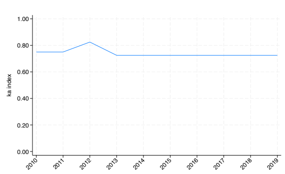
(m) Romania



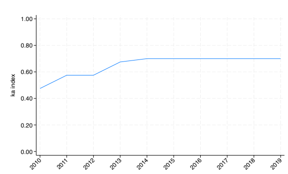
(n) Russia



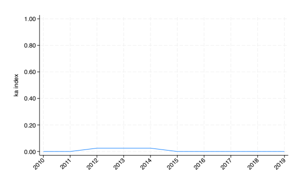
(o) South Africa



(p) Thailand



(q) Turkey



(r) Uruguay

Note: This figure presents the overall capital restriction index (the average of capital inflow and outflow restriction) for each country in our dataset, provided by Fernandez et al. (2016), with data updated to 2019. The measure is in annual frequency.

C Derivation and Proofs

C.1 Model Setup for Inelastic Financial Markets

We follow Itskhoki and Mukhin (2021, 2023a) closely and consider a small open economy, denoted by c. There are four types of agents in the partially segmented financial market where both home and foreign households can hold only government bonds of their own currency. Households demand home-currency bonds $b_{c,t}$, which are shaped by the macroeconomic fundamentals in the economy. There are also three types of agents who can trade both home- and foreign-currency bonds in the international financial market, namely, noise traders, arbitrageurs, and the government, and we assume without loss of generality that they all reside in the home country. We describe the problem of each of these agents below.

Risk-averse arbitrageurs hold a zero-capital portfolio for home- and foreign-currency bonds $(d_{c,t}, d^*_{c,t})$, with the returns on one local-currency unit holding of such portfolio given by $\tilde{i}_{c,t+1} = i_{c,t} - i^*_{c,t} - \mathbb{E}_t \Delta e_{c,t+1}$. Arbitrageurs choose $(d_{c,t}, d^*_{c,t})$ to maximize the mean-variance preferences over profits in the currency carry trade,

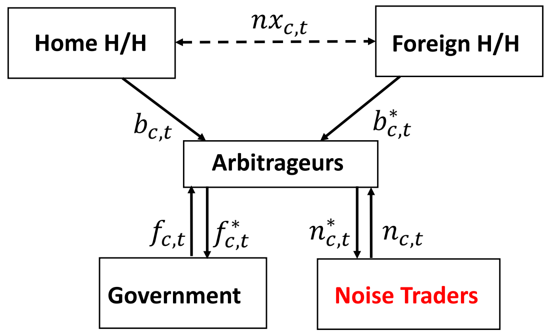
$$d_{c,t} = \frac{1}{\lambda_{c,t}} \big(i_{c,t} - i^*_{c,t} - \mathbb{E}_t \Delta e_{c,t+1} - (\tau_{c,t} + \psi_{c,t}) \big), \tag{A.5}$$

where $\lambda_{c,t} = \omega \ \sigma^2_{e_{c,t}}$ governs the arbitrageurs' risk-bearing ability; parameter ω is the arbitrageurs' risk-aversion coefficient, and $\sigma^2_{e_{c,t}}$ is the equilibrium volatility of exchange rates. The larger the $\lambda_{c,t}$ (or ω and $\sigma^2_{e_{c,t}}$), the lower the arbitrageurs' risk-bearing capacity. We model the risk-bearing capacity to be endogenously dependent on the equilibrium volatility of exchange rates, because our empirical evidence on risk-bearing capacity strongly correlates with exchange rate volatility (Fact 4).

Noise traders hold a zero-capital portfolio $(n_{c,t}, n^*_{c,t})$ and are subject to liquidity demand shocks $\mu_{c,t}$ for local-currency bonds. Importantly, $\mu_{c,t}$ is a random variable uncorrelated with the macroeconomic fundamentals. A positive $\mu_{c,t}$ means that the noise traders sell foreign-currency (US dollar) bonds and buy local-currency bonds.

The government holds a portfolio of $(f_{c,t}, f^*_{c,t})$ units of home- and foreign-currency

Figure C.1: Segmented International Bonds Market



Note: This figure presents the four types of agents in a segmented international bonds market, where home and foreign households (home H/H and foreign H/H, respectively) can hold only government bonds in their own currency. Noise traders' positions are subject to exogenous currency demand shocks that are uncorrelated with the macroeconomic fundamentals.

bonds, where $f_{c,t}$, and $f^*_{c,t}$ are policy instruments corresponding to open market operations in foreign exchange interventions for home- and foreign-currency bonds, respectively. A positive (resp., negative) $f_{c,t}$ means buying (resp., selling) local-currency bonds in the foreign exchange interventions.

We also define $b^*_{c,t}$ as the net foreign asset (NFA) position of the home households and government. In our model with only home and foreign countries, $b^*_{c,t}$ is the foreign households' holdings of foreign-currency bonds, as foreign households cannot hold home currency bonds, owing to the segmented financial market. Figure C.1 presents a diagram on the four types of agents and their positions in a segmented market.

The market clearing condition for home-currency bond states

$$b_{c,t} + n_{c,t} + d_{c,t} + f_{c,t} = 0. \tag{A.6}$$

Using the zero-capital position of the noise traders and arbitrageurs, one can arrive at the following expression for net foreign assets: $b^*_{c,t} = f^*_{c,t} + n^*_{c,t} + d^*_{c,t}$.

Combining equation (A.6) with equation (A.5) and putting exchange rates on the left-hand-side of the equation, we have

$$\mathbb{E}_t \Delta e_{c,t+1} = i_{c,t} - i^*_{c,t} - (\tau_{c,t} + \psi_{c,t}) \ + \ \lambda_{c,t} \left(b_{c,t} + n_{c,t} + f_{c,t} \right), \tag{A.7}$$

where $\lambda_{c,t} = \omega \ \sigma^2_{e_{c,t}}$, and we substitute the arbitrageurs' holdings using the market clearing condition. A currency demand shock $\mu_{c,t}$ on the local-currency bonds moves the noise traders' holdings $n_{c,t}$ and in turn the arbitrageurs' position, which then leads to movements in exchange rates and endogenous deviations in UIP. Specifically, a positive local-currency demand shock (an increase in $\mu_{c,t}$) appreciates exchange rate levels tomorrow (a decrease in $e_{c,t+1}$), with the size of the appreciation governed by the arbitrageurs' risk-bearing capacity $\lambda_{c,t} = \omega\, \sigma^2_{e_{c,t}}$.

C.2 Optimal Policies of the Central Bank

We follow Itskhoki and Mukhin (2023a) and assume the policy objective of the central bank is to maintain the trade-off between the output gap ($x_{c,t}$) stabilization and the international risk sharing wedge ($z_{c,t}$)

$$\min_{x_{c,t}, z_{c,t}, e_{c,t}, b^*_{c,t}, f^*_{c,t}, \sigma^2_{e_{c,t}}} \quad \frac{1}{2}\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \left[\gamma\, z^2_{c,t} + (1-\gamma) x^2_{c,t}\right]$$

$$\text{subject to} \quad \beta\, b^*_{c,t} = b^*_{c,t-1} - z_{c,t}$$

$$\mathbb{E}\Delta z_{c,t+1} = -\omega \sigma^2_{e_{c,t}} (b_{c,t} + n_{c,t} + f_{c,t}),$$

where the two constraints are the country budget constraint and the international risk-sharing wedge. Here $b^*_{c,t}$ is the net foreign asset position of the home country, and the international risk-sharing wedge is measured as the deviation from the uncovered interest parity (UIP) condition. Parameter γ is the weight on the international risk-sharing wedge and a measure of the degree of openness of the economy. Given initial net foreign assets $b^*_{c,-1}$ and the exogenous path of noise trader shocks $n_{c,t}$, monetary policy chooses the direct path of the output gap $x_{c,t}$, while foreign exchange intervention $f_{c,t}$ chooses the path of the risk-sharing wedge $z_{c,t}$. The goal of the policymaker is thus to minimize the weighted average of the volatility of the output gap $x_{c,t}$ and the risk-sharing wedge $z_{c,t}$.

If both policy instruments are available and unconstrained, the optimal policy fully stabilizes both wedges, the output gap $x_{c,t} = 0$ and the risk sharing wedge $z_{c,t} = 0$. The constrained optimum allocation features $x_{c,t} = z_{c,t} = 0$ for all $z_{c,t}$. Such allocation can be delivered by a combination of monetary policy and foreign exchange (FX) policies, with monetary policy stabilizing the output gap ($x_{c,t} = 0$), and optimal FX interventions $\partial f_{c,t}/\partial e_{c,t} = -\partial n_{c,t}/\partial e_{c,t}$ to ensure $z_{c,t} = 0$. As a result, the risk-sharing wedge is fully offset, and the optimal international risk sharing is restored independently of the currency demand shocks $n_{c,t}$.

Optimal policies above can be implemented using a conventional Taylor interest rule targeting the output gap and a similar policy rule for FX interventions that target UIP deviations. Specifically, FX interventions $f_{c,t} = -\mathbb{E}_t \Delta z_{c,t+1}$ and the optimal FX interventions should "lean against the wind" intensively until the UIP wedge is fully eliminated. The implementation does not require observing the shocks and distinguishing between macro-fundamental and non-fundamental sources of variation in exchange rates.

C.3 Proof of Examples 1 and 2

In this section, we provide two model examples and their solutions of exchange rates in response to the noise trader shocks (or currency demand shocks). These impulse responses of exchange rates can be directly mapped into the estimated coefficient from our empirical results. We start with a model with endogenous deviation from uncovered interest parity (UIP) with inelastic financial markets. We then combine the UIP equation with both a partial equilibrium model (Engel and West, 2005) and a general equilibrium model (Itskhoki and Mukhin, 2021) to solve for exchange rates and their impulse response functions to the noise trader shocks.

The modified UIP equation with endogenous UIP deviations as in equation (1) is

$$i_{c,t} - i^*_{c,t} - \mathbb{E}_t \Delta e_{c,t+1} = \tau_{c,t} + \psi_{c,t} - \omega \, \sigma^2_{e_{c,t}} \big(b_{c,t} + n_{c,t} + f_{c,t} \big), \tag{A.8}$$

where we have substituted the risk-bearing capacity $\lambda_{c,t} = \omega \, \sigma^2_{e_{c,t}}$. Capital control taxes ($\tau_{c,t}$) and risk-premium shock ($\psi_{c,t}$) impose exogenous UIP deviations. We can rearrange equation (A.8) as

$$\mathbb{E}_t \Delta e_{c,t+1} = \underbrace{(i_{c,t} - i^*_{c,t}) - \tau_{c,t} - \psi_{c,t}}_{\equiv -x_{c,t}} \quad + \quad \underbrace{\omega\, \sigma^2_{e_{c,t}} (b_{c,t} + n_{c,t} + f_{c,t})}_{\equiv -u_{c,t}}, \tag{A.9}$$

where $x_{c,t}$ is the component of exchange rate $e_{c,t}$ where the classical trilemma constraint holds. The term $u_{c,t}$ is the additional component that generates endogenous UIP deviation when the classical trilemma constraint no longer binds, as FX interventions $f_{c,t}$ can now work as an additional policy tool to stabilize exchange rates under inelastic financial markets. Specifically, under trilemma models where the classical trilemma constraint holds, the risk-bearing capacity of the arbitrageurs $\lambda_{c,t} = 0$, due to either the risk-aversion of the arbitrageurs $\omega = 0$ or to exchange rates being fixed (so that $\sigma^2_{e_c} = 0$). The term non-trilemma $u_{c,t}$ therefore vanishes under trilemma models, whose UIP deviations can come only from exogenous UIP shocks.

If we continue to iterate equation (A.9) forward, we have

$$e_{c,t} = \mathbb{E}_t\, e_{c,\infty} + \mathbb{E}_t \sum_{j=0}^{\infty} x_{c,t+j} + \mathbb{E}_t \sum_{j=0}^{\infty} u_{c,t+j}, \tag{A.10}$$

and the expectation term vanishes as $e_{c,\infty} = 0$ if exchange rate $e_{c,t}$ follows a stationary process.

Example 1. *In the Taylor-rule model (Engel and West, 2005) with exchange rate target $\bar{e}_c$, the home- and foreign monetary policy rates follow the form*

$$\begin{aligned} i_{c,t} &= \beta_0\, (e_{c,t} - \bar{e}_c) + \beta_1\, y_{c,t} + \beta_2\, \pi_{c,t} + \nu_{c,t} \quad, \beta_0 \in (0,1) \\ i^*_{c,t} &= \beta_1\, y^*_{c,t} + \beta_2\, y^*_{c,t} + \nu^*_{c,t}, \end{aligned}$$

where $\bar{e}_c$ is the exchange rate target, $\pi_{c,t} = p_{c,t} - p_{c,t-1}$ is the inflation rate, and $y_{c,t}$ is the output gap of home country c. The policy function of the foreign exchange intervention is given by

$$\frac{\partial e_{c,t}}{\partial f_{c,t}} = -\frac{\partial e_{c,t}}{\partial n_{c,t}} = \frac{1}{(1+\beta_0 - \rho)} \lambda_{c,t},$$

where $\lambda_{c,t} = \omega\, \sigma^2_{e_{c,t}}$, under the assumptions that $n_{c,t} \sim AR(1)$ with persistence ρ, $n_{c,t} \perp f_{c,t}$, *and macro-fundamentals are slow-moving compared with the noise trader shocks.*

We solve for the impulse exchange rate response to the noise trader shocks under a partial equilibrium model with Taylor rule as specified by Engel and West (2005). Let $\pi_{c,t} = p_{c,t} - p_{c,t-1}$ be the inflation rate, and $y_{c,t}$, the output gap of home country c. Monetary policy in the home country (emerging country) follows a Taylor rule of the form

$$i_{c,t} = \beta_0(e_{c,t} - \bar{e}_{c,t}) + \beta_1\, y_{c,t} + \beta_2\, \pi_{c,t} + v_{c,t},$$

where exchange rate target $\bar{e}_t$ ensures purchasing power parity (PPP) so that $\bar{e}_{c,t} = p_{c,t} - p^*_{c,t}$ and $\beta_0 \in (0,1)$.

Monetary policy in the foreign country (U.S.) follows the Taylor rule of the form

$$i^*_{c,t} = \beta_1\, y^*_{c,t} + \beta_2 \pi^*_{c,t} + v^*_{c,t}$$

Interest rate difference $i_{c,t} - i^*_{c,t}$ can thus be written as

$$i_{c,t} - i^*_{c,t} = \beta_0\ (e_{c,t} - \bar{e}_{c,t}) + \beta_1(y_{c,t} - y^*_{c,t}) + \beta_2(\pi_{c,t} - \pi^*_{c,t}) + (v_{c,t} - v^*_{c,t}). \tag{A.11}$$

Now combine the interest rate differential expression in equation (A.11) with the UIP condition in equation (A.9) to substitute out $(i_{c,t} - i^*_{c,t})$

$$\mathbb{E}_t\, e_{c,t+1} = e_{c,t} - \tau_{c,t} - \psi_{c,t} + \beta_0\ (e_{c,t} - \bar{e}_{c,t}) + \beta_1(y_{c,t} - y^*_{c,t}) + \beta_2(\pi_{c,t} - \pi^*_{c,t}) + (v_{c,t} - v^*_{c,t}) - u_{c,t},$$

and it can be further simplified to

$$\begin{aligned}(1+\beta_0)e_{c,t} &= \tau_{c,t} + \psi_{c,t} + \mathbb{E}_t e_{c,t+1} + \beta_0(p_{c,t} - p^*_{c,t}) - \beta_1(y_{c,t} - y^*_{c,t}) - \beta_2(\pi_{c,t} - \pi^*_{c,t}) - (v_{c,t} - v^*_{c,t}) + u_{c,t} \\ \Rightarrow e_{c,t} &= \frac{1}{1+\beta_0}(\tau_{c,t} + \psi_{c,t}) + \frac{\beta_0}{1+\beta_0}(p_{c,t} - p^*_{c,t}) - \frac{\beta_1}{1+\beta_0}(y_{c,t} - y^*_{c,t}) - \frac{\beta_2}{1+\beta_0}(\pi_{c,t} - \pi^*_{c,t}) - \cdots \\ &\quad - \frac{1}{1+\beta_0}(v_{c,t} - v^*_{c,t}) + \frac{1}{1+\beta_0}u_{c,t} + \frac{1}{1+\beta_0}\mathbb{E}_t e_{c,t+1}.\end{aligned}$$

We can write the solution of exchange rate under the Taylor rule as in equation (A.9) by separating its trilemma and non-trilemma components

$$e_{c,t} = X_{c,t} + U_{c,t} + \frac{1}{1+\beta_0}\mathbb{E}_t\, e_{c,t+1}, \tag{A.12}$$

where $\beta_0 \in (0,1)$, $U_{c,t} \equiv \frac{1}{1+\beta_0}u_{c,t} = -\frac{1}{1+\beta_0}\,\omega\sigma^2_{e_{c,t}}(b_{c,t} + n_{c,t} + f_{c,t})$ is the non-trilemma component and $X_t \equiv \frac{1}{1+\beta_0}(\tau_{c,t} + \psi_{c,t}) + \frac{\beta_0}{1+\beta_0}(p_{c,t} - p^*_{c,t}) - \frac{\beta_1}{1+\beta_0}(y_{c,t} - y^*_{c,t}) - \frac{\beta_2}{1+\beta_0}(\pi_{c,t} - \pi^*_{c,t}) - \frac{1}{1+\beta_0}(v_{c,t} - v^*_{c,t})$ is the trilemma component.

Iterating equation (A.12) forward, we have

$$e_{c,t} = \mathbb{E}_t \sum_{j=0}^{\infty}\frac{1}{(1+\beta_0)^j}X_{c,t+j} \;+\; \mathbb{E}_t \sum_{j=0}^{\infty}\frac{1}{(1+\beta_0)^j}U_{c,t+j} \;+\; \mathbb{E}_t \lim_{j\to\infty}\frac{1}{(1+\beta_0)^j}e_{c,\infty},$$

where $\lim_{j\to\infty}\frac{1}{(1+\beta_0)^j} = 0$ in the limit, together with $e_{c,\infty} = 0$ under the stationary process, the expectation term of exchange rates vanishes.

If we impose the assumption that $n_{c,t}$ inside the non-trilemma component $U_{c,t}$ is an AR(1) process with persistence ρ, foreign exchange interventions are independent of noise trader shocks $f_{c,t} \perp n_{c,t}$, and that macro-fundamentals are slow-moving compared with noise trader shocks $n_{c,t}$, we can solve for the impulse response of exchange rate $e_{c,t}$ in response to $n_{c,t}$ as

$$\frac{\partial e_{c,t}}{\partial n_{c,t}} = \frac{-\omega\sigma^2_{e_{c,t}}}{(1+\beta_0-\rho)} < 0. \tag{A.13}$$

On impact, a positive noise trader shock (or a positive local-currency demand shock from the increase in country weight in the GBI-EM Global Diversified index) appreciates home currency and leads to a decrease in exchange rate $e_{c,t}$, which is defined in the number of local currencies per USD. Therefore, the model prediction gives the right sign as suggested by our empirical evidence.

Example 2. *In the general equilibrium model of Itskhoki and Mukhin (2021) that specifies the budget constraint of a country c,* $\beta\; b^*_{c,t} - b^*_{c,t-1} = nx_{c,t} = \kappa\; e_{c,t} + \xi_{c,t}$, *where* $nx_{c,t}$ *is the net exports and* $b^*_{c,t}$ *the net foreign assets of the home country;* β *the discount factor in the household preference and* κ *the openness of the economy. The policy function of the foreign* *exchange intervention is given by*

$$\frac{\partial e_{c,t}}{\partial f_{c,t}} = -\frac{\partial e_{c,t}}{\partial n_{c,t}} = \frac{\beta}{(1-\rho\beta)}\lambda_{c,t},$$

where $\lambda_{c,t} = \omega\ \sigma^2_{e_{c,t}}$, *under the assumptions that* $n_{c,t} \sim AR(1)$ *with persistence* ρ, $n_{c,t} \perp f_{c,t}$, *and macro-fundamentals are slow-moving compared with noise trader shocks.*

We now solve for the impulse exchange rate response to the noise trader shocks under a general equilibrium model with the country's intertemporal budget constraint as specified by Itskhoki and Mukhin (2021). The log-linearized intertemporal budget constraint states

$$\beta b^*_{c,t} - b^*_{c,t-1} = nx_{c,t} = \kappa\ e_{c,t} + \xi_{c,t}, \tag{A.14}$$

where $b^*_{c,t}$ is the net foreign asset position of country c at time t; $nx_{c,t}$ is the net exports; and $e_{c,t}$ is the level of exchange rates. Parameter β is the discount factor; κ (> 0) is a structural parameter pinned down from the price equations in the equilibrium goods market; and $\xi_{c,t}$ is a shock to the net exports $nx_{c,t}$ and is orthogonal to $e_{c,t}$.

We iterate the country budget constraint forward and get

$$b^*_{c,t-1} + \mathbb{E}_t\ \kappa \sum_{j=0}^{\infty} \beta^j e_{c,t+j} = \lim_{T\to\infty} \beta^T b^*_{c,t+T-1} = 0, \tag{A.15}$$

where we impose the no-Ponzi game condition (NPGC) on the country's intertemporal budget constraint.

The country's intertemporal budget constraint uses the net foreign asset position $b^*_{c,t}$ of home households (which equals foreign households' holding of foreign-currency bonds), while the UIP condition in equation (A.9) uses home households' holding of home-currency bonds. The modified UIP condition is still

$$\mathbb{E}_t \Delta e_{c,t+1} = \underbrace{(i_{c,t} - i^*_{c,t}) - \tau_{c,t} - \psi_{c,t}}_{\equiv -x_{c,t}} \quad + \quad \underbrace{\omega\ \sigma^2_{e_{c,t}}\left(\iota\ b_{c,t} + n_{c,t} + f_{c,t}\right)}_{\equiv -u_{c,t}}, \tag{A.16}$$

where we use the market clearing condition of home- and foreign-currency bonds to substitute the zero-capital position of the arbitrageurs' holdings. In addition, we follow

Itskhoki and Mukhin (2023a) and assume $\iota = 0$ in the UIP condition[30] as net foreign assets are slow-moving macroeconomic fundamentals compared to exchange rates and noise trader shocks.

We iterate equation (A.16) forward, as we did for equation (A.9), to derive an expression of $\mathbb{E}_t\, e_{c,t+j}$

$$\mathbb{E}_t\, e_{c,t+j} = \mathbb{E}_t\, e_{c,\infty} + \mathbb{E}_t \sum_{k=0}^{\infty} x_{c,t+j+k} + \mathbb{E}_t \sum_{k=0}^{\infty} u_{c,t+j+k}. \tag{A.17}$$

We then combine equation (A.17) with the country budget constraint in (A.15)

$$\begin{aligned}
& b^*_{c,t-1} + \kappa \sum_{j=0}^{\infty} \beta^j\, \mathbb{E}_t e_{c,t+j} = 0 \\
\Rightarrow & b^*_{c,t-1} + \kappa \sum_{k=0}^{\infty} \beta^j \Big(\mathbb{E}_t\, e_{c,\infty} + \mathbb{E}_t \sum_{k=0}^{\infty} x_{c,t+j+k} + \mathbb{E}_t \sum_{j=0}^{\infty} u_{c,t+j+k} \Big) = 0 \\
\Rightarrow & b^*_{c,t-1} + \frac{\kappa}{1-\beta} \mathbb{E}_t\, e_{c,\infty} + \kappa\, \mathbb{E}_t \sum_j \sum_k \beta^j x_{c,t+j+k} + \kappa\, \mathbb{E}_t \sum_j \sum_k \beta^j u_{c,t+j+k} = 0 \\
\Rightarrow & b^*_{c,t-1} + \frac{\kappa}{1-\beta} \underbrace{\Big(e_{c,t} - \mathbb{E}_t \sum_{j=0}^{\infty} x_{c,t+j} - \mathbb{E}_t \sum_{j=0}^{\infty} u_{c,t+j} \Big)}_{=\mathbb{E}_t\, e_{c,\infty}} + \kappa\, \mathbb{E}_t \sum_j \sum_k \beta^j x_{c,t+j+k} + \kappa\, \mathbb{E}_t \sum_j \sum_k \beta^j u_{c,t+j+k} = 0,
\end{aligned}$$

where the last line substitutes the expression of $\mathbb{E}_t\, e_{c,\infty}$ from equation (A.17).

From above, we have the relation between $e_{c,t}$ and $b^*_{c,t-1}$

$$\frac{\kappa}{1-\beta} e_{c,t} + b^*_{c,t-1} + X_{c,t} + U_{c,t} = 0, \tag{A.18}$$

where $X_{c,t} \equiv -\frac{\kappa}{1-\beta} \sum_j\ \mathbb{E}_t x_{c,t+j} + \kappa \sum_j \sum_k \beta^j\ \mathbb{E}_t x_{c,t+j+k}$ is the trilemma component of the UIP equation, and $U_{c,t} \equiv -\frac{\kappa}{1-\beta} \sum_j\ \mathbb{E}_t u_{c,t+j} + \kappa \sum_j \sum_k \beta^j\ \mathbb{E}_t u_{c,t+j+k}$ is the non-trilemma component of the UIP equation and generates endogenous UIP deviations from the noise trader shocks.

To arrive at the closed-form solution of the exchange rate response to the noise trader

[30] Parameter ι comes from log-linearization of the UIP condition and ι = 0 or 1 depending on the point of linearization. We follow Itskhoki and Mukhin (2023a) and assume $\iota = 0$, which approximates the situation when macroeconomic demand for government bonds is orders of magnitude smaller than financial (liquidity) demand from the noise traders, and therefore $b_{c,t}$ disappears in relative terms in the limit.

shock, we impose the following assumptions, as in the model example 1 under Taylor rule: We assume that noise traders' positions $n_{c,t}$ inside the non-trilemma component follows an AR(1) process with persistence ρ, that foreign exchange interventions $f_{c,t}$ are independent of noise trader shocks ($f_{c,t} \perp n_{c,t}$), and that macro-fundamentals are slow-moving compared with noise trader shocks $n_{c,t}$.

Under these three assumptions, we can simplify equation (A.18) to

$$\frac{\kappa}{1-\beta} e_{c,t} + b^*_{c,t-1} + X_{c,t} + \frac{\beta\kappa \; \omega\sigma^2_{e_{c,t}}}{(1-\rho\beta)(1-\beta)} \; n_{c,t} + \tilde{U}_{c,t} = 0, \tag{A.19}$$

where $\tilde{U}_{c,t} \equiv U_{c,t} + \frac{\beta\kappa \; \omega\sigma^2_{e_{c,t}}}{(1-\rho\beta)(1-\beta)} \; n_{c,t}$ are the residuals of the non-trilemma component of $U_{c,t}$, such as foreign exchange interventions $f_{c,t}$, that are independent of noise traders' positions $n_{c,t}$. We can therefore compute the impulse response of exchange rate level e_t to the noise trader shock as

$$\frac{\partial e_{c,t}}{\partial n_{c,t}} = \frac{-\beta \; \omega\sigma^2_{e_{c,t}}}{(1-\rho\beta)} < 0, \tag{A.20}$$

where $\rho, \beta \in (0,1)$. A positive local-currency demand shock appreciates local currency and results in a decrease in local-currency exchange rate level $e_{c,t}$, which is measured in number of local currencies per dollar. This consistent with the observed empirical evidence on the currency demand shock.