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Keywords

exchange rates, U.S. monetary policy, portfolio adjustment frictions, forward guidance, event study

JEL Classification

E52, F31, F41, G11, G17

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US Monetary Policy, Exchange Rates, and Delayed Portfolio Adjustments[*]

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May 29, 2026

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1 Introduction

Spillovers from U.S. monetary policy to the rest of the world have long attracted the attention of academics and policymakers, reflecting the central role of the United States in the global financial system. Exchange rates, in particular, play a key role in the external adjustment of open economies and have therefore been extensively studied in response to U.S. monetary policy shocks (Albrizio, Choi, Furceri and Yoon, 2020; Bhattarai, Chatterjee and Park, 2021; Dedola, Rivolta and Stracca, 2017; Georgiadis, 2016; Ha, 2021; Hoek, Kamin and Yoldas, 2022; Iacoviello and Navarro, 2019; Maćkowiak, 2007).

Many existing studies examine exchange-rate responses at monthly to quarterly frequencies, motivated by interest in macroeconomic spillovers. However, inference at lower frequencies is complicated by the difficulty of separating policy shocks from systematic policy responses (see, among others, Faust, Rogers, Swanson and Wright, 2003; Kim, Moon and Velasco, 2017; Kim and Roubini, 2000; Scholl and Uhlig, 2008). Studying exchange rate movements at high frequency provides a complementary perspective. High-frequency responses around policy announcements are less contaminated by endogeneity concerns, allow sharper tests of asset-pricing implications such as uncovered interest parity (UIP), and shed light on market efficiency and international capital flow dynamics.

In this paper, we use high-frequency variation around Federal Open Market Committee (FOMC) announcements to study why short-run exchange-rate responses to U.S. monetary policy differ systematically across countries. We focus on daily exchange rate responses of many advanced and emerging market economies and document systematic heterogeneity across countries. Concentrating on short-run dynamics mitigates concerns that observed differences reflect endogenous reactions of domestic central banks rather than direct spillovers from U.S. policy (Choi, Willems and Yoo, 2024; Hnatkovska, Lahiri and Vegh, 2016; Taylor, 2001). Our analysis therefore relates to a growing literature that exploits high-frequency identification to study exchange rate responses to monetary policy shocks (Aizenman, Binici and Hutchison, 2016; Albagli, Ceballos, Claro and Romero, 2024; Georgiadis and Jarociński, 2025; Gürkaynak, Kara, Kısacıkoğlu and Lee, 2021; Inoue and Rossi, 2019; Miranda-Agrippino and Nenova, 2022; Rosa, 2011).

Our main contribution is to show that financial market depth is a key determinant of the magnitude of exchange rate responses to U.S. monetary policy shocks in a wide sample of countries. Empirically, we document that economies with deeper financial markets—proxied by the size of foreign portfolio liabilities, gross foreign portfolio positions, stock market turnover, or broad measures of financial development—experience larger short-run currency depreciations following contractionary U.S. monetary policy.

Following Swanson (2021), we distinguish between federal funds rate shocks, which capture target-rate surprises, and forward-guidance shocks, which capture path surprises. We first show that exchange rates respond more strongly to forward-guidance shocks, consistent with their forward-looking nature and sensitivity to expected future interest rate differentials (Engel, 2016; Galí, 2020). More importantly, the amplifying role of financial market depth in shaping exchange rate depreciations is substantially more pronounced for forward guidance shocks.

We show that this depth amplification is robust across a broad set of alternative specifications. The results are not driven by our baseline decomposition of FOMC announcement surprises into federal-funds-rate, forward-guidance, and asset-purchase components: they continue to hold when we instead reconstruct the original target and path factors from short-rate futures surprises in the spirit of Gürkaynak, Sack and Swanson (2005). The results are also robust to controlling for broad dollar movements, global risk aversion, exchange-rate volatility, institutional quality, exposure to U.S. long-term interest rates, pre-FOMC macroeconomic and financial conditions, domestic monetary policy responses, and foreign exchange market intervention. We further distinguish the market-depth channel from external balance-sheet effects by separating gross foreign portfolio positions from net foreign asset positions. The evidence suggests that the two channels are not mutually exclusive: deeper financial markets amplify short-run depreciation pressures, while stronger net external positions dampen them.

At first glance, our findings may appear at odds with the conventional view that emerging market economies with less developed financial markets are more vulnerable to U.S. monetary policy spillovers (Aizenman, Park, Qureshi, Saadaoui and Uddin, 2024; Albrizio et al., 2020; Bräuning and Ivashina, 2020; Georgiadis, 2016; Iacoviello and Navarro, 2019; Maćkowiak, 2007). We reconcile this tension by emphasizing the role of frequency. Most existing studies examine spillovers using monthly or quarterly data, whereas our analysis isolates daily exchange-rate movements around FOMC announcements. We argue that these short-run responses are shaped more directly by portfolio rebalancing frictions, while responses at monthly or quarterly horizons are more likely to reflect broader macroeconomic fundamentals and policy adjustments.

To interpret these empirical findings, we use a parsimonious theoretical framework that highlights how financial market depth can amplify short-run exchange-rate responses to U.S. monetary policy shocks. Building on the gradual portfolio-adjustment model of Bacchetta and Van Wincoop (2021), we allow effective portfolio adjustment costs to decline with market depth, consistent with standard insights from market microstructure linking transaction costs to market liquidity (Demsetz, 1968; Stoll, 2000). In this setting, the equilibrium exchange rate depends on the present discounted value of current and expected future interest-rate differentials, and lower adjustment costs in deeper markets raise the weight placed on the expected future path of returns. The model is intended as a simple organizing framework rather than a fully structural quantitative account of the empirical magnitudes.

A first implication is that contractionary U.S. monetary policy shocks, interpreted as foreign monetary policy shocks from the perspective of a small open economy, lead to a depreciation of the domestic currency, and that this depreciation is larger when financial markets are deeper. The key mechanism is that lower adjustment costs weaken effective discounting in exchange-rate determination, placing greater weight on expected future interest-rate differentials.

A second, and central, implication concerns the *type* of shock. We model forward guidance as a news shock to the expected future path of interest rates, distinct from a spot shock to the current policy rate. Both shocks raise real interest-rate differentials, but they load differently on the discounted path that determines the exchange rate: a spot shock concentrates its mass on the contemporaneous differential, whereas a news shock shifts the expected future path. Because deeper markets place greater weight on that future path, a given increase in market depth amplifies news shocks more than spot shocks. This shock-dependent amplification arises from the interaction between lower effective discounting and the different timing of spot and news shocks. This is the analytical step that is absent from Bacchetta and Van Wincoop (2021) and allows the model to interpret the forward-guidance-specific heterogeneity we document.

These predictions provide a simple explanation for why short-run exchange-rate responses depend both on the nature of U.S. monetary policy shocks and on cross-country differences in financial market depth. The remainder of the paper is organized as follows. Section 2 presents our empirical strategy, results, and robustness checks. Section 3 develops the theoretical framework and maps its predictions to the data. Section 4 concludes.

2 Empirical Analysis

2.1 Empirical model and data

We estimate the short-run effects of U.S. monetary policy shocks on bilateral exchange rates using a high-frequency event-study design. Specifically, we examine movements in exchange rates against the U.S. dollar around Federal Open Market Committee (FOMC) announcement dates. To study cross-country heterogeneity, we allow exchange-rate responses to vary with each country's financial market depth. Our benchmark specification is

$$\begin{aligned}\Delta E_{i,t}^{h} &= \beta_0 + \beta_1 MPS_t^{orth} + \gamma_1 MPS_t^{orth} \times depth_{i,y-1} + \alpha_1 depth_{i,y-1} \\ &\quad + \delta_1 size_{i,y-1} + \delta_2 GDPPC_{i,y-1} + \delta_3 \Delta RGDP_{i,y} + \delta_4 openness_{i,y} + \delta_5 D.regime_{i,ym} \\ &\quad + \eta_1 \Delta VIX_t^{h} + D.year + D.country + \epsilon_{i,t}^{h},\end{aligned} \tag{1}$$

where $\Delta E_{i,t}^{h}$ denotes the percentage change in the nominal exchange rate of country i against the U.S. dollar from one day before the FOMC announcement to h days after:

$$\Delta E_{i,t}^{h} = \frac{e_{i,t+h} - e_{i,t-1}}{e_{i,t-1}} \times 100,$$

where $e_{i,t}$ is the local-currency price of one U.S. dollar. An increase in $e_{i,t}$ therefore corresponds to a depreciation of currency i. Daily exchange-rate data are from the Bank for International Settlements (BIS).

To isolate exogenous monetary policy shocks, we use the orthogonalized monetary policy surprise series, MPS_t^{orth}, constructed by Bauer and Swanson (2023). These shocks are purged of macroeconomic fundamentals and financial conditions that may independently affect exchange rates, mitigating concerns that policy actions reflect endogenous responses to contemporaneous news. We interact monetary policy shocks with measures of financial market depth, $depth_{i,y-1}$, to test the paper's central hypothesis. The coefficient of interest is therefore γ_1. To facilitate interpretation of the interaction terms, each market-depth measure is demeaned before estimation.

In the benchmark specification, we use lagged log foreign liabilities, $liab_{i,y-1}$, as the proxy for financial market depth. To mitigate reverse causality, foreign liabilities are measured at the end of the previous year, reflecting the annual availability of cross-country liability data. The data are from the IMF's *International Financial Statistics*, converted to real U.S. dollars using the U.S. CPI, and expressed in logarithms. The baseline measure is not normalized by GDP because the mechanism in the paper concerns the absolute size of marketable assets available for international portfolio rebalancing. Instead, all specifications control for the logarithm of real GDP, $size_{i,y-1}$, to absorb differences in economic scale.

We control for additional country characteristics that may systematically affect exchange-rate dynamics. These include the logarithm of real GDP per capita, $GDPPC_{i,y-1}$, which proxies for the level of economic development; and annual real GDP growth, $\Delta RGDP_{i,y}$, which captures cyclical economic conditions. All macroeconomic variables are from the World Bank's World Development Indicators.

We also control for institutional and policy-related factors that may affect exchange-rate behavior. Capital account openness, $openness_{i,y}$, is measured using the annual index of Fernandez, Klein, Rebucci, Schindler and Uribe (2016), which captures restrictions on cross-border financial flows. Exchange-rate regimes, $D.regime_{i,ym}$, are controlled for using the monthly classification of Ilzetzki, Reinhart and Rogoff (2019), which distinguishes among crawling pegs, managed floats, and free floats. These controls account for systematic differences in exchange-rate flexibility and policy frameworks across countries.

Given the close link between U.S. monetary policy and global financial conditions (Rey, 2015), we also include cumulative daily changes in the VIX,

$$\Delta VIX_t^h = \frac{VIX_{t+h} - VIX_{t-1}}{VIX_{t-1}} \times 100,$$

with data from Federal Reserve Economic Data (FRED). Country and year fixed effects absorb time-invariant country characteristics and low-frequency global shocks.

To distinguish between conventional and unconventional monetary policy, we follow Swanson (2021), who identify orthogonalized shocks to the federal funds rate, FFR_t, forward guidance, FG_t, and large-scale asset purchases, $LSAP_t$, using high-frequency financial data around FOMC announcements.[1] We focus on FFR_t and FG_t, excluding $LSAP_t$

[1] Swanson (2021) identify exogenous U.S. monetary policy shocks using a factor model estimated by

because of the limited incidence of quantitative tightening in our sample.[2]

The main regression allows the exchange-rate response to differ between conventional and unconventional monetary policy shocks:

$$\Delta E^h_{i,t} = \beta_0 + \beta_1 FFR_t + \beta_2 FG_t + \gamma_1 FFR_t \times depth_{i,y-1} + \gamma_2 FG_t \times depth_{i,y-1} + \alpha_1 depth_{i,y-1}$$
$$+ \delta_1 size_{i,y-1} + \delta_2 GDPPC_{i,y-1} + \delta_3 \Delta RGDP_{i,y} + \delta_4 openness_{i,y} + \delta_5 D.regime_{i,ym}$$
$$+ \eta_1 \Delta VIX^h_t + D.year + D.country + \epsilon^h_{i,t}. \tag{2}$$

Our sample consists of 4,626 observations covering 148 FOMC meetings between January 2001 and June 2019 for 34 countries. The sample period is determined primarily by the availability of monetary policy shock measures. Countries are required to have exchange-rate data for at least 80 percent of FOMC announcement dates.[3] We exclude euro-area countries because the common currency breaks the link between national financial market depth and bilateral exchange-rate movements. We also exclude special administrative regions such as Hong Kong, although the results are robust to their inclusion. The sample countries are grouped by region as follows:

- North and South America: Argentina, Brazil, Canada, Chile, Colombia, Mexico, Peru, Uruguay;
- Asia and Oceania: Australia, India, Indonesia, Japan, Korea, Malaysia, New

principal component analysis of high-frequency financial market responses around FOMC announcements. The input variables include federal funds futures, Eurodollar futures, and Treasury yields at different maturities. Exchange rates are not used as inputs in the factor extraction; they enter only as outcome variables in subsequent regressions. Under the efficient market assumption, movements in these financial variables within a narrow announcement window reflect unanticipated changes in the stance of U.S. monetary policy. The extracted factors are interpreted as monetary policy shocks and are identified as shocks to the federal funds rate (FFR_t), forward guidance (FG_t), and large-scale asset purchases ($LSAP_t$) based on a set of identifying restrictions: the forward-guidance and asset-purchase factors do not affect the spot federal funds rate, and variation in the asset-purchase factor is minimized prior to the Global Financial Crisis.

[2]Because forward guidance is active throughout the sample period, whereas LSAPs were adopted only after the Global Financial Crisis, we focus on forward guidance as our measure of unconventional monetary policy.

[3]Although the 80 percent threshold is somewhat arbitrary, the results are robust to alternative data-availability thresholds.

Zealand, Philippines, Singapore, Thailand;

- Africa and Middle East: Bahrain, Israel, South Africa, Turkey;
- Europe: Bulgaria, Czech Republic, Denmark, Hungary, Iceland, Norway, Poland, Romania, Russia, Sweden, Switzerland, the United Kingdom.

Table 1 reports summary statistics. Panel A summarizes monetary policy shocks. A one-standard-deviation change in FFR_t corresponds to about 5 basis points. A one-standard-deviation change in FG_t corresponds to about 6 basis points in four-quarter-ahead Eurodollar futures. Panels B–D report exchange-rate and macroeconomic statistics for the full sample and for subsamples split by the median level of foreign liabilities. Countries with higher foreign liabilities exhibit somewhat greater exchange-rate volatility following FOMC announcements, consistent with our central hypothesis.

2.2 Suggestive evidence

As a preliminary illustration of the relationship between financial market depth and exchange-rate responses to U.S. monetary policy shocks, we begin with country-by-country regressions. For each country, we estimate

$$\Delta E_t^{h=1} = \beta_0 + \beta_1 MPS_t^{orth} + \epsilon_t,$$

where $\Delta E_t^{h=1}$ denotes the one-day change in the bilateral exchange rate around FOMC announcements, and MPS_t^{orth} is the orthogonalized monetary policy surprise.

We then relate the estimated country-specific coefficients, β_1, to the average level of log foreign liabilities. Figure 1 shows a positive association between the magnitude of exchange-rate responses and the size of foreign liabilities: countries with deeper financial markets tend to exhibit stronger exchange-rate movements following U.S. monetary

Table 1: Summary statistics

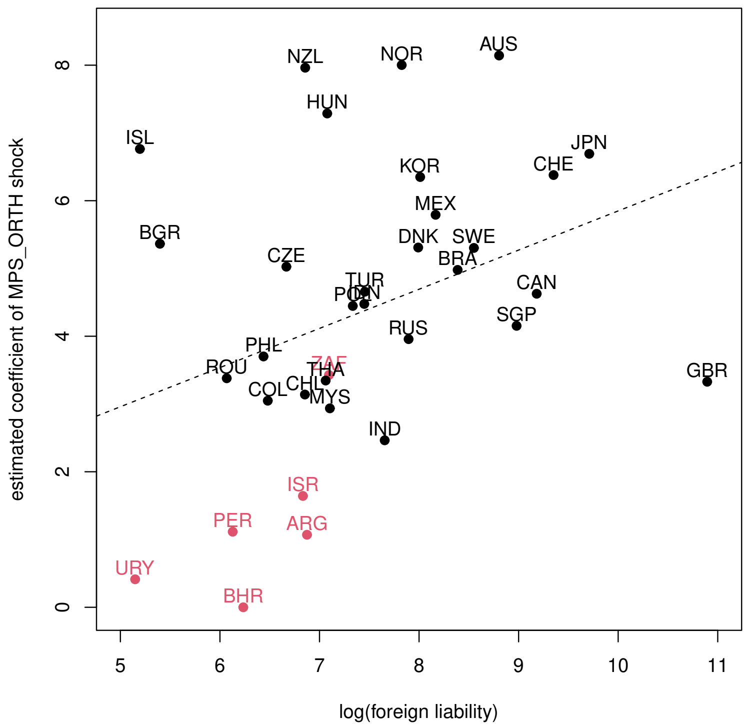
	variable	mean	sd	min	p25	median	p75	max
Panel A: monetary policy shock								
number of FOMC meetings = 148	MPS_t^{orth} (%)	0.00	0.05	-0.23	-0.03	0.00	0.03	0.16
	FFR_t (%)	0.00	0.05	-0.39	0.00	0.01	0.02	0.12
	FG_t (%)	0.00	0.06	-0.19	-0.03	0.00	0.03	0.25
Panel B: full sample								
nobs = 4,626	$\Delta E_t^{h=1}$ (%)	-0.18	1.05	-3.90	-0.68	-0.09	0.32	2.86
	$\Delta E_t^{h=5}$ (%)	-0.05	1.64	-4.88	-0.89	0.00	0.76	5.12
	$\Delta E_t^{h=10}$ (%)	0.17	2.16	-5.65	-0.98	0.01	1.23	7.56
	$\Delta E_t^{h=20}$ (%)	0.14	2.89	-7.03	-1.50	0.00	1.53	10.27
	$\Delta E_t^{h=30}$ (%)	0.09	3.79	-8.75	-2.08	-0.03	1.73	14.44
	$liab$ (ln)	7.43	1.35	3.96	6.57	7.32	8.26	11.31
	$\Delta RGDP$ (%)	3.38	2.99	-10.89	1.80	3.25	5.22	14.52
	$GDPPC$ (%)	9.57	1.07	6.63	8.79	9.41	10.60	11.37
Panel C: low foreign liability country								
nobs = 2,292	$\Delta E_t^{h=1}$	-0.15	1.01	-3.90	-0.57	-0.03	0.27	2.86
	$\Delta E_t^{h=5}$	-0.04	1.61	-4.88	-0.80	0.00	0.68	5.12
	$\Delta E_t^{h=10}$	0.17	2.13	-5.65	-0.86	0.00	1.08	7.56
	$\Delta E_t^{h=20}$	0.15	2.85	-7.03	-1.40	0.00	1.42	10.27
	$\Delta E_t^{h=30}$	0.11	3.69	-8.75	-1.82	0.00	1.65	14.44
	$liab$ (ln)	6.42	0.78	3.96	5.98	6.60	6.97	7.74
	$\Delta RGDP$	3.68	3.06	-10.89	2.21	4.06	5.70	10.43
	$GDPPC$	9.25	0.82	7.52	8.65	9.19	9.77	10.96
Panel D: high foreign liability country								
nobs = 2,334	$\Delta E_t^{h=1}$	-0.21	1.10	-3.90	-0.78	-0.17	0.38	2.86
	$\Delta E_t^{h=5}$	-0.06	1.67	-4.88	-0.97	-0.05	0.88	5.12
	$\Delta E_t^{h=10}$	0.17	2.18	-5.65	-1.10	0.04	1.37	7.56
	$\Delta E_t^{h=20}$	0.13	2.93	-7.03	-1.59	-0.03	1.63	10.27
	$\Delta E_t^{h=30}$	0.08	3.88	-8.75	-2.21	-0.24	1.84	14.44
	$liab$ (ln)	8.42	1.01	6.37	7.78	8.25	9.03	11.31
	$\Delta RGDP$	3.08	2.88	-7.80	1.59	2.83	4.79	14.52
	$GDPPC$	9.88	1.20	6.63	9.08	10.42	10.81	11.37

Note: This table reports summary statistics for key variables in the full sample and in subsamples split by the country-level average of log foreign liabilities. The low group consists of countries below the median, and the high group consists of countries above the median. ΔE_t^h and ΔVIX_t^h are winsorized at the top and bottom 1 percent.

policy shocks.[4] While this evidence is descriptive and does not control for other country characteristics, it motivates the panel regressions below, which formally assess cross-country heterogeneity in exchange-rate responses.

[4]Although not reported for brevity, this relationship remains evident at longer horizons, $h = 5, 10, 20, 30$.

Figure 1: Exchange-rate sensitivity and foreign portfolio liabilities



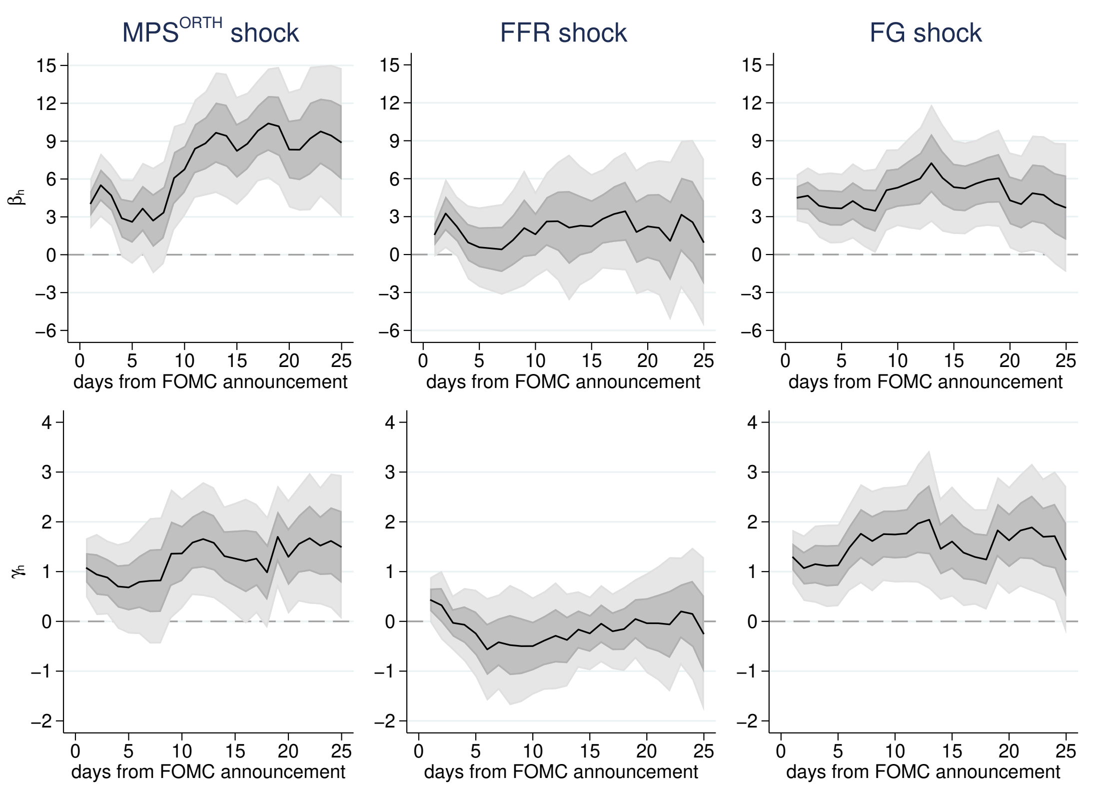
Note: This figure plots country-specific estimates of exchange-rate responses to U.S. monetary policy surprises against the country-level average of log foreign liabilities. Each dot corresponds to a coefficient β_1 estimated from $\Delta E_t^{h=1} = \beta_0 + \beta_1 MPS_t^{orth} + \epsilon_t$. Black dots indicate coefficients that are statistically significant at the 10 percent level.

2.3 Main findings

This section tests whether financial market depth shapes the sensitivity of exchange rates to U.S. monetary policy shocks and whether this effect differs between conventional and unconventional policy instruments. We implement the empirical strategy described in Section 2.1.

We begin by estimating equations (1) and (2). We first examine the average exchange-rate response to exogenous U.S. monetary policy shocks, evaluated at the sample mean of foreign liabilities. The top row of Figure 2 plots the estimated coefficients over horizons, using the aggregate monetary policy surprise, MPS_t^{orth} (first column), and its decomposition into conventional, FFR_t (second column), and forward-guidance, FG_t (third column), components. Shaded areas denote one- and two-standard-deviation confidence intervals, with standard errors clustered at the FOMC-date level.

Figure 2: Response of exchange rates to contractionary U.S. monetary policy shocks, decomposed into baseline effect (top) and depth amplification (bottom)



Note: This figure plots horizon-wise estimates of the baseline coefficient, β_h (top row), and the depth-amplification coefficient, γ_h (bottom row), for the aggregate Bauer–Swanson orthogonalized shock, MPS_t^{orth} (first column), the Swanson (2021) federal funds rate factor, FFR_t (second column), and the forward-guidance factor, FG_t (third column). Each panel reports estimates for varying values of h, the horizon in days after the FOMC announcement. Shaded bands denote one- and two-standard-deviation confidence intervals. Standard errors are clustered at the FOMC-date level.

The estimated coefficients in the first column are positive and statistically significant, indicating that local currencies depreciate against the U.S. dollar following contractionary U.S. monetary policy shocks. This finding is closely related to recent work that revisits deviations from uncovered interest parity using high-frequency event studies and finds UIP-consistent exchange-rate behavior following U.S. monetary policy shocks (Albagli et al., 2024; Rosa, 2011).

The exchange-rate response is larger and more precisely estimated for forward-guidance shocks than for contemporaneous target-rate surprises. This pattern is consistent with Hausman and Wongswan (2011), who document that exchange rates respond more strongly to surprises about the future path of policy rates than to changes in the spot rate, and with Swanson (2021), who shows that path components account for a substantial share of exchange-rate movements on policy announcement days.[5]

The stronger exchange-rate response to forward-guidance shocks is intuitive. By shifting expectations of medium- to long-term interest-rate differentials, forward guidance generates larger present-value effects on exchange rates than an equally sized movement in the spot policy rate. In this regard, Gürkaynak et al. (2021) emphasize that conditioning on path surprises is crucial for understanding exchange-rate behavior around policy announcements and that much of the apparent anomaly in exchange-rate responses dissipates once such surprises are taken into account.[6]

We next turn to the main coefficients of interest, which capture how exchange-rate responses vary with financial market depth. The bottom row of Figure 2 plots the interaction coefficients between monetary policy shocks and lagged foreign portfolio liabilities, our baseline measure of market depth. In the left panel, the interaction between the aggregate monetary policy surprise and foreign liabilities is positive and statistically significant at most horizons. This finding indicates that exchange-rate depreciations following U.S. monetary tightening are larger in countries with deeper financial markets.

Decomposing the aggregate shock, the second and third columns show that the amplification effect is driven primarily by forward-guidance shocks. The interaction between

[5]The weaker estimates for conventional monetary policy shocks, FFR_t, are not driven by the limited variation in target-rate surprises during the zero lower bound period. Restricting the sample to pre-ZLB episodes yields similar results, with exchange rates continuing to respond more strongly to forward-guidance shocks than to contemporaneous target-rate surprises.

[6]Relatedly, evidence from unconventional monetary policy episodes shows that policy announcements affecting expected future monetary conditions can generate sizable exchange-rate movements even when short-term policy rates are constrained, as documented by Neely (2015).

FG_t and foreign liabilities is consistently positive and statistically significant at most horizons, whereas the interaction between FFR_t and foreign liabilities is smaller in magnitude and statistically insignificant. To gauge economic significance, note that the standard deviation of log foreign liabilities in the sample is 1.35. For countries with foreign liabilities one standard deviation below the mean, a one-standard-deviation increase in FG_t implies an exchange-rate depreciation of between 0.10% and 0.14%, depending on the horizon. For countries with foreign liabilities one standard deviation above the mean, the same shock generates a substantially larger depreciation, ranging from 0.26% to 0.38%. These magnitudes indicate substantial cross-country heterogeneity in the transmission of U.S. monetary policy.[7]

These findings are consistent with the literature on Delphic and Odyssean forward guidance and with information-effect interpretations of monetary policy announcements (Fujiwara and Waki, 2022; McKay, Nakamura and Steinsson, 2016). Forward guidance may simultaneously convey information about economic fundamentals and commit to future policy actions, both of which can amplify exchange-rate responses. Rather than assessing the theoretical scope or limits of forward guidance, our goal is to document that the strength of these channels varies systematically with financial market depth and to provide a parsimonious theoretical framework that organizes this cross-country heterogeneity.[8]

[7] To assess whether financial market depth matters beyond the short run, we extend the horizon to 60 trading days. As shown in Figure A1, the foreign-liability interaction coefficients gradually decline after 30 days and converge toward zero, indicating that this channel is primarily relevant at short horizons.

[8] Evidence from models with heterogeneous beliefs (Croitoru, Jiao and Lu, 2024) provides a complementary interpretation of our findings. In such frameworks, exchange rates respond to forward-guidance shocks because these shocks alter expectations about the future path of monetary policy and generate disagreement across market participants. Differences in beliefs affect portfolio allocations and state price densities, leading to larger exchange-rate movements and time-varying risk premia. From this perspective, the stronger exchange-rate response to forward guidance reflects its role in shaping expectations rather than contemporaneous policy-rate changes. Our results are also consistent with this channel and suggest that deeper financial markets amplify these expectation-driven effects by facilitating trading and portfolio rebalancing, regardless of the source of expectation changes.

Table 2: Financial market depth and conventional versus unconventional U.S. monetary policy: summary

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.55*	0.57	1.60	2.23	4.00
	(1.79)	(0.36)	(0.95)	(0.88)	(1.04)
$FFR_t \times liab_{i,y-1}$	0.43*	-0.24	-0.50	-0.03	-0.17
	(1.93)	(-0.56)	(-1.01)	(-0.07)	(-0.19)
FG_t	4.48***	3.65***	5.29***	4.28**	3.64
	(4.88)	(2.66)	(3.49)	(2.26)	(1.27)
$FG_t \times liab_{i,y-1}$	1.30***	1.13***	1.75***	1.63***	0.77
	(4.80)	(2.74)	(3.62)	(3.02)	(0.91)
$liab_{i,y-1}$	-0.08	0.16	0.43**	0.62**	1.01***
	(-0.87)	(1.26)	(2.41)	(2.44)	(3.46)
Adjusted R^2	0.183	0.083	0.110	0.121	0.168
Observations	4626	4626	4626	4626	4626

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table summarizes estimates of equation (2) at selected horizons, $h = 1, 5, 10, 20, 30$. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

2.4 Robustness Checks

We conduct a series of robustness checks to assess the sensitivity of this result to alternative specifications. Throughout, we focus on the key pattern that financial market depth amplifies local-currency depreciations following contractionary U.S. monetary policy shocks, especially for forward-guidance shocks. To conserve space, we report the estimates of interest, β_1, β_2, γ_1, and γ_2 in equation (2), at selected horizons, $h = 1, 5, 10, 20, 30$. To ease comparison, Table 2 summarizes the corresponding coefficients and significance levels from the baseline specification.

Alternative control sets. A first concern is that some controls in the baseline specification are measured at lower frequencies than the daily exchange-rate changes used in our event-study design. If U.S. monetary policy shocks are plausibly exogenous at high frequency, additional controls may be unnecessary and could raise over-controlling concerns.

We therefore re-estimate the regressions sequentially. We begin with a parsimonious specification that includes only monetary policy shocks, their interactions with foreign liabilities, and country and year fixed effects (Panel A). We then add daily changes in the VIX to control for contemporaneous global financial conditions, excluding low-frequency country-level controls (Panel B). Finally, we include slow-moving country-level controls measured at the annual frequency, excluding high-frequency controls (Panel C). Table A1 shows that the interaction between forward guidance shocks and foreign liabilities remains positive, statistically significant, and similar in magnitude across specifications.

FOMC-date fixed effects. We next assess whether the results are driven by unobserved heterogeneity across FOMC meetings. To do so, we replace the direct effects of FFR_t and FG_t with FOMC-date fixed effects, while retaining their interactions with foreign liabilities. This specification allows the average exchange-rate response to vary flexibly across FOMC announcements and absorbs any event-level common shocks. Table A2 shows that only the interaction between forward guidance shocks and foreign liabilities remains positive and statistically significant.

Alternative measures of financial market depth. We next examine whether the main results in Table 2 are sensitive to the choice of financial-depth proxy. Although foreign liabilities provide a natural measure of market depth from the perspective of international investors, they may also reflect equilibrium outcomes rather than underlying financial frictions. We therefore consider two alternative measures. The first is the IMF Financial Development Index, which summarizes the depth, access, and efficiency of financial institutions and markets. The second is the World Bank stock market turnover ratio, defined as the value of shares traded relative to market capitalization, which more directly captures stock-market liquidity.

Tables A3 and A4 report the results. In both cases, countries with more developed or liquid financial markets exhibit larger exchange-rate responses to U.S. monetary policy shocks, especially to forward-guidance shocks. These results indicate that our findings are not specific to foreign liabilities, but instead reflect a broader relationship between financial market depth and exchange-rate sensitivity. They are consistent with the interpretation that portfolio adjustment costs play an important role in generating cross-country heterogeneity.

Gross versus net foreign asset positions. Our baseline measure of financial depth is based on gross foreign liabilities. This measure may conflate two distinct dimensions of external financial integration: the overall scale of cross-border portfolio positions and the net external position. The distinction matters because central banks in countries with large negative net foreign asset positions may have stronger incentives to stabilize nominal exchange rates in order to maintain the capital inflows needed to finance those positions.

To separate these two dimensions, we replace $liab_{i,y-1}$ in equation (2) with gross and net foreign portfolio positions,

$$gross_{i,y-1} = asset_{i,y-1} + liab_{i,y-1}, \qquad net_{i,y-1} = asset_{i,y-1} - liab_{i,y-1},$$

where $asset_{i,y-1}$ and $liab_{i,y-1}$ denote lagged log gross foreign portfolio assets and liabilities, respectively. We refer to $asset_{i,y-1} + liab_{i,y-1}$ as the gross-scale term and $asset_{i,y-1} - liab_{i,y-1}$ as the net-position term. The gross-scale term captures the overall size of cross-border portfolio balance sheets, whereas the net-position term captures external balance-sheet strength. Both terms enter jointly with their interactions with FFR_t and FG_t, so that exchange-rate responses to U.S. monetary policy shocks are allowed to vary separately with each dimension. This specification also provides a direct test of the alternative interpretation: if the heterogeneity were driven entirely by net positions rather than by financial depth, the interaction between FG_t and the gross-scale term should be statistically indistinguishable from zero.

Table A5 shows that this is not the case. The interaction between FG_t and the gross-scale term remains positive and statistically significant conditional on the net-position term, indicating that financial-depth amplification is separately identified rather than an artifact of the net external position. At the same time, the interaction between FG_t and the net-position term is negative and statistically significant at most horizons. Under our convention, in which an increase in $\Delta E^h_{i,t}$ denotes a depreciation against the U.S. dollar, this negative coefficient implies that countries with stronger net external positions depreciate less following U.S. monetary tightening. This pattern is consistent with an external-balance-sheet-strength channel, in which stronger net foreign asset positions dampen depreciation pressures (Corte, Riddiough and Sarno, 2016). It also goes against the stabilization interpretation discussed above: if central banks in countries with large negative net positions were systematically limiting depreciation to preserve capital inflows, we would expect muted depreciations where net positions are most negative. Instead, the estimates imply larger depreciations for countries with weaker net external positions.

Financial-autarky corner. A related concern is that the depth amplification may reflect a corner-solution mechanism rather than gradual portfolio adjustment. If foreign investors hold very few domestic assets initially, they have limited scope to sell those assets after a U.S. tightening shock, especially in the presence of non-negativity constraints on portfolio positions. This mechanism should be most relevant for observations with extremely low foreign liabilities.

To assess this possibility, we re-estimate the baseline specification after excluding the bottom 10 percent of country-year observations by lagged log foreign liabilities. Table A6 shows that the forward-guidance interaction with foreign liabilities remains positive at all reported horizons and statistically significant through $h = 20$. Thus, even after excluding observations most likely to be near the financial-autarky corner, the forward-guidance-specific depth amplification remains intact.

Endogenous responses of domestic monetary policy. Cross-country differences in exchange-rate responses may simply reflect endogenous domestic monetary policy responses to U.S. tightening, especially if such responses are correlated with countries' external positions. Although our high-frequency design mitigates this concern, we examine it directly using daily policy rates obtained from the BIS.

We conduct two exercises. First, we replace the dependent variable in equation (2) with

$$\Delta i_{i,t}^{h} \equiv i_{i,t+h} - i_{i,t-1},$$

the change in country i's policy rate from one day before to h trading days after the FOMC announcement. Table A7 shows that domestic policy rates do not respond systematically to U.S. monetary policy shocks at the horizons relevant for our exchange-rate analysis. Nor is the response systematically related to our measure of financial depth.

Second, we augment the baseline specification with $\Delta i_{i,t}^{h}$ as a level control, which absorbs the component of exchange-rate movements associated with contemporaneous domestic policy-rate adjustments. Because the domestic policy response is itself a post-treatment outcome, this exercise is not our preferred specification. Rather, it is a robustness check that asks whether the depth interaction survives even after conditioning on observed policy-rate changes. Table A8 shows that the depth amplification of exchange-rate responses remains essentially unchanged.

Alternative measure of monetary policy shocks. We next examine whether the results are driven by our baseline three-factor decomposition of FOMC announcement surprises. Following Gürkaynak et al. (2005), we construct the original GSS target and path factors from five short-rate surprises—MP1, MP2, ED2, ED3, and ED4—measured in 30-minute windows around FOMC announcements. The intraday surprises are from the U.S. Monetary Policy Event-Study Database of Acosta, Ajello, Bauer, Loria and Miranda-Agrippino (2025). We restrict the sample to 2001–2019, matching our baseline sample and remaining within the pre-SOFR Eurodollar regime.

This two-factor specification intentionally abstracts from a separate LSAP factor and relies only on short-rate futures. The target factor captures unexpected changes in the current federal funds rate target, while the path factor captures revisions to the expected future path of short rates that are orthogonal to the current target surprise. Thus, the exercise provides a robustness check on whether our results reflect a broader target-versus-path distinction rather than the particular decomposition of federal funds rate, forward-guidance, and LSAP surprises.

Table A9 reports the results. The forward-guidance amplification result carries over to this alternative specification. The interaction between the path factor and financial depth is positive at all horizons and statistically significant at the 1 percent level through $h = 20$. In contrast, the interaction between the target factor and financial depth is small and statistically insignificant throughout. Thus, the path–target asymmetry closely mirrors the baseline forward-guidance–federal-funds-rate asymmetry.

Controlling for dollar strength. Orthogonalized U.S. monetary policy surprises may still be correlated with global risk or other confounding factors that move the dollar broadly. Although the baseline specification controls for daily changes in the VIX, we further address this concern by augmenting the specification with the percent change in the U.S. dollar index, ΔDXY_t^h, measured over the same horizon as the dependent variable. This control absorbs common dollar movements around FOMC announcements, so that the depth interaction is identified from cross-country differences in exchange-rate responses.

Table A10 reports the results. Controlling for dollar strength leaves the depth amplification largely intact. The point estimates are modestly attenuated relative to the baseline, but the qualitative pattern is unchanged. These results indicate that broad dollar movements contribute to the average exchange-rate response but do not account for the cross-country heterogeneity associated with financial depth.

Confounding country characteristics. Our baseline proxy for financial market depth may be correlated with country characteristics that also affect exchange-rate sensitivity. In particular, larger foreign liabilities may partly reflect lower past exchange-rate volatility, stronger institutional quality, or greater exposure to U.S. long-term interest rates. We therefore examine whether the depth amplification survives after controlling for three such characteristics: exchange-rate volatility, institutional quality, and currency exposure to U.S. Treasury yields. For each variable, we augment the baseline specification with the variable itself and its interactions with U.S. monetary policy shocks.[9]

Tables A11–A14 report the results for the FFR–FG decomposition. The interaction between forward guidance and financial depth remains positive and statistically significant at most horizons across these specifications. The additional interaction terms are generally insignificant, although the Treasury-yield beta interactions are occasionally significant. Overall, these results indicate that the depth-conditional response to forward-guidance shocks is not absorbed by pre-existing exchange-rate volatility, institutional quality, or currency exposure to U.S. long-term rates.

Alternative event-level mechanisms. We next examine whether the depth interaction reflects event-level channels rather than financial market depth. We consider two pos-

[9]Exchange-rate volatility, $E^{vol}_{i,m-1}$, is the rolling 12-month standard deviation of daily LCU/USD log returns, measured at the monthly frequency and lagged by one month. Institutional quality, $IQ_{i,y-1}$, is the lagged World Bank Rule of Law index from the Worldwide Governance Indicators. Currency exposure to U.S. Treasury yields is measured by country-specific betas, $\beta^{UST5Y}_{i,m-1}$ and $\beta^{UST10Y}_{i,m-1}$, estimated from rolling 12-month regressions of daily LCU/USD log returns on daily changes in the 5-year and 10-year U.S. Treasury yields, excluding FOMC trading days.

sibilities: changes in global risk aversion around FOMC announcements and residual information effects in U.S. monetary policy surprises.

First, U.S. monetary policy shocks may interact with changes in global risk premia that differentially affect exchange rates. To address this concern, we augment the baseline specification with the change in Moody's Baa corporate bond yield over the 10-year Treasury yield, Baa_t^h, measured over the same horizon as the dependent variable, and its interactions with FFR_t and FG_t.[10] Table A15 shows that the interaction between forward guidance and financial depth remains positive and statistically significant at most horizons. The interaction between forward guidance and the Baa–Treasury spread is significant at longer horizons, suggesting that risk-premium dynamics operate alongside, rather than instead of, the depth channel.

Second, U.S. monetary policy surprises may contain information effects if the Federal Reserve systematically responds to recent macroeconomic or financial news. The Bauer–Swanson orthogonalized shocks are designed to remove such systematic responses. In addition, our specifications with FOMC-date fixed effects absorb all event-level variation (see Table A2). As a direct check, we further add four event-level controls used in the Bauer–Swanson orthogonalization to equation (2): the most recent pre-FOMC nonfarm payroll surprise, the thirteen-week change in the S&P 500 index, the thirteen-week change in the slope of the U.S. Treasury yield curve, and the thirteen-week change in the Bloomberg Commodity Index (Bauer and Swanson, 2023).

Table A16 reports the results. Adding these pre-FOMC macroeconomic and financial controls leaves the depth amplification essentially unchanged. Taken together, the evidence suggests that neither global risk-premium movements nor residual information effects account for the cross-country heterogeneity associated with financial depth.

[10] Our results are also robust to interacting the VIX with both FFR_t and FG_t.

Foreign reserves and central bank intervention. We next examine whether foreign reserve holdings mitigate exchange-rate responses to U.S. monetary policy shocks. Following Chen, Filardo, He and Zhu (2016) and Ahmed, Aizenman, Saadaoui and Uddin (2023), we augment the baseline specification with interactions between monetary policy shocks and the ratio of foreign reserves to GDP. Table A17 shows that foreign reserves appear to dampen exchange-rate movements for rate surprises, although the estimates are generally imprecise. The main result remains unchanged: exchange-rate sensitivity to forward-guidance shocks continues to increase with foreign liabilities.

Relatedly, countries with active foreign exchange intervention may exhibit lower foreign liabilities because of tighter capital-flow restrictions. This raises the possibility that foreign exchange intervention confounds the relation between foreign liabilities and exchange-rate sensitivity (Albagli et al., 2024; Kim, 2003). To address this concern, we interact monetary policy shocks with the monthly country-level measure of foreign currency intervention constructed by Adler, Chang, Mano and Shao (2025), denoted by $FXI_{i,m}$. Table A18 shows that the interaction between forward-guidance shocks and foreign liabilities remains positive and statistically significant at most horizons, even after accounting for foreign exchange intervention.

High- versus middle-income economies. Financial market depth differs systematically between high- and middle-income economies. Thus, the stronger exchange-rate responses observed in deeper markets may partly reflect differences across income groups. For example, Albagli et al. (2024) document systematic differences between advanced and emerging market economies in high-frequency exchange-rate responses to U.S. monetary policy shocks. To address this concern, we estimate equation (2) separately for high- and middle-income countries, based on the World Bank income classification. Table A19 shows that the depth amplification remains present within each income group, indicating that the results are not driven solely by income-level differences.

Alternative standard-error clustering. In the baseline analysis, standard errors are clustered at the FOMC-date level to account for common shocks on announcement days. As an alternative, we implement two-way clustering by FOMC date and country. Table A20 shows that the results are unchanged.

Including the COVID-19 period. Finally, we examine whether the relationship between foreign liabilities and exchange-rate responses changed during the COVID-19 period. We extend the sample through December 2022 and re-estimate equation (1) using the aggregate monetary policy shock, MPS_t^{orth}, because the disaggregated shocks are available only through June 2019. Table A21 shows that the relationship between foreign liabilities and exchange-rate sensitivity continues to hold in the extended sample.

3 Theoretical Explanation

This section develops a tractable two-country, two-period model to organize the empirical findings. The model is not intended to provide a fully structural quantitative account of the empirical magnitudes. Rather, it isolates a mechanism through which gradual portfolio adjustment can generate heterogeneous short-run exchange-rate responses across countries with different levels of financial market depth.

The environment and closed-form representation of the equilibrium exchange rate are inherited from Bacchetta and Van Wincoop (2021). We use this structure for two purposes. First, we allow the effective portfolio adjustment cost to vary with market depth, $\psi = \psi(d_H)$ with $\psi'(d_H) < 0$, which maps cross-country differences in market depth into cross-country differences in effective discounting. Second, and more importantly, we distinguish spot shocks from news shocks and show that market depth selectively amplifies shocks that operate through expected future interest-rate differentials. This news–spot asymmetry is absent from Bacchetta and Van Wincoop (2021) and provides the main analytical link to our empirical finding that the depth interaction is concentrated in forward-guidance shocks rather than contemporaneous target-rate surprises.

Although our empirical analysis focuses on nominal exchange rates, we model the real exchange rate because price-level movements are negligible at the high frequencies studied in the paper, implying near-equivalence of real and nominal exchange-rate dynamics in the short run.

3.1 Benchmark model

The model features overlapping generations of households who live for two periods. There are two countries, indexed by $h = H, F$, and two financial assets: home and foreign bonds. Home households born at time t maximize expected lifetime utility:

$$C_{H,t} + \ln\left(E_t C_{H,t+1}^{1-\gamma}\right)^{\frac{1}{1-\gamma}} - 0.5\psi(d_H)(z_{H,t} - z_{H,t-1})^2, \tag{3}$$

where $C_{H,t}$ denotes consumption and $z_{H,t}$ is the portfolio share invested in foreign bonds. Utility is quasi-linear, with risk aversion parameter γ. Under this specification, savings are fixed at unity, which facilitates the tractable derivation of equilibrium conditions.

The final term represents a quadratic portfolio adjustment cost, capturing the disutility from changing portfolio positions relative to the prior generation. Such adjustment frictions are commonly used to model gradual portfolio rebalancing and may reflect transaction, information-processing, or cognitive costs associated with portfolio changes (Bacchetta, Davenport and Van Wincoop, 2022; Gârleanu and Pedersen, 2013). We allow the effective adjustment cost, $\psi(d_H)$, to decline with the depth of the home bond market, d_H.[11] This is a reduced-form specification. It does not require our country-level depth measures to observe portfolio adjustment costs directly. Rather, it captures the idea that deeper and more liquid markets lower the effective cost of reallocating internationally held assets.

[11] For notational simplicity, we write ψ for $\psi(d_H)$ in the derivations below.

This mapping is motivated by the market microstructure literature, which views bid-ask spreads, price impact, and other liquidity measures as observable counterparts of trading frictions. The bid-ask spread measures the cost of immediacy, while market depth captures the amount of order flow required to move prices by a given amount (Demsetz, 1968; Stoll, 2000). Models of bid-ask spreads with inventory and adverse-selection frictions provide a microfoundation for why transaction prices deviate from frictionless values and why liquidity improves when markets are deeper and more competitive (Copeland and Galai, 1983; Glosten and Milgrom, 1985). Empirically, less liquid assets and markets exhibit larger spreads, greater price impact, and higher estimated round-trip transaction costs. For example, Lesmond, Ogden and Trzcinka (1999) sort firms by the total market value of equity and show that estimated effective transaction costs decline monotonically with firm size.

The formulation should be interpreted as an intensive-margin representation of portfolio rebalancing in markets with positive foreign participation. The model abstracts from occasionally binding portfolio constraints, such as non-negativity constraints on foreign investors' holdings of domestic assets. Such constraints could generate a distinct corner-solution mechanism in economies with extremely low foreign participation. We return to this distinction below when discussing the empirical implications of the model.

The intertemporal budget constraint is:

$$C_{H,t+1} = R^{p,H}_{t+1}(Y_{H,t} - C_{H,t}), \tag{4}$$

where $R^{p,H}_{t+1}$ is the gross portfolio return and $Y_{H,t}$ is a fixed endowment. Since income $Y_{H,t}$ is irrelevant when deriving the equilibrium exchange rate below, we assume that a fixed amount of endowment is given to households.

Portfolio returns are given by:

$$R_{t+1}^{p,H} = \left[z_{H,t} \frac{S_{t+1}}{S_t} e^{i_t^*} + (1 - z_{H,t}) e^{i_t} \right] \frac{P_t}{P_{t+1}}, \tag{5}$$

with i_t and i_t^* denoting home and foreign nominal interest rates, P_t the home CPI, and S_t the nominal exchange rate, expressed as the price of foreign currency in terms of the home currency.

Substituting (4) into (3) and solving yields the optimal portfolio allocation:

$$z_{H,t} - \bar{z}_H = \frac{\psi}{\psi + \gamma\sigma^2}(z_{H,t-1} - \bar{z}_H) + \frac{1}{\psi + \gamma\sigma^2} E_t er_{t+1} \tag{6}$$

where $er_{t+1} = s_{t+1} - s_t + i_t^* - i_t$ denotes the excess return and $\sigma^2 = \mathrm{Var}_t(s_{t+1})$. See Appendix B for the detailed derivation.

Foreign households solve an analogous problem, yielding

$$z_{F,t} - \bar{z}_F = \frac{\psi}{\psi + \gamma\sigma^2}(z_{F,t-1} - \bar{z}_F) + \frac{1}{\psi + \gamma\sigma^2} E_t er_{t+1}. \tag{7}$$

The parameters ψ and γ govern the relative importance of adjustment costs and risk in portfolio choice. A higher ψ raises the cost of changing portfolio positions and therefore makes current portfolios more persistent. A higher γ increases the weight on risk, reducing the sensitivity of portfolio shares to expected excess returns.

Let r_t and r_t^* denote the home and foreign real interest rates, respectively, and let $r_t^D = r_t^* - r_t$ denote the real rate differential. Given $E_t er_{t+1} = E_t q_{t+1} - q_t + r_t^D$ and bond market clearing, the equilibrium real exchange rate satisfies:

$$E_t q_{t+1} - \theta q_t + b\psi q_{t-1} + r_t^D = 0, \tag{8}$$

with $\theta = 1 + \psi b + \gamma\sigma^2 b$ and $b = 0.5\bar{z}_H$.

Solving (8), the equilibrium exchange rate is:

$$q_t = \alpha q_{t-1} + E_t \sum_{i=0}^{\infty} \frac{1}{D^{i+1}} r^D_{t+i}, \tag{9}$$

where

$$\alpha = \frac{\theta - \sqrt{\theta^2 - 4\psi b}}{2}, \tag{10}$$

$$D = \frac{\theta + \sqrt{\theta^2 - 4\psi b}}{2}, \tag{11}$$

where $0 \leq \alpha < 1$ and $D > 1$. The real exchange rate is determined by its lag and the present discounted value of expected future real interest rate differentials.

3.2 Propagation mechanisms: market depth, adjustment costs, and monetary policy shocks

This subsection derives two implications that map directly to the empirical findings. The first is a depth-amplification result: lower adjustment costs in deeper markets increase the exchange-rate response to current or expected real interest-rate differentials. This result connects market depth to the level of exchange-rate sensitivity. The second and central result is a news–spot asymmetry: because lower adjustment costs reduce effective discounting, market depth amplifies shocks to expected future interest-rate differentials more strongly than shocks to the contemporaneous rate. This second result is the main analytical contribution relative to Bacchetta and Van Wincoop (2021) and provides the model counterpart to the empirical asymmetry.

Proposition 1. *Suppose the stable root satisfies $D > 1$. Then an increase in the (current or expected) real interest rate differential induced by a contractionary foreign monetary policy leads to a depreciation of the real exchange rate, with the magnitude of this depreciation increasing as market depth increases.*

The proof can be found in Appendix B.3. The economic intuition behind Proposition 1 is as follows. In economies with shallow financial markets, foreign investors face high costs of portfolio rebalancing. As a result, exchange rate adjustments in response to foreign monetary policy shocks are limited, even when expected interest rate differentials move. By contrast, in deeper financial markets, a larger stock of foreign positions is exposed to changes in current or expected returns, and these positions can be reallocated more readily.

When a contractionary foreign monetary policy shock raises the current or expected real interest rate differential, investors optimally rebalance their portfolios toward higher-yielding assets. In deeper markets, this rebalancing pressure is both larger in scale and more rapidly transmitted to asset prices, leading to a larger exchange rate depreciation. In the model, this mechanism is captured by lower effective discounting of future interest rate differentials as portfolio adjustment costs decline. This theoretical result mirrors the empirical finding in the first column of Figure 2 that currencies of countries with deeper financial markets respond more strongly to contractionary U.S. monetary policy shocks.

A related but distinct mechanism is a financial-autarky or corner-solution channel. If foreign investors initially hold very few domestic assets, a non-negativity constraint may limit their ability to sell those assets after a U.S. tightening shock, muting the exchange-rate response even without adjustment costs. This mechanism operates through an extensive-margin portfolio-availability constraint, whereas our model focuses on intensive-margin rebalancing costs in markets with positive foreign participation. As shown in the robustness checks above, excluding observations with extremely low foreign liabilities does not eliminate the depth amplification.

We next turn to the implication that is new relative to Bacchetta and Van Wincoop (2021). Their framework shows how gradual portfolio adjustment affects exchange-rate dynamics through effective discounting. We use the same structure to ask a different question: whether the effect of market depth depends on the nature of the monetary policy shock. To this end, we discipline the dynamics of the real interest-rate differential, r_t^D, using two distinct policy shocks that mirror the empirical identification: a spot shock and a news (forward-guidance) shock. The two shocks differ not only in their timing but, more importantly, in how their effects are distributed over the expected future path of interest-rate differentials.

The spot shock corresponds to an unanticipated contemporaneous innovation to the real rate differential. Formally, we assume

$$\Delta r_{t+j}^D = \kappa_{\text{spot}} \rho^j, \qquad j \geq 0, \tag{12}$$

where $0 < \rho < 1$ governs persistence and κ_{spot} scales the size of the shock. The impact of the spot shock is therefore concentrated at $t = 0$, with subsequent effects decaying geometrically.

The news shock captures revisions to the expected future path of policy without a contemporaneous change in the current rate. Motivated by the empirical evidence that forward-guidance shocks move medium- and long-term yields more strongly and persistently, we model the news shock as a path shift that lasts for M periods before decaying:

$$\Delta \mathbb{E}_t[r_{t+j}^D] = \begin{cases} 0, & j = 0, \\ \kappa_{\text{news}}, & 1 \leq j \leq M, \\ \kappa_{\text{news}} \rho^{j-M}, & j \geq M+1, \end{cases} \tag{13}$$

where κ_{news} scales the size of the news shock. For simplicity and comparability, we assume that the decay rate in the tail coincides with the persistence parameter ρ.

This specification nests several commonly used formulations as special cases. When $M = 1$, the news shock reduces to a one-period-ahead innovation. As M increases, the shock increasingly resembles a forward-guidance shock that shifts the expected policy path over a longer horizon.

The equilibrium exchange rate at time t can be written as the present discounted value of expected future real interest rate differentials assuming that $q_{t-1} = 0$:

$$q_t = \sum_{j=0}^{\infty} D^{-(j+1)} \, \mathbb{E}_t[r^D_{t+j}], \tag{14}$$

where $D > 1$ is the effective discount factor implied by portfolio adjustment frictions. As shown earlier, higher market depth reduces adjustment costs and therefore lowers D.

We now characterize how market depth shapes the *relative* impact of spot and news shocks.

Proposition 2 (Market depth and news–spot amplification)**.** *Suppose that market depth affects the exchange rate through the adjustment-cost channel, so that deeper markets imply a lower discount factor D. Consider the spot and news shocks defined in* (12)–(13)*. Then:*

1. *The impact response of the exchange rate to a spot shock is*

$$q_0^{spot} = \frac{\kappa_{spot}}{D-\rho}. \tag{15}$$

2. *The impact response to a news shock is*

$$q_0^{news} = \kappa_{news} \left[\frac{1-D^{-M}}{D(D-1)} + \frac{\rho}{D^{M+2}(D-\rho)} \right]. \tag{16}$$

3. *The ratio q_0^{news}/q_0^{spot} is increasing as market depth rises (i.e., as D falls). In particular, deeper markets systematically strengthen the relative importance of news shocks compared to spot shocks.*

The proof can be found in Appendix B.3. Proposition 2 highlights the key analytical prediction of the model. Spot shocks concentrate their mass at impact, while news shocks tilt the expected future path of interest-rate differentials. Market depth, by weakening discounting, increases the relative weight placed on future expectations. Consequently, deeper markets do not amplify all policy shocks by the same degree. Instead, they selectively amplify shocks whose effects materialize through expected future paths. This selective amplification is the main implication that goes beyond the baseline gradual-adjustment environment of Bacchetta and Van Wincoop (2021). It is also the model prediction that maps directly into the empirical finding that the depth interaction is large for forward-guidance shocks but weak for contemporaneous target-rate surprises.

This path–target asymmetry also helps distinguish the adjustment-cost channel from a pure corner-solution mechanism: a binding non-negativity constraint on initial foreign holdings would tend to mute portfolio reallocation broadly, whereas the data show that the depth amplification is concentrated in forward-guidance shocks.

Illustrative sensitivity analysis. To connect the model more directly to the empirical magnitudes, we conduct a simple sensitivity exercise. We follow the calibration in Bacchetta and Van Wincoop (2021) as closely as possible, setting the persistence of the real interest-rate differential to $\rho = 0.95$, the standard deviation of monthly excess returns to $\sigma = 0.027$, and the home-bias parameter to $b = 0.085$. We fix risk aversion at their benchmark value $\gamma = 50$ and vary the adjustment-cost parameter over the range $\psi \in [0, 20]$ considered in their paper. We calibrate the relative sizes of the spot and news shocks to be consistent with the identified FFR and FG shocks used in the empirical analysis.

Appendix Figure B3 reports the results. Panel A shows that, under the benchmark calibration in Bacchetta and Van Wincoop (2021), the model-implied response to news shocks relative to spot shocks increases monotonically as adjustment costs decline. Panel B shows that the region in which $q_0^{news} > q_0^{spot}$ is broad and includes conventional benchmark values. Thus, the model does not require extreme adjustment frictions or parameter choices outside the range considered by Bacchetta and Van Wincoop (2021) to generate the qualitative news-versus-spot asymmetry documented in the data. We do not interpret this exercise as a structural match to the empirical impulse responses or exact interaction coefficients; rather, it shows that plausible variation in adjustment costs can generate the direction and relative ordering of the empirical effects.

Our mechanism is conceptually related to a broad literature that emphasizes financial market imperfections and intermediation frictions in exchange-rate determination (e.g., Bruno and Shin, 2015b; Gabaix and Maggiori, 2015; Itskhoki and Mukhin, 2021). A common insight in this literature is that exchange rates clear international asset markets when investors or intermediaries have limited capacity to absorb currency risk. We view financial market depth as a reduced-form measure of the ease with which international investors can rebalance currency exposure, rather than as a literal measure of intermediary balance-sheet capacity. In our model, deeper and more liquid markets lower effective portfolio adjustment costs, make exchange rates more forward-looking, and increase the weight placed on expected future return differentials.

This mechanism helps reconcile our high-frequency evidence with lower-frequency studies that emphasize balance-sheet vulnerabilities and capital-flow reversals: at daily announcement-window frequencies, exchange-rate movements are more likely to reflect immediate portfolio rebalancing, whereas monthly or quarterly responses also incorporate slower-moving macroeconomic fundamentals and policy adjustments (Akinci and Queralto, 2024; Bruno and Shin, 2015a; Caldara, Ferrante, Iacoviello, Prestipino and Queralto, 2024). It also explains why forward-guidance shocks, which shift the expected future path of interest rates, are amplified more strongly than contemporaneous policy-rate surprises in deeper markets.

4 Conclusion

The growing influence of the Federal Reserve in global financial markets has heightened the importance of understanding the international spillovers of U.S. monetary policy. This paper documents a novel empirical relationship between financial market depth and short-run exchange-rate responses to U.S. monetary tightening. Using high-frequency data around FOMC announcements, we show that countries with deeper financial markets experience larger currency depreciations following contractionary U.S. monetary policy shocks. This depth amplification is concentrated in forward-guidance shocks, consistent with the forward-looking nature of exchange rates.

The effect is economically meaningful and statistically significant at short horizons, but fades within approximately one month. This time profile helps reconcile our findings with lower-frequency evidence on U.S. monetary policy spillovers. Studies using monthly or quarterly data often emphasize the greater vulnerability of emerging markets and financially less developed economies. Our results show that, at daily announcement-window frequencies, deeper markets can exhibit larger exchange-rate responses because portfolio rebalancing is transmitted more quickly to prices.

To interpret these findings, we adapt a gradual portfolio-adjustment model of exchange rates. The model allows effective adjustment costs to decline with market depth and treats forward guidance as a news shock to the expected future path of interest rates. This framework highlights a simple mechanism: deeper markets reduce effective discounting in exchange-rate determination, increasing the weight placed on expected future interest-rate differentials. As a result, shocks that operate through expectations are amplified more strongly in deeper markets than contemporaneous target-rate surprises.

Taken together, the results offer new evidence on the heterogeneous transmission of conventional and unconventional U.S. monetary policy across borders. They also highlight the importance of financial market depth in shaping not only the magnitude, but also the timing, of exchange-rate adjustment.

Declaration of generative AI and AI-assisted technologies in the writing process
While preparing this work, the authors used ChatGPT to reduce grammatical errors and improve readability. After using this tool/service, the authors reviewed and edited the

content as needed and took full responsibility for the content of the publication.

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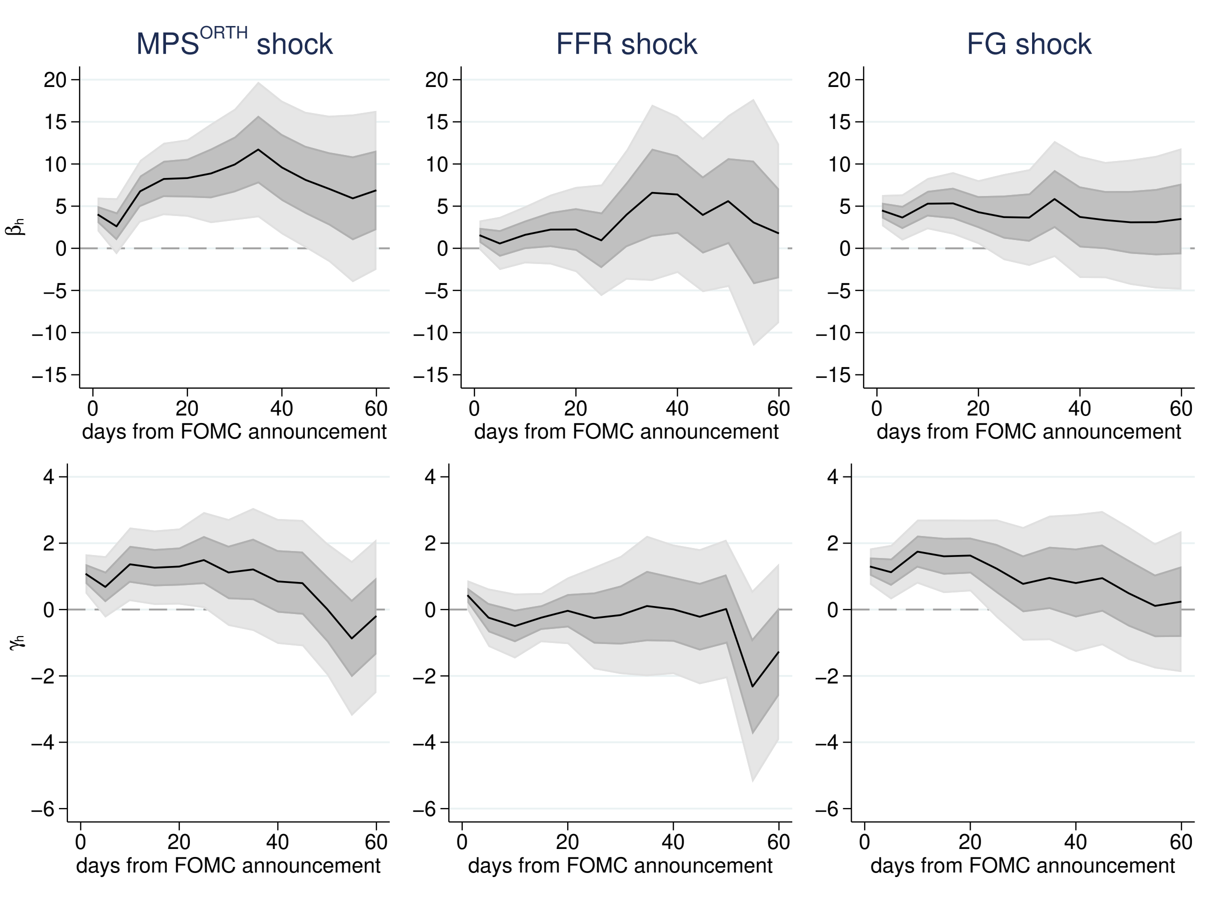
Appendix

A Robustness checks

Figure A1: Response of exchange rates to contractionary U.S. monetary policy shocks: extended horizons



Note: This figure extends Figure 2 up to 60 trading days after the FOMC announcement. Each panel reports estimates for varying values of h, the horizon in days after the FOMC announcement. Shaded bands denote one- and two-standard-deviation confidence intervals. Standard errors are clustered at the FOMC date level.

Table A1: Robustness check: sequential inclusion of covariates

Panel A: without control variables

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.80*	0.58	2.10	4.71	7.11
	(1.92)	(0.36)	(1.17)	(1.65)	(1.38)
$FFR_t \times liab_{i,y-1}$	0.42*	-0.22	-0.49	0.01	-0.17
	(1.92)	(-0.51)	(-0.93)	(0.03)	(-0.15)
FG_t	5.23***	4.17***	6.05***	5.08**	5.64*
	(5.73)	(3.15)	(3.86)	(2.52)	(1.77)
$FG_t \times liab_{i,y-1}$	1.34***	1.13***	1.79***	1.69***	1.01
	(4.89)	(2.78)	(3.66)	(3.07)	(1.13)
$liab_{i,y-1}$	-0.05	0.12	0.36**	0.61***	0.90***
	(-0.71)	(1.12)	(2.28)	(2.74)	(3.37)
Adjusted R^2	0.163	0.076	0.099	0.095	0.121
Observations	4626	4626	4626	4626	4626

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Panel B: adding high-frequency controls only

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.53*	0.57	1.60	2.20	3.96
	(1.77)	(0.36)	(0.94)	(0.87)	(1.02)
$FFR_t \times liab_{i,y-1}$	0.40*	-0.21	-0.44	-0.00	-0.09
	(1.79)	(-0.49)	(-0.88)	(-0.01)	(-0.11)
FG_t	4.48***	3.65***	5.28***	4.27**	3.60
	(4.89)	(2.66)	(3.48)	(2.25)	(1.25)
$FG_t \times liab_{i,y-1}$	1.30***	1.12***	1.74***	1.62***	0.74
	(4.82)	(2.72)	(3.52)	(2.92)	(0.85)
$liab_{i,y-1}$	-0.05	0.12	0.36**	0.61***	0.91***
	(-0.76)	(1.12)	(2.30)	(2.75)	(3.40)
Adjusted R^2	0.183	0.083	0.109	0.119	0.164
Observations	4626	4626	4626	4626	4626

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Panel C: adding low-frequency country-level controls only

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.82*	0.58	2.10	4.74*	7.14
	(1.94)	(0.36)	(1.17)	(1.66)	(1.39)
$FFR_t \times liab_{i,y-1}$	0.45**	-0.25	-0.54	-0.01	-0.24
	(2.08)	(-0.58)	(-1.06)	(-0.02)	(-0.21)
FG_t	5.23***	4.17***	6.06***	5.09**	5.67*
	(5.72)	(3.15)	(3.87)	(2.54)	(1.78)
$FG_t \times liab_{i,y-1}$	1.34***	1.13***	1.80***	1.70***	1.04
	(4.88)	(2.80)	(3.76)	(3.17)	(1.20)
$liab_{i,y-1}$	-0.07	0.16	0.43**	0.62**	1.02***
	(-0.83)	(1.26)	(2.40)	(2.45)	(3.45)
Adjusted R^2	0.163	0.076	0.101	0.097	0.124
Observations	4626	4626	4626	4626	4626

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table reports estimates of equation (2) under alternative sets of control variables. Panel A includes no additional controls. Panel B adds daily changes in the VIX only. Panel C adds low-frequency country-level controls only. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A2: Robustness check: FOMC-date fixed effects

	h=1	h=5	h=10	h=20	h=30
$FFR_t \times liab_{i,y-1}$	0.37**	0.04	-0.23	0.00	-0.23
	(2.08)	(0.20)	(-0.64)	(0.01)	(-0.43)
$FG_t \times liab_{i,y-1}$	0.75***	0.47*	0.97**	1.08***	0.91*
	(3.80)	(1.78)	(2.56)	(2.62)	(1.75)
$liab_{i,y-1}$	-0.07	0.19	0.47**	0.66**	1.07***
	(-0.78)	(1.41)	(2.56)	(2.53)	(3.53)
Adjusted R^2	0.437	0.383	0.378	0.365	0.434
Observations	4626	4626	4626	4626	4626

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) replacing the direct effects of FFR_t and FG_t with FOMC-date fixed effects. The interactions between FFR_t and FG_t and foreign liabilities are retained. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding *t*-statistics reported in parentheses.

Table A3: Robustness check: alternative measure of financial development, IMF financial development index

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.56*	0.76	1.91	2.37	4.00
	(1.76)	(0.51)	(1.18)	(0.98)	(1.09)
$FFR_t \times fdi_{i,y-1}$	1.29**	0.04	0.26	0.70	-0.21
	(2.17)	(0.05)	(0.19)	(0.59)	(-0.14)
FG_t	4.41***	3.56**	5.16***	4.19**	3.69
	(4.75)	(2.60)	(3.38)	(2.22)	(1.31)
$FG_t \times fdi_{i,y-1}$	3.63***	2.93***	5.41***	5.14***	5.43***
	(4.60)	(2.62)	(3.49)	(3.87)	(3.05)
$fdi_{i,y-1}$	-0.15	-0.20	-0.02	-0.27	-0.56
	(-1.11)	(-0.95)	(-0.10)	(-0.67)	(-1.14)
Adjusted R^2	0.181	0.082	0.109	0.120	0.166
Observations	4626	4626	4626	4626	4626

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) using the IMF Financial Development Index of Svirydzenka (2016) as the measure of financial market depth. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A4: Robustness check: alternative measure of financial development, World Bank stock market turnover

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.40	0.85	2.26	2.67	4.22
	(1.59)	(0.59)	(1.42)	(1.13)	(1.18)
$FFR_t \times turnover_{i,y-1}$	0.37*	0.23	-0.25	0.73	0.42
	(1.78)	(1.03)	(-0.75)	(1.56)	(0.63)
FG_t	4.25***	3.52**	5.30***	4.78**	4.05
	(4.73)	(2.57)	(3.47)	(2.53)	(1.43)
$FG_t \times turnover_{i,y-1}$	1.17***	0.72*	1.62***	1.33**	1.80**
	(3.75)	(1.95)	(3.27)	(2.27)	(2.25)
$turnover_{i,y-1}$	0.01	-0.09*	-0.12*	-0.12	-0.33**
	(0.22)	(-1.78)	(-1.71)	(-1.15)	(-2.53)
Adjusted R^2	0.185	0.083	0.107	0.126	0.174
Observations	3864	3864	3864	3864	3864

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) using the World Bank stock market turnover ratio as the measure of financial market depth. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A5: Robustness check: joint estimation of gross and net foreign asset positions

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.40*	0.53	1.49	1.81	3.57
	(1.68)	(0.36)	(0.93)	(0.76)	(1.01)
$FFR_t \times (asset_{i,y-1} + liab_{i,y-1})$	0.28**	-0.036	-0.042	0.43	0.43
	(2.51)	(-0.15)	(-0.16)	(1.16)	(0.65)
$FFR_t \times (asset_{i,y-1} - liab_{i,y-1})$	-1.28*	-1.29	-3.07	-7.14**	-8.04
	(-1.77)	(-0.95)	(-1.57)	(-2.47)	(-1.59)
FG_t	4.15***	3.33**	4.83***	3.93**	3.74
	(4.75)	(2.54)	(3.29)	(2.17)	(1.41)
$FG_t \times (asset_{i,y-1} + liab_{i,y-1})$	0.72***	0.65***	0.97***	0.83**	0.14
	(4.87)	(2.89)	(3.74)	(2.48)	(0.26)
$FG_t \times (asset_{i,y-1} - liab_{i,y-1})$	-1.84**	-2.12*	-2.56*	-1.14	3.68
	(-2.10)	(-1.86)	(-1.85)	(-0.50)	(1.11)
$(asset_{i,y-1} + liab_{i,y-1})$	-0.053	0.087	0.21**	0.31**	0.49***
	(-1.11)	(1.24)	(2.17)	(2.32)	(3.19)
$(asset_{i,y-1} - liab_{i,y-1})$	-0.045	-0.028	-0.23	-0.19	-0.47
	(-0.48)	(-0.20)	(-1.27)	(-0.66)	(-1.48)
Adjusted R^2	0.183	0.083	0.111	0.123	0.169
Observations	4626	4626	4626	4626	4626

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) replacing $depth_{i,y-1}$ with two components of external portfolio positions: the gross-size term, $asset_{i,y-1} + liab_{i,y-1}$ and the net-position term, $asset_{i,y-1} - liab_{i,y-1}$. Both terms enter jointly with their interactions with FFR_t and FG_t. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A6: Robustness check: excluding the bottom 10 percent of foreign liabilities

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.66*	0.62	1.88	2.72	4.63
	(1.84)	(0.40)	(1.10)	(1.01)	(1.17)
$FFR_t \times liab_{i,y-1}$	0.24	-0.52	-1.50**	-1.25**	-1.44*
	(0.86)	(-0.93)	(-2.28)	(-2.58)	(-1.80)
FG_t	4.47***	3.86***	5.57***	4.42**	3.34
	(4.78)	(2.77)	(3.67)	(2.29)	(1.16)
$FG_t \times liab_{i,y-1}$	1.28***	0.71*	1.16**	1.26**	1.01
	(4.36)	(1.71)	(2.36)	(2.36)	(1.16)
$liab_{i,y-1}$	-0.058	0.27*	0.56**	0.72**	1.23***
	(-0.51)	(1.66)	(2.52)	(2.24)	(3.15)
Adjusted R^2	0.189	0.085	0.115	0.128	0.172
Observations	4157	4157	4157	4157	4157

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) after excluding the bottom 10 percent of country-year observations by the level of foreign liabilities. The exclusion cutoff is the 10th percentile of $liab_{i,y-1}$ in the analysis sample. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A7: Robustness check: domestic monetary policy responses to U.S. monetary policy shocks

	h=1	h=5	h=10	h=20	h=30
FFR_t	0.072	-0.022	-0.49	0.83	0.81
	(0.10)	(-0.02)	(-0.50)	(0.26)	(0.25)
$FFR_t \times liab_{i,y-1}$	-0.036	-0.014	0.061	0.21	0.30
	(-0.43)	(-0.13)	(0.54)	(0.60)	(0.82)
FG_t	0.48	0.41	0.42	-0.44	-0.45
	(1.36)	(1.36)	(1.21)	(-0.34)	(-0.34)
$FG_t \times liab_{i,y-1}$	-0.10	-0.12	-0.16	-0.099	-0.054
	(-1.50)	(-1.44)	(-1.45)	(-0.39)	(-0.21)
$liab_{i,y-1}$	-0.84	-0.74	-0.65	2.21	2.11
	(-1.55)	(-1.25)	(-1.08)	(1.01)	(0.96)
Adjusted R^2	0.042	0.026	0.018	0.025	0.023
Observations	4024	4024	4024	4024	4024

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) with the dependent variable replaced by $\Delta i^h_{i,t} \equiv i_{i,t+h} - i_{i,t-1}$, the change in country i's policy rate from one day before to h trading days after the FOMC announcement. Policy rates are from the BIS daily policy rates database. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A8: Robustness check: controlling for domestic monetary policy responses

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.60*	0.65	2.17	2.13	4.15
	(1.73)	(0.40)	(1.24)	(0.80)	(1.03)
$FFR_t \times liab_{i,y-1}$	0.26	-0.49	-1.28**	-0.75	-1.16
	(1.03)	(-0.90)	(-2.06)	(-1.51)	(-1.31)
FG_t	4.53***	3.73**	5.45***	4.45**	3.87
	(4.76)	(2.60)	(3.49)	(2.25)	(1.31)
$FG_t \times liab_{i,y-1}$	1.37***	0.97**	1.57***	1.46**	0.66
	(4.59)	(2.26)	(3.10)	(2.60)	(0.71)
$liab_{i,y-1}$	-0.12	0.12	0.41*	0.56*	1.08***
	(-1.06)	(0.74)	(1.75)	(1.80)	(2.97)
Adjusted R^2	0.187	0.085	0.114	0.127	0.174
Observations	4024	4024	4024	4024	4024

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) augmented with $\Delta i^h_{i,t} \equiv i_{i,t+h} - i_{i,t-1}$ as a level control. $\Delta i^h_{i,t}$ is the change in country i's policy rate from one day before to h trading days after the FOMC announcement. Policy rates are from the BIS daily policy rates database. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A9: Robustness check: using GSS target and path factors

	h=1	h=5	h=10	h=20	h=30
GSS_t^{target}	0.53	0.033	0.97	1.69	3.25
	(0.74)	(0.03)	(0.76)	(0.91)	(1.12)
$GSS_t^{target} \times liab_{i,y-1}$	0.11	-0.36	-0.66	-0.31	-0.26
	(0.55)	(-0.97)	(-1.59)	(-0.78)	(-0.36)
GSS_t^{path}	1.36***	0.98**	1.77***	1.55**	1.57
	(4.43)	(2.17)	(3.25)	(2.40)	(1.62)
$GSS_t^{path} \times liab_{i,y-1}$	0.43***	0.38***	0.60***	0.53***	0.27
	(4.66)	(2.92)	(3.55)	(2.91)	(0.99)
$liab_{i,y-1}$	-0.081	0.16	0.43**	0.62**	1.01***
	(-0.87)	(1.26)	(2.42)	(2.45)	(3.47)
Adjusted R^2	0.166	0.076	0.108	0.122	0.170
Observations	4626	4626	4626	4626	4626

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) replacing FFR_t and FG_t with the target factor, GSS_t^{target}, and the path factor, GSS_t^{path}, constructed following Gürkaynak et al. (2005). The factors are extracted from five short-rate surprise series, MP1, MP2, ED2, ED3, and ED4, measured in 30-minute windows around FOMC announcements. Exchange rates are not used as inputs. Intraday surprises are from the U.S. Monetary Policy Event-Study Database of Acosta et al. (2025). The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A10: Robustness check: controlling for the U.S. dollar index

	h=1	h=5	h=10	h=20	h=30
FFR_t	-0.59	0.62	0.25	1.98	1.93
	(-1.05)	(0.85)	(0.24)	(1.59)	(1.37)
$FFR_t \times liab_{i,y-1}$	0.34*	0.14	-0.17	0.0094	-0.41
	(1.86)	(0.58)	(-0.42)	(0.02)	(-0.72)
FG_t	0.92	0.48	1.82**	2.93**	1.89
	(1.53)	(0.69)	(2.29)	(2.58)	(1.30)
$FG_t \times liab_{i,y-1}$	1.06***	0.57**	1.15**	1.48***	1.39**
	(5.35)	(1.98)	(2.61)	(3.45)	(2.30)
$liab_{i,y-1}$	-0.076	0.19	0.47**	0.66**	1.07***
	(-0.82)	(1.49)	(2.60)	(2.60)	(3.61)
Adjusted R^2	0.347	0.310	0.311	0.308	0.383
Observations	4626	4626	4626	4626	4626

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) augmented with ΔDXY_t^h as an additional control, where ΔDXY_t^h is the percentage change in the U.S. dollar index from one day before to h days after the FOMC announcement. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A11: Robustness check: controlling for past exchange-rate volatility

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.33	0.42	1.41	2.08	3.91
	(1.62)	(0.26)	(0.81)	(0.81)	(0.98)
$FFR_t \times liab_{i,y-1}$	0.38*	-0.31	-0.56	-0.17	-0.30
	(1.76)	(-0.72)	(-1.09)	(-0.35)	(-0.34)
$FFR_t \times E^{vol}_{i,m-1}$	-0.16	-0.10	-0.52	3.39**	2.22
	(-0.30)	(-0.15)	(-0.37)	(2.04)	(1.07)
FG_t	4.32***	3.58***	5.20***	4.18**	3.61
	(5.03)	(2.67)	(3.48)	(2.23)	(1.27)
$FG_t \times liab_{i,y-1}$	1.21***	1.12***	1.71***	1.68***	0.89
	(4.57)	(2.77)	(3.52)	(3.24)	(1.08)
$FG_t \times E^{vol}_{i,m-1}$	2.85	0.13	1.09	-1.95	-4.23
	(1.61)	(0.05)	(0.43)	(-0.59)	(-1.02)
$liab_{i,y-1}$	-0.054	0.21*	0.49***	0.65**	1.09***
	(-0.60)	(1.66)	(2.71)	(2.55)	(3.66)
$E^{vol}_{i,m-1}$	-0.23**	-0.48***	-0.48**	-0.48*	-0.83**
	(-1.99)	(-2.66)	(-2.19)	(-1.67)	(-2.26)
Adjusted R^2	0.190	0.087	0.113	0.124	0.171
Observations	4626	4626	4626	4626	4626

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) including additional interaction terms $FFR_t \times E^{vol}_{i,m-1}$ and $FG_t \times E^{vol}_{i,m-1}$. $E^{vol}_{i,m-1}$ is the rolling standard deviation of daily LCU/USD log returns over the prior twelve months, evaluated at the end of the calendar month preceding the FOMC announcement and demeaned across the sample. The level term $E^{vol}_{i,m-1}$ is also included. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A12: Robustness check: controlling for institutional quality

	h=1	h=5	h=10	h=20	h=30
FFR_t	2.02	-1.04	0.88	3.47	6.95
	(1.44)	(-0.32)	(0.25)	(0.66)	(0.86)
$FFR_t \times liab_{i,y-1}$	-0.27	-1.15*	-2.75***	-1.60	-2.66*
	(-0.51)	(-1.79)	(-3.68)	(-1.19)	(-1.94)
$FFR_t \times IQ_{i,y-1}$	1.60**	0.44	2.67	1.93	2.91
	(2.45)	(0.34)	(1.62)	(0.99)	(1.15)
FG_t	4.83***	4.38***	6.03***	4.60**	3.59
	(4.71)	(2.88)	(3.51)	(2.19)	(1.10)
$FG_t \times liab_{i,y-1}$	1.07***	1.00**	1.43***	1.61**	0.36
	(3.63)	(2.07)	(2.82)	(2.30)	(0.33)
$FG_t \times IQ_{i,y-1}$	0.59	0.033	0.42	-0.66	1.12
	(1.45)	(0.04)	(0.41)	(-0.64)	(0.77)
$liab_{i,y-1}$	-0.12	0.11	0.45**	0.65**	1.15***
	(-1.07)	(0.76)	(2.18)	(2.35)	(3.83)
$IQ_{i,y-1}$	0.10	0.021	-0.14	-0.25	-0.12
	(0.98)	(0.12)	(-0.67)	(-0.84)	(-0.36)
Adjusted R^2	0.197	0.091	0.118	0.130	0.174
Observations	4165	4165	4165	4165	4165

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) including additional interaction terms $FFR_t \times IQ_{i,y-1}$ and $FG_t \times IQ_{i,y-1}$. $IQ_{i,y-1}$ is the lagged Rule of Law index from the World Bank's Worldwide Governance Indicators, demeaned across the sample. The level term $IQ_{i,y-1}$ is also included. The reduced number of observations relative to the baseline reflects country-years for which WGI data are unavailable. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A13: Robustness check: controlling for currency exposure to the 5-year U.S. Treasury yield

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.16	0.54	1.77	3.16	5.94
	(1.24)	(0.31)	(0.83)	(0.97)	(1.29)
$FFR_t \times liab_{i,y-1}$	0.41	-0.33	-0.66	-0.33	-0.58
	(1.65)	(-0.77)	(-1.29)	(-0.76)	(-0.78)
$FFR_t \times \beta^{TR5Y}_{i,m-1}$	0.44	-0.27	-0.80	-2.05	-3.08
	(0.62)	(-0.34)	(-0.59)	(-1.24)	(-1.56)
FG_t	4.86***	3.99***	5.85***	4.59**	3.28
	(4.95)	(2.64)	(3.73)	(2.30)	(1.05)
$FG_t \times liab_{i,y-1}$	1.15***	0.97***	1.50***	1.43***	0.75
	(4.68)	(2.65)	(3.25)	(2.78)	(0.96)
$FG_t \times \beta^{TR5Y}_{i,m-1}$	-0.97*	-1.16	-2.03**	-1.66	0.37
	(-1.92)	(-1.27)	(-2.25)	(-1.52)	(0.23)
$liab_{i,y-1}$	-0.14	0.11	0.42*	0.59**	1.08***
	(-1.20)	(0.71)	(1.96)	(1.99)	(3.13)
$\beta^{TR5Y}_{i,m-1}$	0.043	0.020	0.020	0.030	0.11
	(1.56)	(0.47)	(0.36)	(0.48)	(1.39)
Adjusted R^2	0.193	0.088	0.118	0.127	0.174
Observations	4491	4491	4491	4491	4491

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) including additional interaction terms $FFR_t \times \beta^{UST5Y}_{i,m-1}$ and $FG_t \times \beta^{UST5Y}_{i,m-1}$. $\beta^{UST5Y}_{i,m-1}$ is the country-specific currency beta to changes in the 5-year U.S. Treasury yield, estimated from rolling 12-month regressions of daily LCU/USD log returns on daily yield changes, excluding FOMC trading days. The beta is demeaned across the sample, and its level term is also included. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A14: Robustness check: controlling for currency exposure to the 10-year U.S. Treasury yield

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.13	0.14	1.51	2.25	5.04
	(1.15)	(0.08)	(0.67)	(0.63)	(1.00)
$FFR_t \times liab_{i,y-1}$	0.40	-0.27	-0.61	-0.15	-0.40
	(1.64)	(-0.62)	(-1.25)	(-0.30)	(-0.52)
$FFR_t \times \beta^{TR10Y}_{i,m-1}$	0.39	0.38	-0.26	-0.29	-1.17
	(0.58)	(0.48)	(-0.23)	(-0.14)	(-0.54)
FG_t	4.84***	4.01***	5.83***	4.70**	3.42
	(4.97)	(2.68)	(3.75)	(2.37)	(1.10)
$FG_t \times liab_{i,y-1}$	1.12***	0.96***	1.48***	1.41***	0.74
	(4.59)	(2.61)	(3.18)	(2.72)	(0.96)
$FG_t \times \beta^{TR10Y}_{i,m-1}$	-1.06**	-1.09	-1.93**	-1.51	0.47
	(-2.09)	(-1.20)	(-2.12)	(-1.36)	(0.29)
$liab_{i,y-1}$	-0.14	0.11	0.42*	0.58**	1.06***
	(-1.21)	(0.71)	(1.97)	(1.98)	(3.09)
$\beta^{TR10Y}_{i,m-1}$	0.059**	0.032	0.040	0.026	0.12
	(2.11)	(0.74)	(0.72)	(0.41)	(1.36)
Adjusted R^2	0.196	0.088	0.117	0.125	0.173
Observations	4491	4491	4491	4491	4491

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) including additional interaction terms $FFR_t \times \beta^{UST10Y}_{i,m-1}$ and $FG_t \times \beta^{UST10Y}_{i,m-1}$. $\beta^{UST10Y}_{i,m-1}$ is the country-specific currency beta to changes in the 10-year U.S. Treasury yield, estimated from rolling 12-month regressions of daily LCU/USD log returns on daily yield changes, excluding FOMC trading days. The beta is demeaned across the sample, and its level term is also included. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A15: Robustness check: controlling for global risk aversion

	h=1	h=5	h=10	h=20	h=30
FFR_t	2.05**	-0.48	0.17	3.23	4.24
	(2.06)	(-0.22)	(0.10)	(1.23)	(1.18)
$FFR_t \times liab_{i,y-1}$	0.48**	-0.091	-0.11	0.053	-0.17
	(2.09)	(-0.26)	(-0.30)	(0.12)	(-0.23)
$FFR_t \times Baa_t^h$	11.2	-21.2	-12.5	8.90*	9.08
	(0.79)	(-1.08)	(-1.59)	(1.67)	(1.36)
FG_t	3.88***	3.77**	5.70***	5.40***	6.11**
	(4.29)	(2.42)	(3.86)	(3.10)	(2.26)
$FG_t \times liab_{i,y-1}$	1.27***	1.04**	1.61***	1.62***	1.17
	(4.94)	(2.49)	(3.20)	(2.96)	(1.45)
$FG_t \times Baa_t^h$	5.84	6.77	8.38	22.8**	20.7*
	(0.41)	(0.42)	(0.74)	(2.23)	(1.87)
$liab_{i,y-1}$	-0.080	0.17	0.46**	0.62**	1.02***
	(-0.86)	(1.30)	(2.54)	(2.45)	(3.48)
Baa_t^h	-2.51	0.24	1.27	1.81**	3.36***
	(-1.59)	(0.19)	(1.35)	(2.35)	(4.71)
Adjusted R^2	0.197	0.080	0.127	0.142	0.227
Observations	4594	4592	4558	4626	4626

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) including additional interaction terms $FFR_t \times Baa_t^h$ and $FG_t \times Baa_t^h$. Baa_t^h is the change in Moody's Baa corporate bond yield over the 10-year Treasury yield from one day before to h days after the FOMC announcement. The level term Baa_t^h is also included. Because Baa_t^h is horizon-specific, its coefficient and corresponding interactions appear only in the column matching the relevant horizon. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A16: Robustness check: controlling for pre-FOMC macroeconomic and financial conditions

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.76**	0.49	2.44	4.64**	6.77**
	(1.99)	(0.29)	(1.48)	(2.44)	(2.24)
$FFR_t \times liab_{i,y-1}$	0.43*	-0.22	-0.35	0.27	0.24
	(1.94)	(-0.50)	(-0.75)	(0.62)	(0.34)
FG_t	4.76***	4.22***	6.44***	6.27***	6.97**
	(5.12)	(3.02)	(3.98)	(3.47)	(2.61)
$FG_t \times liab_{i,y-1}$	1.33***	1.18***	1.82***	1.77***	1.03
	(4.95)	(2.87)	(3.78)	(3.43)	(1.31)
$liab_{i,y-1}$	-0.080	0.16	0.44**	0.62**	1.02***
	(-0.86)	(1.26)	(2.43)	(2.45)	(3.45)
NFP_t^{surp}	-0.00037	0.0014	0.00027	-0.0017	-0.00080
	(-0.59)	(1.25)	(0.21)	(-1.18)	(-0.37)
$SP500_t^{3M}$	1.21	-2.55	-2.17	-0.43	-4.73
	(1.39)	(-1.56)	(-1.35)	(-0.23)	(-1.47)
$Slope_t^{3M}$	-0.22	-0.29	-0.19	-0.37	-0.91**
	(-1.61)	(-1.33)	(-0.68)	(-1.14)	(-2.32)
$BCOM_t^{3M}$	-0.87	0.48	-1.43	-4.75***	-4.79*
	(-1.08)	(0.41)	(-1.06)	(-2.92)	(-1.94)
Adjusted R^2	0.193	0.096	0.120	0.145	0.202
Observations	4626	4626	4626	4626	4626

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) augmented with four event-level controls used in the Bauer–Swanson orthogonalization (Bauer and Swanson, 2023). NFP_t^{surp} is the most recent pre-FOMC nonfarm payroll surprise, measured as the actual BLS release minus the Bloomberg consensus expectation. $SP500_t^{3M}$ is the log change in the S&P 500 index over the thirteen weeks ending the day before the FOMC announcement. $Slope_t^{3M}$ is the change in the 10-year minus 2-year Treasury yield spread over the same window. $BCOM_t^{3M}$ is the log change in the Bloomberg Commodity Index over the same window. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A17: Robustness check: controlling for foreign reserves

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.59*	0.46	1.29	1.88	3.59
	(1.93)	(0.28)	(0.75)	(0.71)	(0.91)
$FFR_t \times liab_{i,y-1}$	0.44*	-0.27	-0.57	-0.13	-0.28
	(1.83)	(-0.58)	(-1.17)	(-0.25)	(-0.31)
$FFR_t \times reserve_{i,y-1}$	0.83	-1.83	-5.82	-7.37	-9.08
	(0.28)	(-0.54)	(-1.44)	(-1.00)	(-0.90)
FG_t	4.50***	3.68***	5.33***	4.28**	3.59
	(4.92)	(2.70)	(3.52)	(2.28)	(1.26)
$FG_t \times liab_{i,y-1}$	1.30***	1.13***	1.74***	1.62***	0.77
	(4.80)	(2.75)	(3.65)	(3.01)	(0.90)
$FG_t \times reserve_{i,y-1}$	1.10	1.78	1.09	-2.81	-6.88
	(0.64)	(0.84)	(0.42)	(-0.70)	(-1.36)
$liab_{i,y-1}$	-0.08	0.17	0.46**	0.64**	1.05***
	(-0.85)	(1.30)	(2.48)	(2.49)	(3.53)
$reserve_{i,y-1}$	0.05	-0.15	-0.54	-0.60	-1.02*
	(0.31)	(-0.54)	(-1.59)	(-1.29)	(-1.80)
Adjusted R^2	0.182	0.083	0.110	0.121	0.168
Observations	4626	4626	4626	4626	4626

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) including additional interaction terms between monetary policy shocks and the lagged foreign-reserves-to-GDP ratio $reserve_{i,y-1}$. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A18: Robustness check: controlling for foreign currency intervention

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.59*	0.51	1.91	2.60	3.98
	(1.71)	(0.30)	(0.99)	(0.87)	(0.87)
$FFR_t \times liab_{i,y-1}$	0.58*	-0.19	-0.01	0.73	0.16
	(1.96)	(-0.31)	(-0.01)	(0.81)	(0.11)
$FFR_t \times FXI_{i,m-1}$	0.04	0.16	0.66	-0.42	-0.35
	(0.09)	(0.24)	(0.83)	(-0.33)	(-0.23)
FG_t	4.76***	4.10***	5.90***	4.68**	3.70
	(5.05)	(2.82)	(3.76)	(2.33)	(1.18)
$FG_t \times liab_{i,y-1}$	1.77***	1.86***	2.69***	2.32***	1.05
	(5.58)	(3.35)	(4.55)	(3.13)	(0.86)
$FG_t \times FXI_{i,m-1}$	-0.20	-0.31	-0.59	-0.53	0.17
	(-0.70)	(-0.65)	(-1.14)	(-0.81)	(0.23)
$liab_{i,y-1}$	-0.06	0.18	0.44**	0.62**	0.97***
	(-0.75)	(1.39)	(2.45)	(2.47)	(3.36)
$FXI_{i,m-1}$	-0.01	-0.02	-0.01	-0.01	0.00
	(-0.53)	(-0.98)	(-0.39)	(-0.28)	(0.02)
Adjusted R^2	0.191	0.090	0.116	0.130	0.177
Observations	4346	4346	4346	4346	4346

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) including additional interaction terms between monetary policy shocks and the lagged foreign currency intervention measure, $FXI_{i,m-1}$, from Adler et al. (2025). Bahrain and Iceland are excluded because of missing FXI data. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A19: Robustness check: high- versus middle-income countries

Panel A: high-income countries

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.58*	0.43	1.86	1.50	3.31
	(1.67)	(0.27)	(1.09)	(0.66)	(0.89)
$FFR_t \times liab_{i,y-1}$	0.48*	-0.22	-0.80*	-0.52	-0.91
	(1.68)	(-0.53)	(-1.82)	(-1.01)	(-1.02)
FG_t	4.64***	3.35**	5.25***	3.81**	4.43
	(4.73)	(2.43)	(3.29)	(2.06)	(1.57)
$FG_t \times liab_{i,y-1}$	1.04***	0.88**	1.30***	0.95	0.13
	(3.58)	(2.24)	(2.62)	(1.63)	(0.16)
$liab_{i,y-1}$	0.01	0.26	0.46**	0.71**	1.01***
	(0.06)	(1.34)	(1.99)	(2.32)	(2.73)
Adjusted R^2	0.187	0.091	0.132	0.132	0.174
Observations	2588	2588	2588	2588	2588

Panel B: middle-income countries

	h=1	h=5	h=10	h=20	h=30
FFR_t	0.88	-0.11	1.42	5.88	8.71
	(0.83)	(-0.06)	(0.53)	(1.34)	(1.47)
$FFR_t \times liab_{i,y-1}$	-0.07	-0.83	-0.07	2.81	3.41
	(-0.13)	(-1.01)	(-0.07)	(1.53)	(1.33)
FG_t	5.03***	5.22***	6.77***	7.10***	3.40
	(5.10)	(3.08)	(3.74)	(2.78)	(0.88)
$FG_t \times liab_{i,y-1}$	2.26***	2.77***	3.71***	4.89***	1.97
	(3.80)	(2.70)	(3.43)	(3.63)	(0.98)
$liab_{i,y-1}$	-0.21	0.13	0.57*	0.66	1.08**
	(-1.22)	(0.51)	(1.76)	(1.58)	(2.13)
Adjusted R^2	0.188	0.088	0.091	0.117	0.162
Observations	2038	2038	2038	2038	2038

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) separately by income group, based on the World Bank income classification. Panel A reports estimates for high-income countries, and Panel B reports estimates for middle-income countries. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

Table A20: Robustness check: alternative standard-error clustering

	h=1	h=5	h=10	h=20	h=30
FFR_t	1.55**	0.57	1.60	2.23	4.00
	(2.72)	(0.39)	(0.83)	(0.87)	(1.40)
$FFR_t \times liab_{i,y-1}$	0.43*	-0.24	-0.50	-0.03	-0.17
	(1.97)	(-0.50)	(-0.70)	(-0.07)	(-0.26)
FG_t	4.48***	3.65**	5.29***	4.28**	3.64
	(4.94)	(2.67)	(3.27)	(2.19)	(1.46)
$FG_t \times liab_{i,y-1}$	1.30***	1.13**	1.75**	1.63**	0.77
	(3.51)	(2.18)	(2.45)	(2.54)	(1.11)
$liab_{i,y-1}$	-0.08	0.16	0.43*	0.62*	1.01**
	(-0.82)	(1.09)	(1.81)	(1.96)	(2.55)
Adjusted R^2	0.182	0.083	0.110	0.121	0.168
Observations	4626	4626	4626	4626	4626

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (2) using two-way clustered standard errors by FOMC date and country. The remaining controls and fixed effects follow Table 2. The corresponding t-statistics are reported in parentheses.

Table A21: Robustness check: COVID-19 pandemic, January 3, 2001–December 14, 2022

	h=1	h=5	h=10	h=20	h=30
MPS_t^{orth}	4.03***	2.66	6.62***	8.24***	9.72***
	(4.16)	(1.62)	(3.63)	(3.66)	(2.99)
$MPS_t^{orth} \times liab_{i,y-1}$	1.06***	0.68	1.29**	1.20**	0.94
	(3.70)	(1.50)	(2.38)	(2.10)	(1.18)
$liab_{i,y-1}$	-0.06	0.18	0.42**	0.62**	0.98***
	(-0.73)	(1.46)	(2.40)	(2.53)	(3.38)
Adjusted R^2	0.155	0.074	0.114	0.134	0.177
Observations	4750	4750	4750	4750	4750

t statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: This table re-estimates equation (1) using the extended sample through December 2022. The specification uses the aggregate orthogonalized monetary policy shock, MPS_t^{orth}. The remaining controls and fixed effects follow Table 2. Standard errors are clustered at the FOMC-date level, with the corresponding t-statistics reported in parentheses.

B Model derivation

B.1 Households' optimality conditions

Since the first-order condition for consumption gives $Y_{H,t} - C_{H,t} = 1$, saving is always one and households invest a wealth of one in home and foreign assets. The first-order condition for the portfolio share is

$$\frac{E_t e^{-\gamma r^{p,H}_{t+1} + s_{t+1} - s_t + i^*_t - \pi_{t+1}} - E_t e^{-\gamma r^{p,H}_{t+1} + i_t - \pi_{t+1}}}{E_t e^{(1-\gamma) r^{p,H}_{t+1}}} - \psi(z_{H,t} - z_{H,t-1}) = 0, \tag{B1}$$

where $\pi_{t+1} = p_{t+1} - p_t$ denotes inflation from time t to $t+1$ and p_t is the log price level. Linearizing around zero values of exponents and evaluating the expectations of exponentials by assuming log normality, the following relationship can be derived:

$$E_t s_{t+1} - s_t + i^*_t - i_t + 0.5 var_t(s_{t+1}) - cov_t(\gamma r^{p,H}_{t+1} + p_{t+1}, s_{t+1}) - \psi(z_{H,t} - z_{H,t-1}) = 0. \tag{B2}$$

A first-order approximation of the log portfolio return can be expressed as

$$r^{p,H}_{t+1} = z_{H,t} er_{t+1} + i_t - \pi_{t+1}, \tag{B3}$$

where the excess returns present $er_{t+1} = s_{t+1} - s_t + i^*_t - i_t$. Substituting (B3) into (B2), the optimal portfolio can be expressed as

$$z_{H,t} - \bar{z}_H = \frac{\psi}{\psi + \gamma\sigma^2}(z_{H,t-1} - \bar{z}_H) + \frac{1}{\psi + \gamma\sigma^2} E_t er_{t+1} \tag{B4}$$

where $\sigma^2 = var_t(s_{t+1})$.[12]

[12] From (B2), the steady-state fraction of asset invested in the Foreign bond by Home agents, $\bar{z}_H$ can be expressed as $\bar{z}_H = \frac{1}{2\gamma} + \frac{\gamma-1}{\gamma}\frac{\sigma_{s,p}}{\sigma^2}$, where $\sigma_{s,p} = cov_t(s_{t+1}, p_{t+1})$. As agents shift from risk-neutral to risk-averse behavior, second moments influence allocation decisions. Specifically, increased exchange rate volatility leads agents to reduce foreign investment. Conversely, when inflation and exchange rates exhibit

B.2 Bond market clearing condition

Due to symmetry between home and foreign households, $\bar{z}_F = 1 - \bar{z}_H$ holds. The real supply of bonds is fixed at one in terms of the purchasing power of the respective countries. Let P_t^* be the consumer price index of the foreign country measured in the foreign currency and $Q_t = S_t P_t^* / P_t$ be the real exchange rate. Due to Walras' law, it is sufficient to consider the foreign bond market equilibrium. The value of the foreign bond supply in terms of home purchasing power is Q_t, while the wealth of home and foreign households in terms of home purchasing power is 1 and Q_t, respectively. Then, foreign bond market equilibrium can be expressed as

$$z_{H,t} + z_{F,t} Q_t = Q_t. \tag{B5}$$

Linearizing around the steady-state log real exchange rate, which is zero, the above equation becomes

$$z_t^A = 0.5 + b q_t \tag{B6}$$

where $z_t^A = 0.5(z_{H,t} + z_{F,t})$ is the average fraction invested in foreign bonds across Home and Foreign.

B.3 Proof of propositions

Proof of Proposition 1. Under the stability condition $D > 1$, the equilibrium real exchange rate admits the present-value representation

$$q_t \;=\; \alpha(\psi)\, q_{t-1} \;+\; E_t \sum_{i=0}^{\infty} \frac{1}{D(\psi)^{i+1}}\, r_{t+i}^D. \tag{B7}$$

The proof begins with the following lemma.

stronger co-movement, holding foreign assets helps stabilize purchasing power in the event of inflation, thereby increasing the proportion of assets invested abroad.

Lemma 1. *Let*

$$D(\psi) = \frac{\theta(\psi) + \sqrt{\Delta(\psi)}}{2}, \qquad \Delta(\psi) \equiv \theta(\psi)^2 - 4\psi b,$$

with

$$\theta(\psi) = 1 + \psi b + \gamma\sigma^2 b,$$

where $b > 0$, $\gamma > 0$, and $\sigma^2 > 0$. If the stable root satisfies $D(\psi) > 1$, then $D'(\psi) > 0$.

Proof. Differentiating $D(\psi)$ with respect to ψ yields

$$D'(\psi) = \frac{1}{2}\left(\theta'(\psi) + \frac{\Delta'(\psi)}{2\sqrt{\Delta(\psi)}}\right).$$

Since $\theta'(\psi) = b$ and $\Delta'(\psi) = 2\theta(\psi)\theta'(\psi) - 4b = 2b(\theta(\psi) - 2)$, we obtain

$$D'(\psi) = \frac{b}{2}\left(1 + \frac{\theta(\psi) - 2}{\sqrt{\Delta(\psi)}}\right) = \frac{b}{2}\left(1 - \frac{2 - \theta(\psi)}{\sqrt{\Delta(\psi)}}\right).$$

Now impose the stability condition $D(\psi) > 1$. By definition,

$$D(\psi) > 1 \quad \Longleftrightarrow \quad \frac{\theta(\psi) + \sqrt{\Delta(\psi)}}{2} > 1 \quad \Longleftrightarrow \quad \sqrt{\Delta(\psi)} > 2 - \theta(\psi).$$

If $2 - \theta(\psi) \leq 0$, then the term in parentheses is strictly larger than 1. If $2 - \theta(\psi) > 0$, the inequality $\sqrt{\Delta(\psi)} > 2 - \theta(\psi)$ implies

$$0 < \frac{2 - \theta(\psi)}{\sqrt{\Delta(\psi)}} < 1,$$

so the term in parentheses remains strictly positive.

In either case,

$$1 - \frac{2 - \theta(\psi)}{\sqrt{\Delta(\psi)}} > 0.$$

Since $b > 0$, it follows immediately that

$$D'(\psi) > 0.$$

$\square$

Fix any information set at time t and consider a perturbation to the (current or expected) real interest rate differential that weakly increases the path $\{E_t r^D_{t+i}\}_{i \geq 0}$, holding q_{t-1} fixed. The induced change in q_t is

$$\Delta q_t \;=\; \sum_{i=0}^{\infty} \frac{1}{D(\psi)^{\,i+1}}\, \Delta E_t\left[r^D_{t+i}\right], \tag{B8}$$

where $\Delta E_t[r^D_{t+i}]$ denotes the change in the conditional expectation at horizon i. For any $i \geq 0$, the weight $w_i(\psi) \equiv D(\psi)^{-(i+1)}$ is strictly decreasing in ψ:

$$w_i'(\psi) \;=\; -(i+1)\, D(\psi)^{-(i+2)}\, D'(\psi) \;<\; 0,$$

because $D(\psi) > 1$ and, by Lemma 1, $D'(\psi) > 0$.

Hence, for any perturbation satisfying $\Delta E_t[r^D_{t+i}] \geq 0$ for all i and $\Delta E_t[r^D_{t+j}] > 0$ for at least one j, we have $\Delta q_t > 0$ (a real depreciation), and moreover Δq_t is strictly decreasing in ψ:

$$\frac{\partial}{\partial \psi} \Delta q_t \;=\; \sum_{i=0}^{\infty} w_i'(\psi)\, \Delta E_t[r^D_{t+i}] \;<\; 0.$$

Equivalently, lowering the portfolio adjustment cost ψ raises the magnitude of the depreciation induced by an increase in the (current or expected) real interest rate differential.

The proposition follows directly from the above result and the assumption that $\psi'(d) < 0$, by applying the chain rule.

$\square$

Proof of Proposition 2. The expressions in (15) and (16) follow directly from substituting the respective shock paths into the present-value representation (14) and summing geometric series.

Below, we prove that the impact ratio $R(D) \equiv q_0^{\text{news}}/q_0^{\text{spot}}$ is globally decreasing in D for all $D > 1$.

Let

$$w_j(D) \equiv D^{-(j+1)}\rho^j, \qquad j \geq 0,$$

and define the normalized weights

$$p_j(D) \equiv \frac{w_j(D)}{\sum_{k=0}^{\infty} w_k(D)}.$$

Since $w_j(D) > 0$ and $\sum_{j\geq 0} w_j(D) < \infty$ for $D > 1$ and $0 < \rho < 1$, we have $p_j(D) > 0$ and $\sum_{j\geq 0} p_j(D) = 1$.

A direct calculation yields the closed form

$$p_j(D) = (1-\lambda)\lambda^j, \qquad \lambda \equiv \rho/D \in (0,1). \tag{B9}$$

Hence, changes in D map one-to-one into changes in λ: if $D_1 < D_2$, then $\lambda_1 > \lambda_2$.

Under the spot shock, the real interest rate differential follows

$$\Delta r_j^D = \kappa_{\text{spot}}\rho^j, \qquad j \geq 0.$$

Under the news shock with a plateau of length $M \geq 1$ and common decay rate ρ,

$$\Delta\mathbb{E}_0[r_j^D] = \begin{cases} 0, & j = 0, \\ \kappa_{\text{news}}, & 1 \leq j \leq M, \\ \kappa_{\text{news}}\rho^{j-M}, & j \geq M+1. \end{cases}$$

The impact response of the exchange rate satisfies

$$q_0 = \sum_{j=0}^{\infty} D^{-(j+1)} \Delta \mathbb{E}_0[r_j^D].$$

Hence,

$$q_0^{\text{spot}} = \kappa_{\text{spot}} \sum_{j=0}^{\infty} D^{-(j+1)} \rho^j, \qquad q_0^{\text{news}} = \kappa_{\text{news}} \sum_{j=0}^{\infty} D^{-(j+1)} b_j,$$

where the sequence $\{b_j\}_{j\geq 0}$ is defined by

$$b_0 = 0, \qquad b_j = 1 \ (1 \leq j \leq M), \qquad b_j = \rho^{j-M} \ (j \geq M+1).$$

Define the auxiliary sequence $\{a_j\}_{j\geq 0}$ by

$$a_0 \equiv 0, \qquad a_j \equiv \frac{b_j}{\rho^j} = \begin{cases} \rho^{-j}, & 1 \leq j \leq M, \\ \rho^{-M}, & j \geq M+1. \end{cases}$$

By construction, $b_j = \rho^j a_j$ for all $j \geq 0$. Substituting yields

$$q_0^{\text{news}} = \kappa_{\text{news}} \sum_{j=0}^{\infty} D^{-(j+1)} \rho^j a_j.$$

Given the common weight sequence

$$w_j(D) \equiv D^{-(j+1)} \rho^j,$$

The impact responses can be expressed as

$$q_0^{\text{spot}} = \kappa_{\text{spot}} \sum_{j=0}^{\infty} w_j(D), \qquad q_0^{\text{news}} = \kappa_{\text{news}} \sum_{j=0}^{\infty} w_j(D)\, a_j.$$

The ratio of impact responses can therefore be written as

$$R(D) \equiv \frac{q_0^{\text{news}}}{q_0^{\text{spot}}} = \frac{\kappa_{\text{news}}}{\kappa_{\text{spot}}} \frac{\sum_{j=0}^{\infty} w_j(D)\, a_j}{\sum_{j=0}^{\infty} w_j(D)}.$$

Using normalized weights $p_j(D)$, we obtain

$$R(D) = \frac{\kappa_{\text{news}}}{\kappa_{\text{spot}}} \sum_{j=0}^{\infty} p_j(D)\, a_j. \tag{B10}$$

Thus, up to the scale factor $\kappa_{\text{news}}/\kappa_{\text{spot}}$, the impact ratio is a weighted average of the nondecreasing sequence $\{a_j\}$.

Define the tail sums

$$P_D(\geq j) \equiv \sum_{k=j}^{\infty} p_k(D).$$

From (B9),

$$P_D(\geq j) = \sum_{k=j}^{\infty} (1-\lambda)\lambda^k = \lambda^j. \tag{B11}$$

Therefore, if $D_1 < D_2$ (equivalently $\lambda_1 > \lambda_2$), then for every $j \geq 0$,

$$P_{D_1}(\geq j) = \lambda_1^j \;>\; \lambda_2^j = P_{D_2}(\geq j). \tag{B12}$$

That is, lowering D shifts weight toward larger horizons in the sense of increasing all tail sums.

We now use a standard summation-by-parts identity. For any sequence $\{a_j\}_{j\geq 0}$ with $a_{-1} \equiv 0$,

$$\sum_{j=0}^{\infty} p_j(D)\, a_j = \sum_{j=0}^{\infty} \left(a_j - a_{j-1}\right) P_D(\geq j). \tag{B13}$$

To verify (B13), expand the right-hand side and note that the coefficient on each a_j telescopes to $p_j(D)$.

Because a_j is nondecreasing, we have $\Delta a_j \equiv a_j - a_{j-1} \geq 0$ for all $j \geq 0$. Combining

(B12) and (B13), for any $D_1 < D_2$,

$$\sum_{j=0}^{\infty} p_j(D_1)\, a_j - \sum_{j=0}^{\infty} p_j(D_2)\, a_j = \sum_{j=0}^{\infty} \Delta a_j \Big(P_{D_1}(\geq j) - P_{D_2}(\geq j) \Big) > 0,$$

since each $\Delta a_j \geq 0$ and each tail difference is strictly positive by (B12). The final inequality is strict because $\Delta a_1 = \rho^{-1} > 0$ and $P_{D_1}(\geq 1) - P_{D_2}(\geq 1) = \rho/D_1 - \rho/D_2 > 0$. Hence,

$$D_1 < D_2 \quad \Longrightarrow \quad \sum_{j=0}^{\infty} p_j(D_1)\, a_j > \sum_{j=0}^{\infty} p_j(D_2)\, a_j. \tag{B14}$$

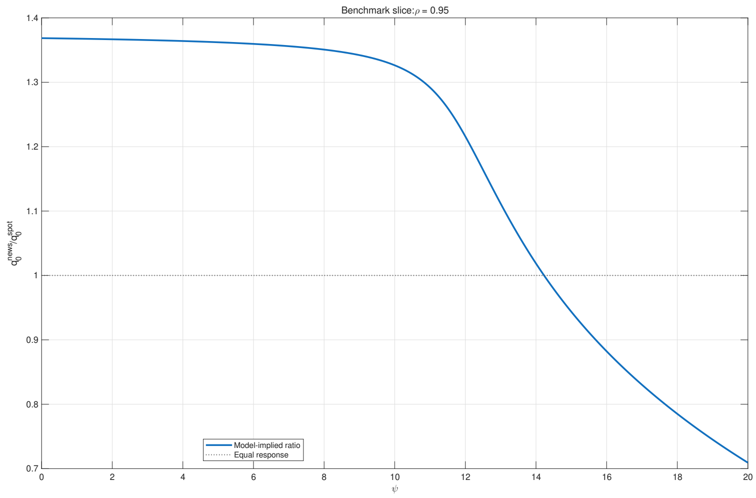
Finally, multiplying (B14) by the positive constant $\kappa_{\text{news}}/\kappa_{\text{spot}}$ in (B10), we obtain

$$D_1 < D_2 \quad \Longrightarrow \quad R(D_1) > R(D_2).$$

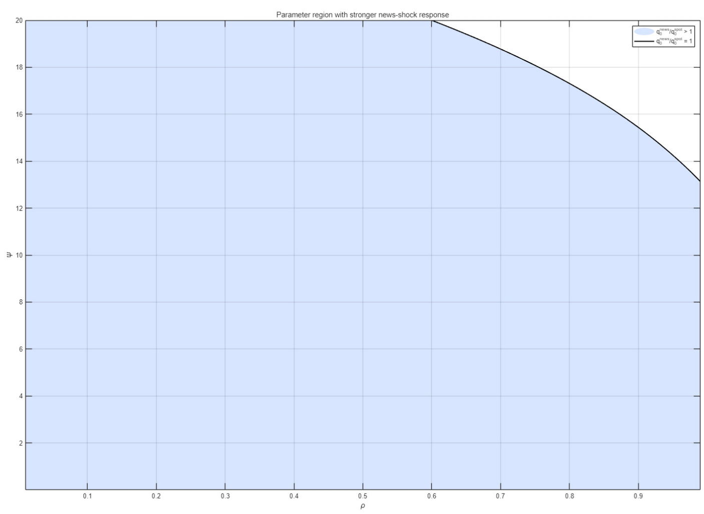
Therefore, $R(D)$ is strictly decreasing in D for all $D > 1$. Since deeper markets correspond to lower values of D, the ratio $q_0^{\text{news}}/q_0^{\text{spot}}$ is strictly increasing in market depth.

□

Figure B3: Illustrative sensitivity analysis of news and spot responses



Panel A. Benchmark slice at $\rho = 0.95$



Panel B. Parameter region with $q_0^{news} > q_0^{spot}$

Note: Panel A plots the model-implied relative impact response, q_0^{news}/q_0^{spot}, as the adjustment-cost parameter ψ varies under the benchmark persistence calibration $\rho = 0.95$. Panel B reports the region of the admissible (ρ, ψ) parameter space in which the model-implied impact response to a news shock exceeds that to a spot shock. Unless otherwise varied, the remaining model primitives are fixed at the benchmark values in Bacchetta and Van Wincoop (2021): $\sigma = 0.027$, $b = 0.085$, and $\gamma = 50$. The figures are intended as qualitative sensitivity exercises rather than quantitative IRF-matching exercises.