Do Central Banks Follow the Fed?*

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Abstract

Using a newly constructed dataset covering over 9,500 monetary policy meetings across 59 economies, we show that emerging market economy (EME) central banks systematically move in unison with the Fed, whereas advanced economy central banks do not. We present evidence that this heterogeneity arises because, in EMEs, U.S. monetary shocks are perceived as inflationary, transmitted through exchange rate fluctuations that affect import prices and firms' costs, owing to the greater prevalence of dollar-invoiced imports and dollar debt. A three-bloc open-economy New Keynesian model with dominant currency pricing and partial debt dollarization rationalizes these empirical patterns. The model highlights that EME responses stem from the immediate impact of Fed shocks on exchange rates and, in turn, on inflation.

JEL Classification: E58, F31, F41

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"Whatever changes in the U.S. rates, [we] will follow and do the same."

—Anonymous emerging market central banker, Huertas (2022).

"There is a view that in matters of monetary policy, the Reserve Bank is guided by the principle of 'follow the Fed'. I would like to unambiguously state that while we do keep a watch on whether clouds are building up or clearing out in the distant horizon, we play the game according to the local weather and pitch conditions."

—Governor of the Reserve Bank of India, June 2024.

1 Introduction

The decisions of the United States Federal Reserve System (the Fed) have sizable spillovers on the global economy, including effects on variables that central banks monitor on an ongoing basis, such as exchange rates, prices, and economic activity. This naturally raises a set of questions: Do central bankers react to U.S. monetary decisions? If so, when, to what extent, and why? As suggested by the contrasting views above, answering these questions is important because, if central banks do respond to the Fed, a broader issue arises about whether they do so mechanically or in response to the effects of U.S. monetary policy on domestic conditions.

To answer these questions, we construct a new dataset on monetary policy meetings and show that emerging market economy (EME) central banks move in tandem with the Fed, whereas advanced economy (AE) central banks do not. We demonstrate that the difference in central bank responses across these two groups of economies stems from heterogeneity in the revision of inflation expectations following a U.S. monetary shock and in the degree to which an economy's inflation is exposed to such shocks. Finally, we demonstrate that a three-bloc open economy New Keynesian model with multiple currency pricing and partial debt dollarization reproduces the empirical patterns we document. The model highlights that the response of EME central banks to a U.S. monetary shock is rooted in the immediate effect of such shocks on the nominal exchange rate and, therefore, on inflation.

We begin our analysis by introducing the Monetary Policy Meetings dataset, a new dataset that records the monetary policy decisions in over 9,500 monetary policy meetings between August 1990 and December 2024 by fifty-nine central banks, including the Fed.[1] One key advantage of the MPM-dataset over alternative policy rate datasets widely used in the literature (e.g., the IMF's *International Financial Statistics*) is that it records the exact dates of monetary policy meetings for an extensive set of countries, a feature that we use to study two margins

[1] In this paper, we use the words *meetings* and *policy meetings*, as synonyms for *monetary policy meetings*.

through which central banks may follow the Fed: the timing of policy meetings and policy rate decisions.

First, we document that central banks meet on average nine times per year to decide on their monetary stance and find that the timing of their meetings heavily bunches in the days immediately following FOMC meetings. Specifically, a central bank's first policy meeting after an FOMC announcement occurs within 14 days in 50–60% of cases. This pattern is unlikely to arise by chance: under random uniform scheduling, the corresponding figure would be close to 30%. Furthermore, we show that this bunching pattern is similar for EMEs and AEs and is not driven by central banks in specific regions, income groups, or by the release of quarterly macroeconomic data.

Second, we demonstrate that EME central banks follow the Fed's policy rate decisions, but AE central banks do not. Specifically, we use an instrumental variable (IV) regression to show that an exogenous 100 basis points increase in the Fed funds rate leads EME central banks to increase their policy rate by 50 basis points in their first meeting after the Fed announces their policy decision whereas AE central banks keep their policy rate unchanged. We complement this finding with local projection regressions à la Jordà (2005, 2023) to study the dynamic response of policy rates to a U.S. monetary shock. Our estimates reveal that EMEs policy rates are kept above the trend level for at least a year, whereas AE policy rates do not deviate from their trend.

The meeting-level dataset is crucial component for this empirical result. By contrast, if we aggregate the data at a monthly frequency (similar to existing datasets from the IMF or BIS) and approximate the effect of the Fed on central bank decisions by regressing policy rate changes on the Fed fund rate changes, we obtain a statistically insignificant coefficient that is roughly two-thirds as large as our baseline estimate —consistent with previous estimates in the literature (e.g., Camara, Christiano, & Dalgic, 2024; Degasperi, Hong, & Ricco, 2024; De Leo, Gopinath, & Kalemli-Özcan, 2024). The reason is that when the FOMC meeting occurs later in the month, central banks are less likely to meet afterward within that same month. As a result, monthly data blur the contemporaneous relationship. We resolve this ambiguity by using a more precise empirical setup that underscores the importance of precisely identifying policy meeting dates to properly capture the influence of Fed decisions on central bank decisions.[2]

The facts we document are robust across several dimensions, and we find that the Fed is

[2]Pelin (2025) provides direct evidence for a Fed information effect by showing that FOMC announcements tend to reveal private signals about industrial production, which then influence financial markets and macroeconomic expectations. Our approach is broadly in the same spirit: we exploit the timing of policy meetings to isolate how the Fed's decisions causally affect other central banks' decisions.

the only central bank at the epicenter of other central banks' decisions. Specifically, we verify that our results are robust to using alternative monetary shock series as instruments —rather than the Bauer and Swanson (2023b) shocks employed in our baseline results— as well as to alternative classifications of emerging economies. Moreover, we show that central banks do not follow the European Central Bank and that the Fed does not follow any other central bank. For this reason, we focus on understanding why EME central banks choose to follow the Fed.

Having established that EME central banks follow the Fed while AE central banks do not, we examine factors that could explain central bank responses. More specifically, we study the effect of a Fed tightening on the expectations and realized values of key macroeconomic variables monitored by central banks. We show that a 100 basis points Fed tightening is followed by a 40 basis points increase in inflation expectations and a 100 basis points rise in realized inflation in EMEs, whereas these variables are not significantly affected in AEs. Interestingly, these differences arise despite the nominal exchange rate depreciating by roughly 3% in both groups of economies. Moreover, we find that growth expectations in EMEs decline by about 20 basis points.

We show that, consistent with the behavior of a credible central bank focused on its inflation target, central banks in economies where inflation is more exposed to U.S. monetary shocks through movements in the nominal exchange rate respond more strongly to such shocks.[3] Specifically, we regress policy rate changes on the interaction between U.S. monetary shocks and three measures of inflation exposure: the degree of financial dollarization, the imports-to-GDP ratio, and the ratio of imports invoiced in dollars to GDP. Across all three measures, we obtain positive coefficients. These results indicate that EME central banks are not following the Fed mechanically: prices in EMEs are more sensitive to U.S. monetary shocks, providing EME central banks with a rationale to adjust policy in the same direction.

To understand how much of the response of EME central banks to a Fed tightening can be explained by its effects on firms' marginal costs, we propose a three-bloc open economy New Keynesian model with multiple currency pricing and firm dollar debt.

Our model replicates the observed response of central banks following a Fed tightening and emphasizes that, in EMEs, the imports prices are more exposed to U.S. monetary shocks due to dollar invoicing. In addition, firms in EMEs face a working-capital constraint that requires them to borrow, partly in dollars, to prepay for production inputs. These features lead to higher import prices and increased marginal costs after a Fed tightening, generating upward pressure on domestic prices and, consequently, on CPI inflation. As a result, an inflation-

[3]Throughout the paper, we define the nominal exchange rate as the number of units of local currency required to purchase one U.S. dollar. Thus, a depreciation of the local currency means that more local currency units are needed to buy one dollar.

targeting EME central bank responds by raising its policy rate. In particular, following a 100 basis points increase in the Fed funds rate, the EME central bank raises its policy rate by roughly 50 basis points, whereas the AE central bank response is close to zero, at barely 10 basis points. This difference arises because CPI inflation and firms' marginal costs increase sharply in EMEs, so the EME central banks response is stronger the higher the degree of debt dollarization and dollar invoicing.

Related Literature. This paper contributes to three main strands of the literature. First, we contribute to the literature examining the effects of U.S. monetary shocks on the global economy (Rey, 2015; Dedola, Rivolta, and Stracca, 2017; Vicondoa, 2019; Miranda-Agrippino and Rey, 2020, 2022; Huertas, 2022; Camara et al., 2024; Degasperi et al., 2024; De Leo et al., 2024; Arbatli-Saxegaard, Firat, Furceri, and Verrier, 2025; Camara, 2025). While this literature has not focused directly on central bank decisions at meetings following FOMC announcements, some studies have analyzed how policy rates at a monthly frequency respond to monetary shocks (e.g., Camara et al., 2024; Degasperi et al., 2024). We show that the absence of meeting-level data limits the ability of these approaches to identify the immediate responses of central banks to U.S. monetary shocks. Relative to these studies, we construct a new high-frequency dataset of monetary policy meetings (in the spirit of Cook and Hahn, 1989; Kuttner, 2001) and provide evidence that EME central banks adjust their policy rates in the same direction as U.S. monetary shocks.

Second, we contribute to the literature on monetary policy decision-making. A large body of work, beginning with Taylor (1993), has noted that policy rates can be described as linear combinations of macroeconomic fundamentals (see for instance Ball, 1999; Leitemo and Söderström, 2005; Galí and Monacelli, 2005; Davig and Leeper, 2007; and Faia and Monacelli, 2008). Using meeting-frequency data, we contribute to this literature by showing that the Fed plays a central role in the policy decisions of EME central banks. Moreover, we provide direct evidence that these decisions are consistent with attempts to prevent an inflation-expectations de-anchoring; an objective recently highlighted by Nakamura, Riblier, and Steinsson (2025) and Kakhbod, Kermani, and Maciel (2025) in the case of the Fed.

Finally, we connect our findings to the recent literature employing New Keynesian models with dominant currency pricing (e.g., Gopinath et al., 2020; Gopinath and Itskhoki, 2022; Akinci and Queralto, 2023; Akinci, Benigno, Pelin, and Turek, 2024; Cabezas, Candia, and Xu, 2024; McLeay and Tenreyro, 2025; Ahmed, Akinci, and Queralto, in press). We extend this class of models by incorporating firms that borrow in dollars, and show that the framework is able to replicate the empirical patterns we document. When exports are invoiced in dollars, exchange rates cannot stabilize economic activity, as the classical expenditure-switching

channel is muted. A large literature has demonstrated that exchange-rate pass-through into destination prices and quantities depends critically on the invoicing currency, with a substantial share of world trade denominated in U.S. dollars (Goldberg & Tille, 2008; Gopinath, 2015; Gopinath et al., 2020). In addition, foreign-currency borrowing by EME firms can render currency depreciation contractionary through negative balance sheet effects and higher marginal costs (Aguiar, 2005; Bleakley & Cowan, 2008; Bruno & Shin, 2015; Céspedes, Chang, & Velasco, 2004; Kalemli-Ozcan, Kamil, & Villegas-Sanchez, 2016). We therefore consider both of these roles of the U.S. dollar as potential channels amplifying the effects of U.S. monetary policy shocks on EMEs.

Layout. The paper is organized as follows. Section 2 discusses the data. Section 3 presents our main empirical result: EME central banks move in tandem with the Fed, AE central banks do not. Section 4 advances a hypothesis rationalizing why central banks follow the Fed and provides evidence in its support. Section 5 shows that, by incorporating dollar-denominated debt for firms into a model with dominant currency pricing, this class of models can replicate the empirical patterns we document. Section 6 concludes.

2 A New Database of Monetary Policy Meetings

To analyze how the Federal Reserve's monetary policy decisions affect the policy decisions of other central banks, it is ideal to rely on data that satisfy three properties. First, the data must record every monetary policy decision, including multiple decisions within the same month, to ensure an accurate measurement of policy decisions. Second, the data must allow for the identification of the set of macroeconomic information available at the time central banks make their policy decisions. Third, the data must be representative of a broad group of countries whose monetary policy frameworks share similar objectives and instruments. A dataset meeting these criteria does not currently exist. In this section we introduce a novel dataset that satisfies these criteria.

2.1 The MPM-Dataset

We construct the *Monetary Policy Meetings* dataset (MPM-dataset), a novel dataset that combines information on 9,511 monetary policy meetings —recorded on the exact day they took place— with macroeconomic data reflecting the information set available to policymakers at the time of their decisions. The sample is an unbalanced panel of policy rate decisions from 50 emerging market economies (EMEs) and 10 advanced economies (AEs), recorded on the

day the decisions were made. We restrict attention to periods in which central banks used a short-term interest rate as their primary policy instrument, excluding episodes in which they targeted monetary aggregates or pegged their currency to that of another country.[4]

The AEs in the MPM-dataset are Australia, Canada, the European Union, Japan, New Zealand, Norway, Sweden, Switzerland, the United Kingdom, and the United States. All other economies are classified as EMEs. This set of AEs closely aligns with those considered in the academic literature (e.g. Camara et al., 2024) and by practitioners (e.g., the S&P Global Country Classification).

Regarding data sources, we obtain the meeting dates and the associated monetary policy decisions from press releases and monetary policy minutes available on central bank websites. The macroeconomic variables include the inflation rate and GDP growth observable on the day the decisions were made, as well as standard measures of monetary policy shocks for the Federal Reserve and the European Central Bank (ECB). Appendix A provides a detailed description of the country-period coverage for each variable, the data sources, and the monetary shocks accompanying the dataset.

Importantly, the Fed and ECB monetary shock series that we include in the MPM-dataset consist of high-frequency monetary policy surprises constructed from changes in short-term interest rate instruments within a narrow 30-minute window surrounding FOMC announcements. The identifying assumption is that all information that is public at the beginning of the announcement window has already been incorporated into financial markets. Consequently, variation in asset prices within this narrow window primarily reflects unexpected monetary policy news revealed by the FOMC announcement itself, rather than broader macroeconomic shocks that may trigger similar policy responses by both the Federal Reserve and non-U.S. central banks. In other words, using high-frequency monetary policy surprises ensures that the variation in unexpected monetary policy changes is not driven by broad events such as the terrorist attacks on 9/11/2001, the financial stress associated with the Global Financial Crisis, the COVID pandemic, or the war in Ukraine.

2.2 Advantages of the MPM-dataset

The MPM-dataset has three main advantages over commonly used sources of data on policy rates such as the IFS or the Bank for International Settlements' Data Portal (BIS-DP).

The first advantage of the MPM-dataset is that it records policy rates only during periods in

[4]For example, between March 19, 2001 and March 8, 2006, and again between April 3, 2013 and September 20, 2016, the Bank of Japan used the outstanding balance of current accounts at the Bank of Japan as its policy instrument. We exclude meetings during these periods from the dataset.

which a well-defined policy rate exists. Hence, it excludes episodes where the implied policy rate is highly sensitive to asset-demand fluctuations.[5] In this regard, the MPM-dataset is not affected by the measurement errors that arise in standard datasets.

Second, unlike standard datasets, the MPM-dataset explicitly records the exact dates of policy meetings. When standard datasets are used to study the behavior of policy rates, researchers must adopt one of two approaches: *i)* assume that decisions were made only on the dates when the policy rate changed; or *ii)* aggregate the data to a given frequency (typically monthly) and assume that if the policy rate does not change from one period to the next, then a zero change was decided. The first approach assumes that central banks never decide to keep the policy rate unchanged, whereas the second introduces an excessive imputation of zero decisions. As will become clear when we discuss the summary statistics of our data, on average the first approach ignores information from two-thirds of policy meetings, while the second assumes there are policy decisions every month, which is not the case in the data. By incorporating the full set of policy meetings, the MPM-dataset eliminates the need to rely on either approach and therefore avoids artificially imputing or excluding policy decisions.

The third and final advantage of the MPM-dataset is that it dominates existing datasets in terms of both country coverage and frequency. Relative to the BIS-DP, which provides daily policy rates for 34 countries with an explicit rate in 2024, the MPM-dataset includes 25 additional countries. In contrast, while the IFS offers broad country coverage, it reports policy rates only at a monthly frequency. This limitation makes it difficult to study how central bankers react to unexpected events, since researchers must impose stronger assumptions about the joint structure of multiple shocks. By recording policy decisions at a meeting frequency, the MPM-dataset overcomes this problem.

2.3 Summary Statistics

Table 1 reports the mean, median, and standard deviation of the main macroeconomic indicators in the dataset, separately for EMEs and AEs.Consistent with common wisdom, inflation and nominal policy rates are higher and more volatile in EMEs than in AEs over the sample period. In addition, consistent with the findings of recent papers (e.g. Mendoza-Fernández & Meyer, 2025), we find that AEs grow at a slower pace than EMEs.

An interesting pattern in Table 1 is that the average number of policy meetings is close to nine per year for both EMEs and AEs —a figure similar to the eight meetings typically held

[5] As an example, before 2008 Mexico did not have an explicit policy rate, and the implied rate could swing sharply from one day to the next: it was 8.2% on January 14, 2003, dropped to 6.1% on January 15, and then returned to 8.1% on January 16.

Table 1. Summary Statistics

	Advanced economies			Emerging economies		
	Mean	Median	S.D.	Mean	Median	S.D.
Macroeconomic indicators:						
Inflation	1.9	2.2	0.8	6.6	4.6	7.7
GDP growth	2.3	1.9	1.1	3.2	3.3	1.7
Monetary policy rate	2.3	2.4	1.2	7.6	5.8	6.4
Monetary meetings per year	9.1	10.6	2.3	9.0	8.5	3.2
Share of first meetings after an FOMC announcement:						
0-14 days after	51.9	50.5	20.6	58.0	59.5	16.9
15+ days after	48.1	49.5	20.6	42.0	40.5	16.9

Notes: This table reports the summary statistics of the main variables in our dataset. We first compute the average of each quantity indicated in the rows, first column, of this table for each country, and then we take the mean, median, and standard deviation (S.D.) of such averages across countries to obtain the figures shown in this table.

by the Federal Reserve in a normal year. Moreover, in 50 to 60% of cases, a central bank's first policy meeting after an FOMC meeting takes place within 14 days, suggesting that many central banks strategically schedule their meetings shortly after U.S. policy announcements.

3 Do Central Banks Follow the Fed?

In this section, we use the MPM-dataset to study two margins through which the Fed may influence central banks' behavior. First, we examine the timing margin —how monetary policy meetings are scheduled relative to FOMC announcements— and find that both EME and AE central banks bunch a large share of their meetings in the days immediately following an FOMC announcement. Second, we analyze the quantitative margin and show that EME central banks follow the Fed in terms of their policy decision, whereas AE central banks do not.

3.1 Bunching of Meetings After FOMC Announcements

In Table 1 we show that non-U.S. central banks schedule close to nine meetings annually. In a typical year, the FOMC holds eight meetings per year, which is equivalent to scheduling one meeting every 40 to 45 days. Hence, if non-U.S. monetary policy meetings were scheduled at random, only about one-third of them would fall within 14 days after an FOMC

announcement.[6] Strikingly, Table 1 shows this is not the case. A central bank's first policy meeting following an FOMC meeting falls within 14 days after an FOMC meeting in 50–60% of cases, suggesting that central banks deliberately bunch their policy meetings immediately after FOMC meetings. In what follows, we restrict attention to central banks' first policy meetings after an FOMC announcement and provide additional descriptive evidence in support of this point.

Figure 1 presents graphical evidence supporting the bunching of policy meetings after the FOMC announcements. Panel A shows the histogram of the number of days between a local policy meeting and the closest FOMC announcement. The figure shows that the number of central bank meetings that happen in the week following an FOMC announcement is approximately four times larger than in the week preceding it.

Panel B complements by showing the histogram of the number of days between an FOMC announcement and the first policy meeting of a central bank after such an announcement, truncated at 49 days. This panel reveals that the number of policy decisions made within the first week following an FOMC announcement is roughly twice the number of policy decisions made in each of the subsequent weeks, indicating that a disproportionately large share of policy meetings occur either on the same day as the FOMC announcement or within the following seven days.

Importantly, the bunching pattern is prevalent across several dimensions. In Appendix B.1 we show that results hold for the modal central bank in each of the five regions considered in this study (Figures B2-B6), by income level (Figures B7-B8) and type of economy (Figures B9-B10). Additionally, Figure B11 shows that the same pattern holds after excluding meetings occurring between three and seven weeks after the end of a quarter, which are typically the dates when GDP figures are released, meaning that the bunching pattern is not driven by dates when economic activity data are released. Lastly, Figure B12 shows that for forty-five out of the fifty-eight non-U.S. central banks (i.e., 80%), the first meeting after an FOMC announcement takes place within 14 days after the announcement. The pattern also applies to the People's Bank of China, the Bank of England, the Bank of Japan, and the European Central Bank —four central banks of great interest in international macroeconomics (e.g., in Miranda-Agrippino & Rey, 2020, 2022).

Taken together, these findings indicate that a majority of central banks systematically schedule their policy meetings after U.S. monetary policy announcements, regardless of geography, income level, market type, or the timing of macroeconomic data releases. We there-

[6] This calculation is straightforward: the Fed typically spaces its meetings 40-45 days apart. If a non-U.S. central bank also holds about eight meetings per year, scheduled uniformly at random, then the chance that its next meeting falls within the 14-day window following an FOMC meeting is approximately $14/42 = 1/3$.

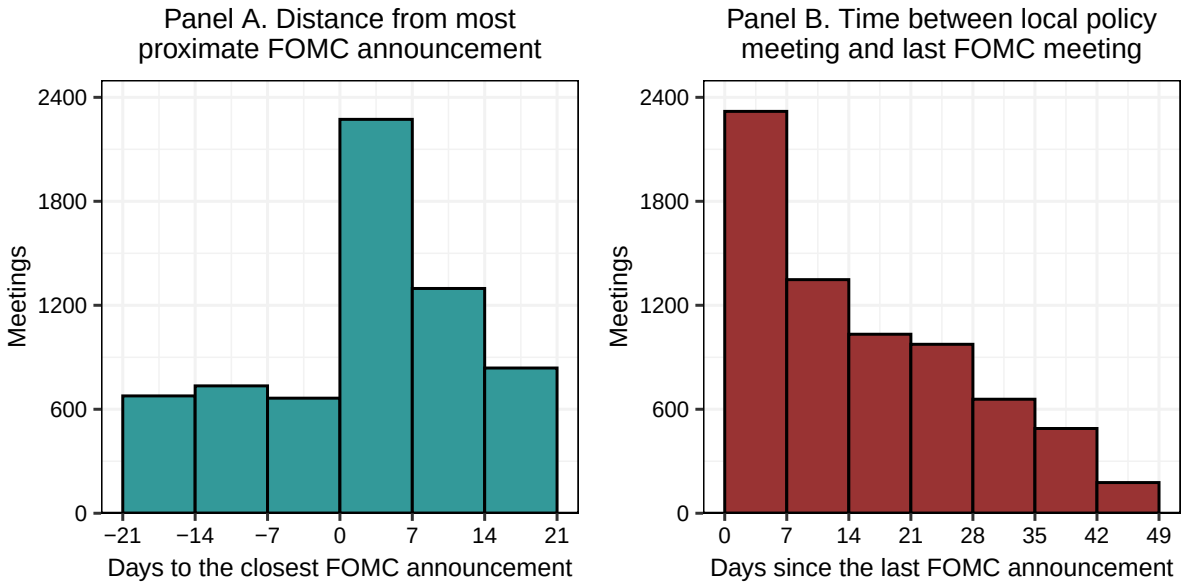


Figure 1. Bunching of Policy Meetings

Notes: This figure shows histograms of the the number of days between a policy meeting and the closest FOMC announcement (Panel A) and the number of days between an FOMC announcement and the next policy meeting of a central bank (Panel B). Panel A restricts the sample to meetings that occurred within three weeks before or after an FOMC announcement, while Panel B includes all meetings.

fore interpret this pattern as evidence that central banks take into account the Fed's policy calendar when scheduling their own policy meetings.[7]

Intuitively, there is a benefit to scheduling monetary policy meetings in the days following Fed announcements: if the Fed takes a decision that could put a central bank's inflation and output objectives at risk, it is desirable for the central bank to be able to react quickly in order to keep expectations about its medium and long run objectives anchored. Furthermore, in the case of emerging economies, there may be an additional credibility benefit, as economic agents may interpret the timing of their meetings relative to those of the Fed as a signal of commitment to maintaining macroeconomic stability in the face of foreign shocks affecting the domestic economy.

[7]The FOMC meeting schedule for a given year is published well in advance. Press releases available on the Fed's website since 2006 indicate that the latest such schedule has ever been published was in 2024, when the schedules for 2025 and 2026 were released on August 9. Moreover, other central banks that announce their policy meeting calendars for the following year (e.g., the central banks of Albania, Canada, Chile, and Costa Rica) typically publish their calendars on dates several weeks or months later than the Fed's announcement.

3.2 Immediate Adjustments After FOMC Announcements

What do central banks decide in the meetings that follow an FOMC announcement? Figure 2 shows that in their first monetary policy meeting after an FOMC announcement, central banks adjust their policy rates in the same direction as the Fed. In what follows, we demonstrate that this relationship is causal.

Specifically, we quantify the effect of a change in the Fed funds rate on the most proximate policy rate decision made by a central bank after the FOMC announcement. For this purpose, we estimate regression equation:

$$\Delta i_{c,\tau+\epsilon} = \alpha_c + \beta \cdot \Delta i_{\tau}^{US} + \gamma' \boldsymbol{X}_{c,\tau+\epsilon} + e_{c,\tau+\epsilon}, \tag{1}$$

where the outcome variable $\Delta i_{c,\tau+\epsilon}$ is the change in the policy rate of country c decided at the first meeting following FOMC announcement τ and before FOMC announcement $\tau+1$; the main explanatory variable Δi_{τ}^{US} measures the change in the Fed funds rate according to FOMC announcement τ; the α_c are country fixed effects; $\boldsymbol{X}_{c,\tau+\epsilon}$ is a vector including the most recent monthly change in annual CPI inflation and GDP growth observable at meeting $\tau+\epsilon$;[8] and $e_{c,\tau+\epsilon}$ is an error term. We interpret the coefficient β as the causal effect of Fed decisions on other central banks' policy rates.

Table 2 presents our baseline estimates, obtained using ordinary least squares (OLS) in Panel A and instrumental variable (IV) regression in Panel B. For the IV specifications, we instrument the explanatory variable Δi_{τ}^{US} using a monetary policy shock series identified through high-frequency methods, specifically we use the series from Bauer and Swanson (2023b).[9] Estimates are reported separately for the full sample, EMEs, and AEs, and are based on observations for which the policy rate is at most 20% (98% of the sample), excluding observations where the dependent variable falls below the 1st percentile or above the 99th percentile. We report standard errors à la Driscoll and Kraay (1998), which allow regression errors to be correlated across both central banks and time.

[8] The choice for our baseline results embodies the notion that central banks follow policy rules that respond to inflation and growth (Orphanides, 2003; Taylor, 1993). As we discuss at the end of this section, our results are robust to using alternative sets of controls.

[9] Bauer and Swanson construct their monetary shock series by estimating the first principal component of changes in short-term interest rate instruments within a 30-minute window surrounding FOMC announcements, and subsequently purging the resulting series from predictable responses to macroeconomic and financial conditions. The shocks are normalized such that a 100 basis points shock corresponds to a 100 basis points exogenous increase in the Fed funds rate. By relying on variation within a narrow announcement window, the identified shocks are plausibly exogenous and unlikely to reflect common information about macroeconomic or geopolitical events that may simultaneously affect the policy decisions of both the Federal Reserve and non-U.S. central banks. We download both the non-purged and purged shock series from Michael Bauer's website.

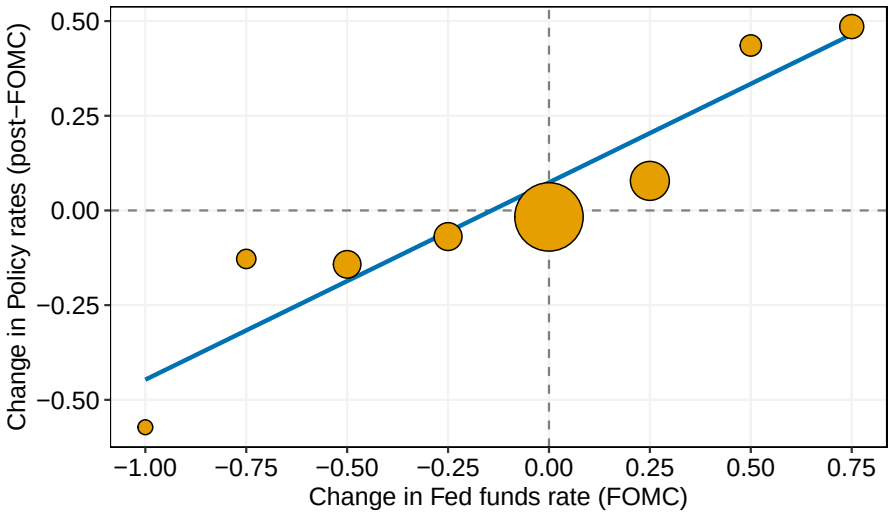


FIGURE 2. POLICY RATE DECISION SINCE LAST FOMC ANNOUNCEMENT

Notes: This figure shows the binscatter of the changes in the Fed funds rate against the changes in the policy rates of other central banks decided in the meetings following the FOMC announcement.

The OLS estimates of β indicate that the positive correlation between FOMC decisions and subsequent policy rate decisions by EME and AE central banks (shown in Figure 2) remains strong even after controlling for inflation and output growth. This result suggest that there is a non-trivial response of non-U.S. central banks to policy decisions by the Fed. However, the IV estimates indicate that this relationship can only be interpreted as causal for EMEs. For AEs, the IV estimate of β is statistically zero, meaning that the correlation Fed decisions and subsequent policy rate decisions by AE central banks likely reflects common macroeconomic shocks or shared economic conditions, rather than a causal response to exogenous U.S. monetary policy shocks.

Delving into the details, for EMEs the IV estimates of β are close to 0.5, implying that a 100 basis point increase in the Fed funds rate leads their central banks to tighten policy rates by about 50 basis points—stronger than the 29 basis points implied by the OLS estimates. For AEs, by contrast, the estimated β suggests an increase of roughly 13 basis points, but these estimates are not statistically significant at conventional confidence levels (90% or higher).

Our estimates confirm, with a high degree of statistical precision, that an unexpected increase in the Fed funds rate causes an increase in the policy rates of EMEs. Other recent studies such as Camara et al. (2024), Degasperi et al. (2024), De Leo et al. (2024), and Huertas (2022) find that a 100 basis points increase in the Fed funds rate is associated with an immediate rise of between 8 and 29 basis points in policy rates, but all of their estimates come with large standard errors, preventing them from establishing the sign of the relationship at confidence

Table 2. Policy Rates Decisions Following an FOMC Announcement

	Policy rate decision $\Delta i_{c,\tau+\epsilon}$		
	All	Emerging	Advanced
Panel A. OLS Estimates			
Δi_{τ}^{US}	0.319***	0.286***	0.429***
	(0.020)	(0.024)	(0.031)
Observations	6,730	5,178	1,552
R^2	0.142	0.129	0.261
Panel B. IV Estimates			
Δi_{τ}^{US}	0.430***	0.550***	0.135
	(0.157)	(0.210)	(0.304)
Observations	6,010	4,582	1,428
F-stat (1st stage)	186.4	135.5	48.5

Notes: The table reports the coefficient β of equation (1) estimated using ordinary least squares (Panel A) and using U.S. monetary policy shocks à la Bauer and Swanson (2023b) as instruments for the Fed funds rate (Panel B). The vector of control variables $\boldsymbol{X}$ includes country fixed effects, the most recent change in annual CPI inflation, and output growth. Driscoll and Kraay (1998) standard errors are reported in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

levels of 90% or higher —in all cases, the standard error exceeds 20 basis points.

In Table 3, we show that this difference in statistical significance can be attributed to the frequency of the data used in our estimates compared with that in the other studies. Specifically, when we regress changes in the end-of-the-month policy rate on changes in the end-of-the-month Fed funds rate, the coefficient β for EMEs falls to 0.39 with a standard error of 0.36.[10] . These figures are comparable to those reported in the studies cited above and emphasize that meetings frequency data are necessary to identify the causal effect of Fed decisions on non-U.S. central banks' policy rate decisions.

The intuition for why data aggregation alters the results can be illustrated with an example. If an FOMC meeting takes place in the last week of a month, the probability that an

[10]The regression we estimate using the MPM-dataset at monthly frequency is analogous to (1). That is, the regression includes the monthly version of the vector of controls in (1) and is restricted to periods when the policy rate is at most 20%, excluding observations for which the dependent variable falls below the 1st or above the 99th percentile. Including the outliers does not lead to statistically significant results.

Table 3. Contemporaneous Relation Between Policy Rates and the Fed Funds Rate

	Monthly policy rate change $\Delta i_{c,t}$		
	All	Emerging	Advanced
Δi_t^{US}	0.326	0.386	0.233
	(0.211)	(0.356)	(0.265)
Observations	11,314	8,643	2,671
F-stat (1st stage)	121.8	71.6	50.0

Notes: The table reports the coefficient β of the monthly version of equation (1) estimated using U.S. monetary policy shocks à la Bauer and Swanson (2023b) as instruments for the Fed funds rate. The vector of controls includes country fixed effects, the most recent change in annual CPI inflation, and output growth. Driscoll and Kraay (1998) standard errors are reported in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

EME central bank meets during the remainder of that month is roughly 33%.This probability increases if the FOMC meets in the last 15 days of the month, but it still does not reach 60%. Consequently, in monthly data there will appear to be many cases in which central banks do not react to Fed actions immediately, even though this is not actually the case.[11] Put simply, aggregation to monthly frequency biases the estimated coefficient from a regression of policy rates on the Fed funds rate toward zero.

In Appendix B.2.1 we confirm this intuition by estimating regressions of changes in end-of-the-month policy rates on changes in the end-of-the-month Fed funds rate using two sets of country-month observations: *i)* months in which the FOMC meeting occurred in the first 10 days; and *ii)* months in which it occurred in the last 10 days. For the former sub-sample we estimate a statistically significant β of 0.45 for EMEs (Panel A, Appendix Table B1), whereas for the latter sub-sample estimates of β are close to zero and far from statistically significant (Panel B, Appendix Table B1). This exercise indicates that using the full sample of monthly data attenuates and adds noise to β because it ignores that some central banks do not hold their first meeting after an FOMC announcement in the same month.

We conclude this subsection by addressing concerns about the validity of our empirical strategy. To this end, we re-estimate equation (1) while varying three aspects of the baseline procedure. First, we add lags of the policy decision to the vector of controls $\boldsymbol{X}_{c,\tau+\epsilon}$. Appendix Table B2 shows that including these lags results in an estimate of β that is barely different from our baseline and remains highly statistically significant. Second, the identifying assumption

[11] Since the Fed started pre-announcing their policy meetings in 1995, 67% of FOMC meetings were hold in the second half of a month.

behind our causal estimates is that the monetary shocks we use are truly exogenous to other central banks. Appendix Table B3 reports the results of a placebo test in which we estimate equation (1) using the last policy rate decision before the FOMC announcement, $\Delta i_{c,\tau-\epsilon}$, instead of the decision made afterward, $\Delta i_{c,\tau+\epsilon}$. In this case, we find no effect, suggesting that the information contained in the monetary shocks is not anticipated by other central banks —otherwise, it would be puzzling that EME central banks wait until the shock is realized to adjust their policy rates. Third, we use all of our sample, including country-meeting pairs for which the changes in the policy rate were below and above the 1st and 99th percentiles, respectively. Regardless of the inclusion of outliers, our main result continues to hold (see Appendix Table B4).

3.3 Dynamic Response of Policy Rates

We extend the analysis beyond the immediate policy decision that follows an FOMC announcement by examining how policy rates adjust over time to U.S. monetary surprises using a local projections framework à la Jordà (2005, 2023).

Specifically, we estimate the response of policy rates to a 100-basis-point increase in the Fed funds rate over a 360-day horizon using the following local projection regressions:

$$i_{c,\tau+h} = \alpha_c^h + \beta^h \cdot \text{MPS}_\tau^{US} + \boldsymbol{\gamma}^h \cdot \boldsymbol{X}'_{c,\tau} + e_{c,\tau+h}, \quad h = -240, ..., 360, \tag{2}$$

where $i_{c,\tau+h}$ denotes the policy rate of central bank c observed h days after the monetary policy surprise MPS_τ^{US} is revealed after FOMC meeting τ. The vector of controls $\boldsymbol{X}_{c,\tau}$ includes the most recent observables on annual inflation and GDP growth, twelve monthly lags of the local policy rate at time τ, and the Fed funds rate announced at the previous six FOMC meetings. We also estimate the regressions for $h < 0$ to test for pre-trends in central bank decision-making. For these cases, we follow the convention of lagging the vector of controls to $\boldsymbol{X}_{c,\tau-|h|}$.

The coefficients of interest are the β^h, which trace the cumulative change in a central bank's policy rate in response to an exogenous 100 basis point increase in the Fed funds rate. We estimate β^h using an ordinary least square estimation procedure proxying MPS_τ with the shocks identified by Bauer and Swanson (2023b). As in the previous subsection, we exclude country–meeting observations with policy rates above 20% and winsorize the dependent variables $i_{c,\tau+h}$ at the 1st and 99th percentiles. Additionally, for our baseline results we exclude observations $i_{c,\tau+h}$ for which the interval between τ and $\tau + h$ includes dates between July 2008–June 2009 or January 2020–December 2020. This choice reduces the influence of extreme

episodes such as the Global Financial Crisis (GFC) and the COVID pandemic. We discuss results under alternative data treatments after presenting the baseline estimates.

Figure 3 plots the estimated β^h coefficients from 240 days before to 360 days after an FOMC announcement, together with 68% and 90% confidence bands.[12] Results are shown separately for EMEs and AEs in Panels A and B, respectively. In both panels, the absence of systematic policy rate movements prior to FOMC announcements indicates that the estimates are not driven by pre-trends—supporting the validity of our identification strategy and confirming that the instrument captures unexpected variation in the Fed funds rate. Post-announcement responses reveal a meaningful difference: EME central banks tighten following a Fed hike, AE central banks do not.

In line with the estimates in Table 2, Panel A shows that EME policy rates increase by about 50 basis points within one month following a 100-basis-point surprise tightening in the Fed funds rate. Policy rates continue to rise for roughly eight months, reaching a peak increase of about 200-250 basis points before gradually returning to their pre-shock levels. Panel B indicates that policy rates in AEs move only modestly in tandem with the Fed, and these responses are not statistically significant at conventional levels over short horizons. Taken together, the estimates suggest that AE central banks do not adjust their policy rates promptly in response to a surprise Fed tightening, but may respond gradually over the medium run.

Previous studies provide inconclusive evidence on whether EME policy rates respond to U.S. monetary shocks. For example, Camara et al. (2024) and Degasperi et al. (2024) examine the effects of U.S. monetary shocks on a wide range of outcomes in EMEs and find no significant impact on policy rates at horizons between zero and 24 months. By contrast, our results provide clear evidence that following an exogenous increase in the Fed funds rate, EME central banks raise their policy rates above their pre-shock levels and keep them elevated for several months.

In Appendix B.3, we assess the sensitivity of our baseline results to three alternative specifications. First, we show that estimating equation (2) using the Fed funds rate instrumented with the SB monetary policy shocks yields results that are nearly identical to the baseline (see Figure B13). Second, we examine how including observations from the GFC and the onset of the COVID-19 pandemic affects our findings. When these periods are included, our main result for emerging market economies (EMEs) remains unchanged: EME central banks continue to raise policy rates for several months following a Fed tightening. In contrast, the estimates

[12] We use lag-augmented heteroskedasticity-corrected standard errors, following the recommendations of Montiel Olea, Plagborg-Møller, Qian, and Wolf (2025) for time-series analysis.

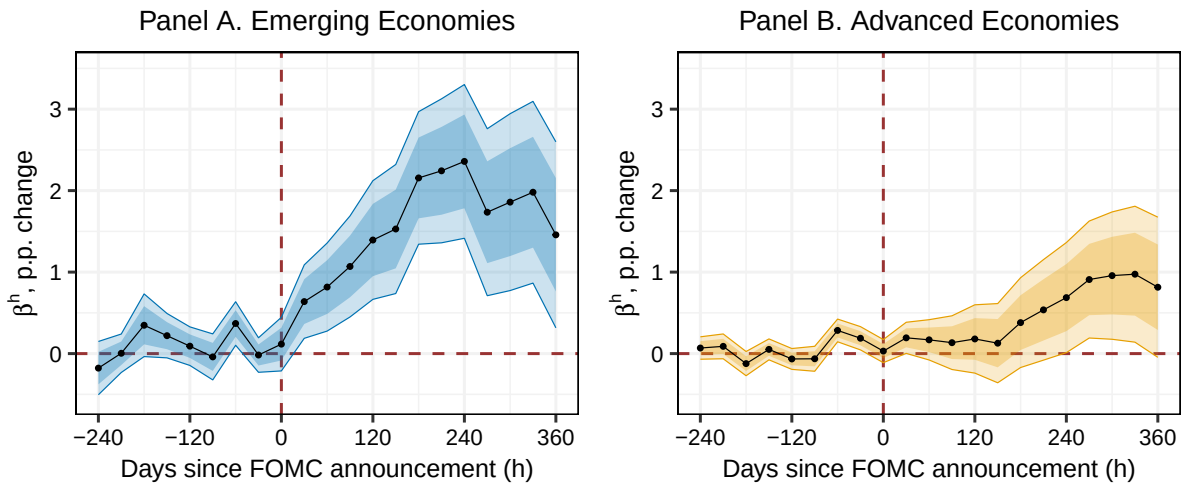


FIGURE 3. RESPONSE OF POLICY RATES TO A U.S. MONETARY SHOCK

Notes: This figure shows the cumulative response of central bank policy rates to U.S. monetary shocks for EMEs (Panel A) and AEs (Panel B). The shock series used to obtain these estimates is the Bauer and Swanson (2023b) monetary shock series. The light (dark) shaded areas represent the 90% (68%) symmetric confidence intervals, constructed using lag-augmented heteroskedasticity-corrected standard errors.

for advanced economies (AEs) differ from the baseline: when all periods are included, AE central banks appear to lean against the Fed (see Figure B14). However, as discussed in the Appendix, these differences are driven primarily by observations from the GFC alone (Figures B15–B16), an episode that arguably represents an extreme environment in which the Fed reached the zero lower bound (ZLB) while many other central banks still had room to lower their policy rates. Third, we exclude observations from the ZLB period (January 2009–December 2015) and find that the baseline results remain unchanged (see Figure B17).

3.4 Additional Results

We conclude this section by addressing three related questions that arise in our context: Do central banks follow other central banks besides the Fed? Does the Fed follow other central banks? How do our results relate to the post-covid tightening cycle? We briefly summarize the answers to these questions below.

Response to ECB monetary shocks. When news about the Fed's policy rate is released, central banks in other economies tend to revise their own policy rates at the meeting scheduled immediately afterward. A natural question is whether central banks from other major economies exert a similar influence. In Appendix B.4.1, we show that no such pattern ex-

ists with respect to news about ECB monetary policy decisions.[13] We briefly summarize the procedure and results here.

We apply the methodology from Section 3.2 to measure the average effect of an exogenous change in the ECB's Deposit Facility Rate (DFR) on policy rate decisions by non-E.U. central banks. Overall, we find no evidence that central banks revise their policy rates following a surprise change in the ECB's monetary stance. This result holds across all economies, whether the effect is estimated jointly or separately for EMEs and AEs (Table B5).

Fed's response to other central banks. In Appendix B.4.2, we use the *meeting-frequency* feature of our dataset to show that the Fed is unresponsive to the policy decisions of other central banks. For each central bank c, we estimate an IV regression of the change in the Fed funds rate at the FOMC meeting that follows a meeting of central bank c on the policy rate change decided by c in that meeting.[14] Our results indicate that 55 of the 58 estimated coefficients are not statistically significant at the 5% level (Appendix Figure B18), including those for central banks such as the ECB, the Bank of England, and the People's Bank of China. We therefore conclude that the Fed does not systematically follow any other central bank.

Post-Covid tightening cycle. During the post-Covid tightening cycle many EMEs have been known to tighten ahead of the Fed, see Nakamura et al. (2025). In a recent article, Mendoza-Fernández, Obstfeld, and Pelin (2026) analyze why and how EMEs have followed the Fed during this tightening cycle. We discuss in Appendix B.4.3 that this pattern is driven primarily by the expected federal funds rate and the U.S. dollar.

4 Why Emerging Economy Central Banks Follow the Fed

Why do EME central banks adjust policy rates in the same direction as U.S. monetary shocks whereas AE central banks do not? We hypothesize that this pattern can be rationalized in a simple way if U.S. monetary shocks are inflationary in EMEs but not in AEs. Under this hypothesis, EME central banks have an incentive to move in the same direction as the Fed, since U.S. monetary shocks directly affect the variable they seek to stabilize —namely, inflation. In contrast, if U.S. monetary shocks do not generate inflationary pressures in AEs, then AE central banks have little reason to adjust policy rates in response to such shocks. In this section, we provide evidence in support of this hypothesis.

[13] We focus on the ECB because updated high-frequency monetary shock series are available, which allow us to adapt the methodology of section 3.2 to the ECB as the reference bank.

[14] In these regressions, we instrument the policy rate change with monetary policy shocks constructed from a country-specific Taylor rule purged of global variation. For details, see Appendix A.2.

Section 4.1 shows that a U.S. tightening raises expected inflation and exchange rate depreciation, and lowers expected output growth in EMEs. In contrast, in AEs a U.S. tightening has no systematic effect on inflation or output. Section 4.2 shows that these patterns extend to realized outcomes, with U.S. monetary shocks affecting inflation and the nominal exchange rate in EMEs but not in AEs. Finally, Section 4.3 shows that central banks in economies where inflation is more exposed to U.S. monetary shocks, through the exchange rate, respond more strongly to those shocks, consistent with the mechanism outlined above.

4.1 A Fed Tightening Is Perceived As Inflationary in EMEs

Monetary policymaking goes beyond what reduced-form Taylor rules capture. In practice, expectations are a central ingredient in decisions taken at monetary meetings.[15] Therefore, we first examine how U.S. monetary policy shocks change expectations about EMEs and AEs.

To estimate the immediate perceived impact of U.S. monetary policy shocks on the rest of the world, we focus on expectations of professional forecasters. Specifically, we combine the MPM-dataset with the monthly survey forecasts of Consensus Economics to study how macroeconomics expectations of professional forecasters respond to a surprise U.S. monetary policy decision.[16] Figure 4 illustrates the setup.

We quantify the response of the one-month revision in the consensus forecast of a given variable around an FOMC announcement using a similar setup to that of Nakamura and Steinsson (2018) and Bauer and Swanson (2023a), who study forecast revisions to monetary shocks in the U.S. context. Concretely, we estimate regressions of the form:

$$\Delta \mathbb{E}_\tau \left[x_{c,j} \right] = \alpha_c + \beta \cdot \Delta i_\tau^{US} + \gamma' \cdot \boldsymbol{X}_{c,\tau} + e_{c,\tau}, \tag{3}$$

where the outcome variable, $\Delta \mathbb{E}_\tau [x_{c,j}] \equiv \mathbb{E}_{\tau+\epsilon}[x_{c,j}] - \mathbb{E}_{\tau-\epsilon}[x_{c,j}]$, is the revision in the j-month-ahead forecast of variable x for country c (excluding the U.S.) around FOMC meeting τ;[17] the treatment variable Δi_τ^{US} is the change in the Fed funds rate at meeting τ; and $\boldsymbol{X}_{c,\tau}$ is a

[15] Nakamura et al. (2025) show that the Taylor rule is a poor descriptor of how the Fed operates after the early 2000s; the Fed operates *beyond the Taylor rule*. They emphasize that central banks act to keep inflation expectations anchored. Coibion and Gorodnichenko (2011) and Kakhbod et al. (2025) provide complementary evidence that an expectations-augmented Taylor rule provides a better description of the Fed's behavior.

[16] The Consensus Economics dataset is widely used by researchers, and existing evidence suggests that it provides an accurate proxy for the expectations of market participants. For example, De Marco, Macchiavelli, and Valchev (2022) show that it closely tracks expectations in sovereign bond markets.

[17] Consensus Economics regional surveys report forecasts for the current and next years. We weight these forecasts based on the month of the survey to impute the one-year-ahead forecast, allowing accurate comparison of forecast revisions across survey months. For example, for February 2018 the one-year-ahead forecast is computed as $(10/12) \cdot F(\textit{current year}) + (2/12) \cdot F(\textit{next year})$, where $F(\cdot)$ denotes the forecast.

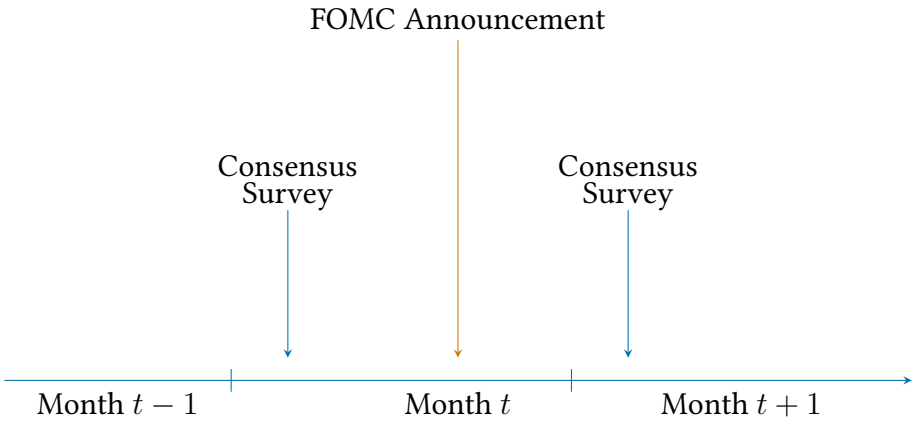


FIGURE 4. ILLUSTRATION: FORECAST REVISIONS AROUND FOMC ANNOUNCEMENT

Notes: The Consensus Economics survey of forecasters is conducted around the middle of each month, while FOMC announcements can occur at any point within the month.

vector of controls including the last twelve revisions of x at horizon j.[18] Finally, we instrument the change in the Fed funds rate using the shocks constructed by Bauer and Swanson (2023b) to ensure causal identification.

Our baseline results are based on an unbalanced panel covering the period from July 1995 to December 2024. To reduce the influence of outliers, we remove all observations of each dependent variable that fall below the 1st or above the 99th percentile of its distribution. This approach preserves the majority of the data while systematically removing extreme observations that could disproportionately affect the regression coefficients.

Table 4 reports the IV estimates of β for forecast revisions of inflation, exchange rate depreciation, and output growth. The results indicate that U.S. monetary shocks are perceived as cost-push shocks in survey expectations: in EMEs, expectations of inflation and exchange rate depreciation are revised upwards, and output growth is revised downwards following a U.S. monetary shock, whereas in AEs, inflation and output expectations remain unchanged despite a tendency for exchange rate depreciation expectations to increase. More precisely, for EMEs, columns (1) through (3) show that a surprise 100 basis points U.S. tightening induces an upward revision of 36, 250, and 190 basis points in expected inflation 12 months ahead and expected depreciation 3 and 12 months ahead, respectively. Column (4) shows that expected output growth 12 months ahead is revised downward by 22 basis points.

[18] Coibion and Gorodnichenko (2012) find a high degree of rigidity in forecast revisions for U.S. inflation. Farmer, Nakamura, and Steinsson (2024) show that forecast revisions for U.S. Treasury bill rates display a non-trivial degree of autocorrelation. Based on these findings, we include lagged revisions in our specification.

Table 4. Forecast Revisions around FOMC Announcements

	Forecast revisions for:			
	12-month inflation $\mathbb{E}[\pi_{c,12}]$	3-month depreciation $\mathbb{E}[\Delta e_{c,3}]$	12-month depreciation $\mathbb{E}[\Delta e_{c,12}]$	12-month growth $\mathbb{E}[\Delta y_{c,12}]$
	(1)	(2)	(3)	(4)
Panel A. Emerging Economies				
Δi_τ^{US}	0.363**	2.544**	1.911**	-0.224**
	(0.143)	(1.015)	(0.951)	(0.089)
Observations	5,819	3,296	3,244	6,067
F-stat (1st stage)	193.7	201.4	199.9	197.3
Panel B. Advanced Economies				
Δi_τ^{US}	0.047	4.146***	1.404	-0.078
	(0.099)	(1.367)	(0.896)	(0.079)
Observations	1,593	1,326	1,285	1,691
F-stat (1st stage)	76.5	75.2	80.5	115.2

Notes: The table reports the coefficient β of equation (3) estimated using U.S. monetary policy shocks à la Bauer and Swanson (2023b) as the instrument for the Fed funds rate. Lag-augmented heteroskedasticity-corrected standard errors are in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Figure 5 complements these findings by presenting estimates of β for revisions in $j = 1, \ldots, 12$ month-ahead inflation forecasts. We show that, for EMEs, inflation following a surprise Fed tightening is expected to rise over time.[19] That is, forecasters do not expect an immediate jump in inflation following the Fed's tightening, but instead expect inflation to keep rising for at least up to 1 year after the FOMC announcement. This is a salient observation in particular because it contradicts sharply the implications of canonical international macro models, where in response to a foreign unexpected monetary tightening, CPI inflation in the home economy overshoots through the exchange rate overshooting.

Together, these estimates indicate that U.S. monetary shocks are perceived as inflationary in emerging market economies (EMEs), but not in advanced economies (AEs). We interpret these findings as suggesting that, in EMEs, U.S. monetary shocks are expected to generate

[19]To construct horizon-specific inflation expectations from monthly survey data, we combine current- and next-year CPI forecasts.

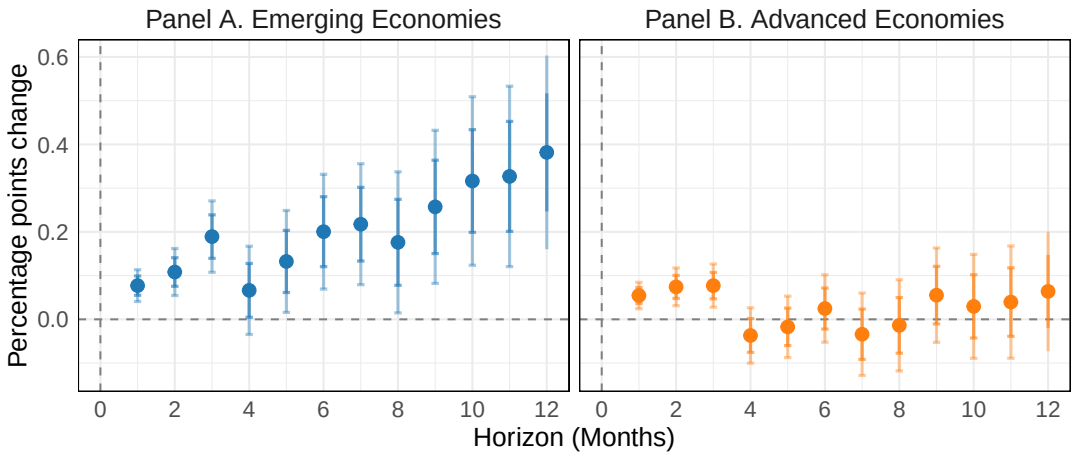


FIGURE 5. PASS-THROUGH TO INFLATION EXPECTATIONS ACROSS HORIZONS

Notes: The pass-through is defined as the estimated coefficient β of equation (3). The light (dark) shaded areas represent the 90% (68%) symmetric confidence intervals, constructed using lag-augmented heteroskedasticity-corrected standard errors.

persistent currency depreciation, which in turn feeds into higher inflation, possibly through channels such as imported goods prices and fluctuations in the domestic-currency value of dollar-denominated private debt.[20] In contrast, in AEs, the expected depreciation is short-lived, which limits the pressure for price adjustments.

Under this interpretation, an EME central bank operating under an inflation-targeting regime, whose expectations can be approximated by those of professional forecasters, has an incentive to move in the same direction as the Fed. By responding to a Fed tightening with an increase in its own policy rate, the central bank mitigates the inflationary pressures induced by the U.S. monetary shock through the exchange rate channel.

An alternative interpretation of our results is that shifts in U.S. monetary policy could affect expectations about foreign economies indirectly, through their effects on the U.S. economy and demand. Appendix Table C1 shows our estimates of forecast revisions for U.S. variables around FOMC announcements. In response to a surprise Fed tightening, forecasters do not significantly revise their expectations for the U.S. economy. This contrast is striking, as forecasters adjust their expectations substantially for foreign economies in the same setting. Therefore, we rule out the possibility that the results above are driven by expectations about U.S. activity or demand, and instead conclude that they operate through exchange rate and

[20] In economies where borrowing in dollars is allowed, such loans are typically issued at long maturities with monthly payments in dollars. These features create an environment in which firms may face currency mismatches, earning revenues in local currency while servicing debt in foreign currency. As a result, exchange rate movements can have direct effects on their production costs and thereby on domestic inflation.

financial channels.

In Appendix C.1 we assess the robustness of the results in this section. We show that our baseline results are robust to the inclusion of outliers (see Table C3), and to instrumenting the change in the Fed funds rate with alternative monetary shock series (see Table C2 and Figure C20).

4.2 A Fed Tightening Is Inflationary in EMEs

We next examine the response of the realized nominal exchange rate, consumer price index (CPI), and unemployment of non-U.S. economies to a U.S. monetary shock. For this purpose, we estimate local projection regressions of the form:

$$y_{c,\tau+h} = \alpha_c^h + \beta^h \cdot \text{MPS}_\tau^{US} + \gamma^{h\prime} \cdot \boldsymbol{X}_{c,\tau} + e_{c,\tau+h}, \quad h = 0, ..., 12, \tag{4}$$

where $y_{c,\tau+h}$ is one of the three outcome variables measured h months after FOMC announcement τ; MPS_τ^{US} is the U.S. monetary policy surprise; $\boldsymbol{X}_{c,\tau}$ is a vector of controls including twelve lags of the outcome variable and the six most recent FOMC decisions on the Fed funds rate; and α_c^h and $e_{c,\tau+h}$ denote country fixed effects and error terms, respectively.

For the baseline estimation of equation (4), we follow the same approach used in Section 3.3. Specifically, we exclude observations affected by the GFC and the COVID-19 pandemic, as well as observations for which the outcome variable falls below the 1st or above the 99th percentile. Alternative estimates obtained by including outliers or observations from the GFC and the COVID-19 period do not differ materially from our baseline results.

Figure 6 reports the estimated responses β^h, along with 90% confidence intervals, for the three outcome variables, separately for EMEs and AEs. The estimate of the response functions show that both groups initially experience a currency depreciation following a contractionary U.S. monetary shock, but the depreciation in EMEs becomes more pronounced over time. After twelve months, the exchange rate in AEs begins to revert toward its pre-shock level, whereas in emerging economies it remains persistently depreciated.

Turning to consumer prices, the price level in EMEs rises immediately above its pre-shock trend following a Fed tightening and remains elevated over the subsequent twelve months. In contrast, consumer prices in AEs do not exhibit a statistically significant response. Hence, these results indicate that U.S. monetary shocks are inflationary in EMEs but not in AEs, despite the broadly similar initial exchange rate movements across the two groups.

Additionally, Figure 6 shows that for economic activity, measured by the unemployment rate, EMEs and AEs exhibit little sensitivity. Thus, even if EME central banks may also have explicit mandates concerning output, U.S. monetary shocks do not pose a significant threat to

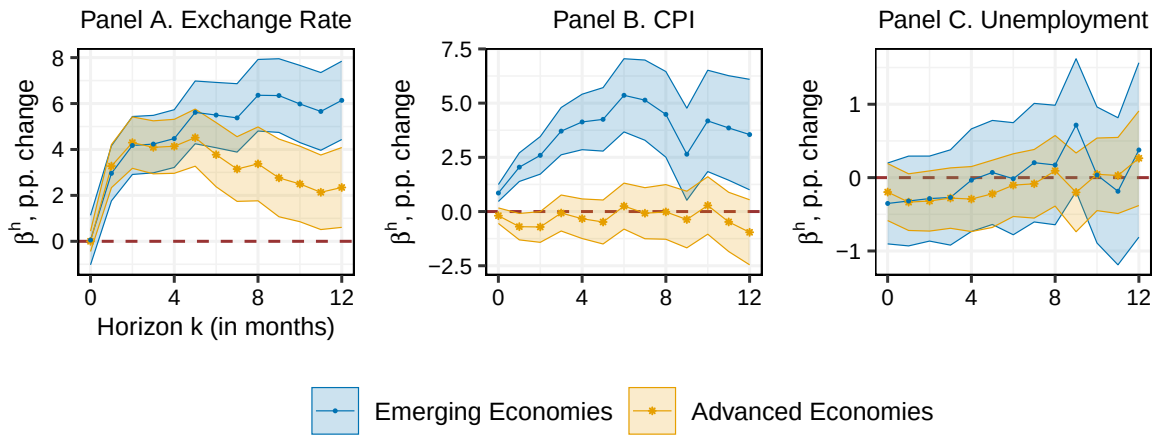


FIGURE 6. RESPONSE OF MACROECONOMIC FUNDAMENTALS TO A U.S. MONETARY SHOCK

Notes: This figure shows the response of the nominal exchange rate (Panel A), consumer price index (Panel B), and unemployment (Panel C) to an exogenous 100 basis points increase in the Fed funds rate. The instrument used to obtain these estimates is the Bauer and Swanson (2023b) monetary shock series. The shaded areas are the 90% confidence intervals, constructed using lag-augmented heteroskedasticity-corrected standard errors.

this objective. From their perspective, raising policy rates to counteract inflationary pressures is a rational policy response.

4.3 Inflation Exposure and Monetary Policy Responses

At this point in the paper, we have established that EME central banks adjust their policy rates in response to U.S. monetary shocks, and that their behavior is consistent with the inflationary effects of a Fed tightening. A natural next question is how the exposure of inflation to U.S. monetary shocks shapes the intensity with which central banks respond to Fed decisions.

To answer this question, we estimate regressions of the following form:

$$\Delta i_{c,\tau+\epsilon} = \alpha_c + \alpha_\tau + \delta \cdot \left(\text{MPS}_\tau^{US} \times \text{Exposure}_{c,\tau}\right) + \gamma' \boldsymbol{X}_{c,\tau+\epsilon} + e_{c,\tau+\epsilon}, \tag{5}$$

where α_c and α_τ denote country and FOMC announcement fixed effects; MPSτ^{US} is the U.S. monetary policy shock at FOMC announcement τ, measured using the shocks from Bauer and Swanson (2023b); Exposure$_{c,\tau}$ captures a country's inflation exposure to U.S. monetary shocks through the exchange rate channel; $\boldsymbol{X}_{c,\tau+\epsilon}$ is a vector of controls including the exposure measure, the most recent change in inflation, and GDP growth; and $e_{c,\tau+\epsilon}$ is the error term.

The coefficient δ captures how the response of a central bank's policy rate to U.S. monetary shocks varies with the economy's inflationary exposure to such shocks. A positive value of δ implies that central banks in more exposed economies adjust their policy rates more

strongly in response to a given U.S. monetary shock. Quantitatively, a one-unit increase in the exposure measure is associated with an additional δ percentage point increase in the policy rate following a 100 basis points Fed tightening.

We require measures of countries' inflation exposure to U.S. monetary shocks. We consider three alternative proxies. First, we use a measure of financial dollarization, defined as the share of deposits denominated in U.S. dollars in the economy (compiled as part of the MPM-dataset). This measure captures the extent to which firms' borrowing costs are exposed to exchange rate fluctuations induced by U.S. monetary shocks (see Christiano, Dalgic, & Nurbekyan, 2021, for a detailed discussion on why this is a good proxy).[21] Second, we use the ratio of imports to GDP (from United Nations, 2025), which reflects that economies that are more dependent on imports may face greater inflationary pressures through exchange rate pass-through. Third, we use the ratio of imports invoiced in U.S. dollars (from Gopinath et al., 2020) to GDP, which accounts for the fact that exchange rate pass-through depends not only on the volume of imports, but also on the share of imports invoiced in U.S. dollars. Together, these three measures capture the idea that U.S. monetary shocks affect countries' price level through exchange rate fluctuations.

Table 5 reports estimates of δ using ordinary least squares for each exposure measure separately (columns 1 through 3), as well as jointly (column 4). In all specifications, the coefficients indicate that central banks in countries with greater inflation exposure adjust their policy rates more strongly in response to a Fed tightening. Quantitatively, a 10% increase in financial dollarization, the imports-to-GDP ratio, and the share of imports invoiced in U.S. dollars is associated with an additional increase in policy rates of 12, 15, and 19 basis points, respectively, following a 100 basis points Fed tightening. Moreover, the estimated relationships are statistically significant at conventional levels across specifications, and remain robust in the joint estimation.

The exposure measures we use are not exogenous, and therefore the estimates of δ should not be interpreted as causal. These measures are likely correlated with other structural characteristics of the economy; for example, in our sample, the correlation between the ratios of imports and exports to GDP is 70%. Nevertheless, the fact that we obtain a positive estimate of δ across three distinct measures of inflation exposure —capturing channels operating through both the financial and real sectors— provides consistent evidence in support of our mechanism. Hence, a natural interpretation is that central banks in economies that are more

[21] We use the share of deposits denominated in U.S. dollars, rather than loans, because most central banks in our sample report deposits disaggregated by currency, while loans are typically reported under a single category (namely, foreign currency). As a result, publicly available data on loan dollarization are prone to substantial measurement error.

Table 5. Marginal-Cost Exposure to U.S. Monetary Shocks Regressions

	Policy rate decision $\Delta i_{c,\tau+\epsilon}$			
MPS$^{US}\times$	(1)	(2)	(3)	(4)
Share of deposits in USD (%)	0.012*			0.021**
	(0.007)			(0.010)
Imports to GDP (%)		0.015***		0.016**
		(0.004)		(0.007)
Share of imports invoiced in USD (%)			0.019**	0.022**
			(0.008)	(0.010)
Observations	2,780	5,732	4,006	1,806
R^2	0.364	0.262	0.269	0.397

Notes: The table reports the ordinary least squares estimates of coefficient δ from equation (5), using the monetary shocks from Bauer and Swanson (2023b). The vector of controls in specifications (1)–(4) includes the most recent change in annual CPI inflation, output growth, and the measure (s) of exposure under consideration. Driscoll and Kraay (1998) standard errors are reported in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

exposed to the inflationary effects of U.S. monetary shocks, operating through the exchange rate, respond more strongly by adjusting their policy rates in the same direction as the Fed.

Additional evidence from studies using micro-data. Evidence from microeconomic studies suggests that inflation in EMEs is, at least in part, driven by U.S. monetary policy. For instance, Cabezas et al. (2024) use Chilean firm-level data to show that a 1% exchange rate depreciation —which can be induced by a U.S. monetary shock— raises the prices of goods used by firms by about 0.5% on impact. Likewise, Morais, Peydró, Roldán-Peña, and Ruiz-Ortega (2019) find that a one standard deviation increase in the Fed funds rate reduces the volume of U.S. loans to Mexico by 6%. This contraction likely raises equilibrium lending rates and, in turn, the marginal costs faced by firms in EMEs, ultimately putting upward pressure on domestic inflation.

5 A Model of U.S. Monetary Policy Spillovers

In this section, we introduce a three-country model with price rigidities, in which EME imports are denominated in foreign currency and a share of firms' loans is denominated in U.S. dollars. The model highlights that, in general equilibrium, a Fed tightening appreciates the

dollar, increasing the cost of imports and firms' debt service in EMEs, and thereby generating inflationary pressures. In contrast, AEs are less exposed in the model because they are able to borrow and export in their local currency. As a result, a Fed tightening induces the EME central bank to move in the same direction as the Fed, whereas the AE central bank does not need to adjust its policy rate to the same extent.

Section 5.1 presents the details of the model. Section 5.2 describes our calibration strategy. Finally, Section 5.3 discusses the effects of a Fed tightening through the lens of the model and relates them to the empirical results documented in Sections 3 and 4.

5.1 Model

We study a three-bloc New Keynesian model with the United States, an advanced economy (AE), and an emerging market economy (EME), building on the framework of Akinci et al. (2024). For convenience, we refer to the United States as the Home economy (H), the EME as Foreign (F or *), and the AE as the Rest of the World (W or **).

5.1.1 Households

Households in each bloc maximize expected lifetime utility given by:

$$\mathbb{E}_t \sum_{j=0}^{\infty} \beta^j \left(\frac{\sigma}{\sigma - 1} C_{t+j}^{\frac{\sigma-1}{\sigma}} - \frac{\chi_0}{1+\chi} L_{t+j}^{1+\chi} \right), \tag{6}$$

where C_t denotes aggregate consumption of the final good; L_t denotes total hours worked; the parameter β is the subjective discount factor; σ is the elasticity of intertemporal substitution; and χ is the inverse of the Frisch elasticity of labor supply.

Consumption is structured in two layers. At the top level, the final consumption bundle C_t is a CES aggregate of tradeable and non-tradeable goods, denoted by C_t^T and C_t^N. At the second level, the tradeable component C_t^T is itself a CES composite of tradeable goods produced in the three blocs:

$$C_t^T = \left\{ \left(\nu_H^i\right)^{\frac{1}{\rho_c}} C_{Ht}^{\frac{\rho_c-1}{\rho_c}} + \left(\nu_F^i\right)^{\frac{1}{\rho_c}} C_{Ft}^{\frac{\rho_c-1}{\rho_c}} + \left(\nu_W^i\right)^{\frac{1}{\rho_c}} C_{Wt}^{\frac{\rho_c-1}{\rho_c}} \right\}^{\frac{\rho_c}{\rho_c-1}}, \tag{7}$$

where ν_j^i denotes the weight assigned by country i to imports from country j in its tradeable consumption bundle, and ρ_c governs the elasticity of substitution across tradeable goods.

The U.S. household faces the following period budget constraint:

$$P_{C,t} C_t + \frac{B_t}{R_t^n} \leq B_{t-1} + W_t L_t + \sum_{i=T,N} \left(R_{K,t}^i \overline{K}^i + \Gamma_t^i \right) \tag{8}$$

where B_t denotes nominal domestic bond holdings; R_t^n is the gross nominal interest rate; $W_t L_t$ represents labor income (with wages assumed to be equal across the tradable and non-tradable sectors); and households receive income from capital and firm ownership in both sectors $i \in T, N$, where $\overline{K}^i$ denotes the fixed capital endowment earning return $R_{K,t}^i$, and Γ_t^i denotes firms' profits.

We assume that EMEs have access to one-period, risk-free international bonds denominated in U.S. dollars. The representative EM household is subject to the following budget constraint:

$$P_{C,t}^* C_t^* + \frac{B_{Ht}^*}{S_{Ft}^H R_t^n \phi_{Ht}^*} \leq \frac{B_{Ht-1}^*}{S_{Ft}^H} + W_t^* L_t^* + \sum_{i=T,N} \left(R_{K,t}^{i\,*} \overline{K}^{*i} + \Gamma_t^{*i} \right) \tag{9}$$

where B_{Ht} denotes a non-state-contingent bond denominated in U.S. dollars that pays the interest rate R_t^n. EME households face a small premium on their holdings of these bonds, captured by ϕ_{Ht}.[22] In the case of AEs, households face analogous period budget constraints from which one can derive an international risk-sharing condition that mirrors the one faced by households in EMEs.

5.1.2 Production

In each bloc, final good $i = \{T, N\}$ is produced according the Cobb-Douglas production function:

$$Y_t^i = \overline{K}^{i\alpha} (Z_t^i L_t^i)^{\gamma_p} \left(M_t^i \right)^{1-\alpha-\gamma_p} \tag{10}$$

where L_t^i denotes labor employed in sector i, satisfying $\sum_i L_t^i = L_t$, and M_t^i denotes the bundle of intermediate inputs used in that sector. The parameters α and γ_p govern the contributions of capital and labor to production. We model M_t as a CES aggregate of the domestic non-tradable intermediate good and tradable intermediate goods sourced from each bloc.

In this open economy, intermediate inputs are assumed to have the same price as final goods. Thus, in the case of the United States, the tradable final good Y^T and the domestic tradable input M_H^i share the price P_H. Likewise, the non-tradable final good Y^N and the corresponding intermediate input M_N^i are both priced at P_N. This assumption applies to prices in the three blocs of our model's the global economy.

[22] $\phi_{H,t} = \exp\left(-\nu_H \frac{B_{Ht}}{S_{Ft}^H P_{C,t} Y_t^*}\right)$ is the adjustment cost associated with U.S. bond holdings for EME households, which ensures stationarity of the model. See Schmitt-Grohe and Uribe (2003) for details.

5.1.3 Working capital constraint in EMEs

For sector $i \in \{T, N\}$, final-good producers in EMEs face non-capital payments given by

$$\Upsilon_t^i = P_{M,t}^i M_t^i + W_t L_t^i, \tag{11}$$

which represent the total payments for intermediate goods and labor. We assume that production is completed at the end of period t, implying that firms must borrow in advance to finance these factor payments.

Specifically, we model the effective cost of intermediate goods and labor faced by firms in sector i of the EM as

$$\Upsilon_t^i \cdot \left\{ 1 + \underbrace{\theta \left[(1-\kappa^*) R_t^{n*} + \kappa^* R_{\text{risk},t}^n - 1 \right]}_{\text{Effective borrowing rate}} \right\}, \tag{12}$$

where θ captures the mark-up over the net policy rates, κ^* denotes the share of borrowing denominated in U.S. dollars, R_t^{n*} is the EM gross interest rate, and $R_{\text{risk},t}^n$ represents the U.S. interest rate R_t^n adjusted for a time-varying risk premium, proxied by the nominal exchange rate depreciation against the dollar, given by

$$R_{risk,t}^n = R_t^n + \left(\frac{S_{H,t}^F}{S_{H,t-1}^F} - 1 \right), \tag{13}$$

with S_H^F the nominal exchange rate of EMs (units of EM currency per U.S. dollar).

Assumption (12) captures two key empirical features emphasized in the literature. First, as discussed since Ize and Levy Yeyati (2003) and more recently by Gutierrez, Ivashina, and Salomao (2023), firms in EMEs borrow in both local and U.S. dollar currency. In our framework, this feature is represented by the parameter κ^*.[23] Second, borrowers face lending rates that include a mark-up over policy rates.[24] We capture this empirical regularity through the parameter θ.

Assumption (13) embeds the idea that borrowing costs vary with exchange rate movements. Bruno and Shin (2015) document that a stronger U.S. dollar is associated with tighter

[23] We treat the share of U.S. dollar borrowing as exogenous. This assumption can be micro-founded, as in Dalgic (2024), where entrepreneurs optimally choose the currency composition of their liabilities. Here, we follow the approach in Camara et al. (2024) and Fukui, Nakamura, and Steinsson (2025).

[24] For instance, Barquero-Romero, Loaiza-Marín, and Mendoza-Fernández (2021) use micro-level data to show that market rates in Costa Rica on both local and U.S. dollar loans are, on average, about three times higher than the corresponding policy rates.

financing conditions for firms. Moreover, Akinci and Queralto (2023) provide a microfoundation for this mechanism.[25] In our framework, we capture this feature by allowing dollar-denominated borrowing costs to rise when the dollar appreciates.

The final good-producing firm in sector $i \in \{T, N\}$, maximizes profits:

$$\Gamma_t^{i*} = P_t^* Y_t^{i*} - R_{K,t}^{i*} \bar{K}^{i*} - \left\{1 + \theta \left[(1-\kappa^*) R_t^{n*} + \kappa^* R_{\text{risk},t}^n - 1\right]\right\} \Upsilon_t^i \tag{14}$$

where P_t^* is the production price associated with Y_t^{i*}, and the last term represents the gross interest payment to working capital. Appendix D presents the derivations relating to the firm's profit maximization and the exposure of the firm's marginal cost to changes in the exchange rate.

5.1.4 Price Setting

For each sector $i \in \{T, N\}$, there is a continuum of retail firms $\ell \in [0, 1]$ that produces sector i's good, which is aggregated into final output Y_t^i as the CES composite

$$Y_t^i = \left(\int_0^1 Y_{\ell,t}^{i\ \frac{1}{1+\theta_p}}\, d\ell \right)^{1+\theta_p}, \tag{15}$$

where $Y_{\ell,t}^i$ is sector i output by retailer ℓ and θ_p denotes the elasticity of substitution among retailers' output. Due to the CES properties, the price level of domestic sector i final output is $P_t^i = \left(\int_0^1 P_{\ell,t}^{i\ -\frac{1}{\theta_p}} dl \right)^{-\theta_p}$ where $P_{\ell,t}^i$ is the price set by Home retailer ℓ. Cost minimization by users of sector i final output results in the demand function for firm ℓ in sector i's output:

$$Y_{\ell,t}^i = \left(\frac{P_{\ell,t}^i}{P_t^i} \right)^{-\frac{1+\theta_p}{\theta_p}} Y_t^i. \tag{16}$$

We assume that domestic retail producer ℓ in sector i produces according to the Cobb-Douglas production function

$$Y_{lt}^i = \overline{K}_{lt}^{i\ \alpha} (Z_{lt}^i L_{lt}^i)^{\gamma_p} \left(M_{lt}^i\right)^{1-\alpha-\gamma_p} \tag{17}$$

and that it resets its price with probability $1-\xi_p$. Otherwise, they keep their prices unchanged.

Our model incorporates the notion of multiple currency pricing (MCP) introduced by Gopinath et al. (2020). Specifically, we assume that EMEs' exports to the United States are invoiced in U.S. dollars, the dominant currency in international trade. Under the MCP assumption, EMEs' firms price their exports to the U.S. in dollars, while those directed to AEs

[25] Akinci and Queralto (2023) develop a general equilibrium model linking the risk premium faced by EMEs to the depreciation of their exchange rates following unexpected U.S. monetary tightening.

are priced in their own currency. In contrast, exporters in the U.S. and other AEs set export prices in their respective domestic currencies, consistent with the standard producer-currency-pricing framework.[26]

5.1.5 Central Banks

In each bloc, central banks target the output gap and CPI inflation. In particular, central banks follow a Taylor rule of the form:

$$R_t^n = \left(R_{t-1}^n\right)^{\gamma_r} \left(\beta^{-1} \left(\Pi_t\right)^{\gamma_\pi} \left(\frac{Y_t}{Y_{ss}}\right)^{\gamma_y}\right)^{1-\gamma_r} \exp\left(\epsilon_t^{mp}\right) \tag{18}$$

where $\Pi_t \equiv P_{C,t}/P_{C,t-1}$ is gross consumption based price inflation; Y_{ss} is the potential level of output; the parameters γ_π and γ_y govern the weight placed on inflation and output stabilization, while γ_r captures the degree of interest rate smoothing; and ϵ_t^{mp} is a monetary policy shock, modeled as a first-order autoregressive process with persistance ρ_m:

$$\epsilon_t^{mp} = \rho_{mp} \cdot \epsilon_{t-1}^{mp} + \eta_t^{mp}, \tag{19}$$

where the η_t^m are mean-zero independent and identically distributed shocks.

5.2 Calibration

We calibrate the model closely to Akinci et al. (2024), and the references therein.[27] We calibrate the Home bloc to the U.S. and take the Foreign and Rest of the World to represent the EMEs and AEs, respectively, with strong trade and financial linkages to the U.S. Our calibration is asymmetric with the U.S. being relatively closed to trade.

We depart from Akinci et al. (2024) in three ways. First, we allow central banks to stabilize the output gap along with CPI inflation. To calibrate the Taylor rule in each bloc, we draw on the estimates in De Leo et al. (2024). These estimates imply that EMEs exhibit less interest rate inertia and place greater effective weight on inflation stabilization relative to AEs due to higher interest rate smoothing (see Table 6 below).[28] Second, we allow the share of working capital that is dollarized in EMEs, denoted by κ^*, to be less than one. In the baseline, we set $\kappa^* = 0.30$, which is close to the average level of dollarization observed in our sample. Third, we calibrate the model at a monthly, rather than quarterly, frequency to study how non-U.S. central banks respond to U.S. monetary shocks in a higher-frequency setting.

[26] This pricing assumption is motivated by empirical evidence. Several studies, including Goldberg and Tille (2008), Gopinath et al. (2020), and Boz et al. (2025, 2022), document that international trade is largely conducted in a few currencies, with the U.S. dollar playing the dominant role.

[27] We describe the full list of the model calibration along with the equilibrium conditions in Appendix D.

[28] In appendix D.3, we rationalize why EME central bankers are more hawkish than AE central bankers.

Table 6. Taylor Rule Parameter Descriptions and Estimates

Parameter	Description	Calibration
United States and Advanced Economy		
γ_r	Inertia in Taylor rule	0.92
γ_π	Taylor rule response to inflation	3.40
γ_y	Taylor rule response to output gap	1.20
Emerging Economy		
γ_r^*	Inertia in Taylor rule	0.83
γ_π^*	Taylor rule response to inflation	2.41
γ_y^*	Taylor rule response to output gap	0.34

Notes: Value of parameters based on the estimated of De Leo et al. (2024).

5.3 Results

We use the model to study the general equilibrium effects of a U.S. monetary tightening. Specifically, we consider a U.S. monetary policy shock analogous to that in our empirical analysis and examine its macroeconomic effects across countries, as well as its implications for the policy decisions of EME and AE central banks.

5.3.1 Global effects of a U.S. monetary tightening

Figure 7 presents the deviations from steady state of selected variables in the EME and AE blocs in response to a 100 basis points increase in the Fed funds rate. The impulse responses for the EME and AE blocs are shown in blue and yellow, respectively. Interest rate changes are reported in basis points to facilitate comparison with the empirical results.

Overall, the model matches our main empirical observations: *i)* EMEs raise rates by almost half of the Fed tightening; *ii)* EMEs face significant inflation following a Fed tightening and *iii)* inflation and policy rate responses are muted for AEs.

The top row of Figure 7 shows that, following a contractionary Fed shock, the EME central bank increases its policy rate by 47 basis points, whereas the AE central bank raises its rate by 13 basis points. These responses are comparable to the empirical estimates reported in Table 2. The difference in responses is primarily driven by the larger increase in inflation in EMEs, which rises by about 2.5%, roughly two to three times the increase observed in AEs. From an inflation-targeting perspective, this implies that the EME central bank has a stronger

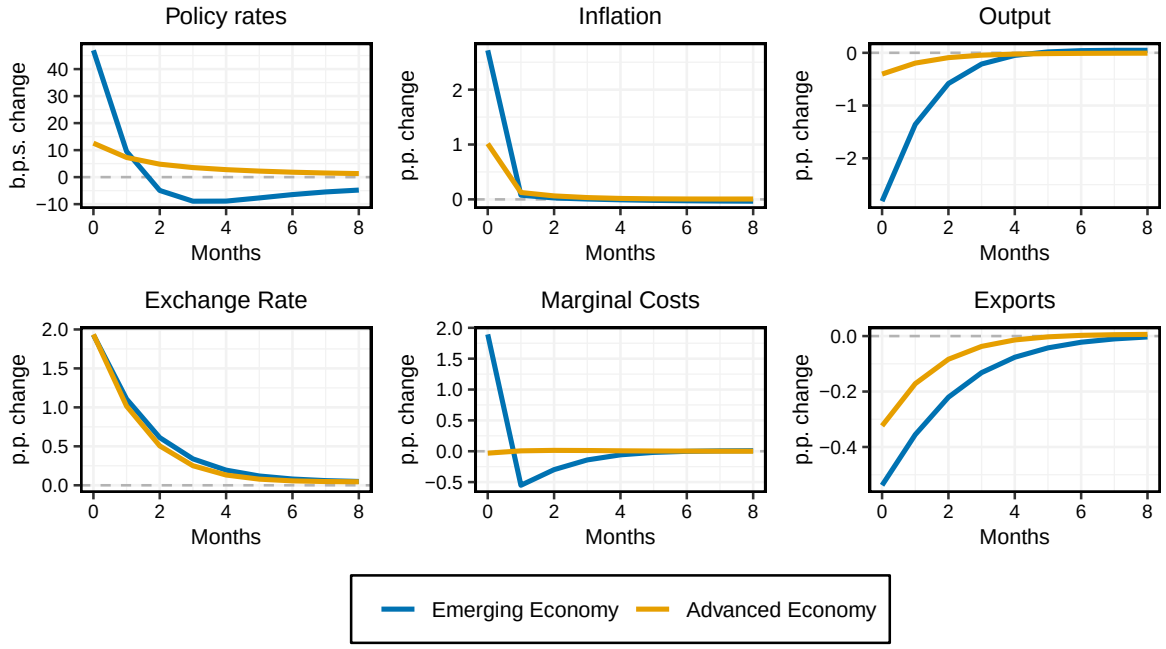


FIGURE 7. MODEL IRFS IN RESPONSE TO U.S. TIGHTENING

Notes: The figure shows the macroeconomic effects of a 100 basis points increase in the Fed funds rate in our model.

incentive to tighten policy in response to the Fed. In addition, the EME central bank places seven times more weight on inflation stabilization than on the output gap, so the contraction in output, approximately 2.8%, is more than offset by the rise in inflation. In appendix D.3, we show through the lens of the model why EME central bankers put more weight on inflation relative to the output gap.

The second row of Figure 7 provides insight into the mechanisms through which the U.S. monetary shock operates. On impact, a U.S. monetary tightening depreciates the currencies of both the EME and AE. In the EME, this depreciation has a significant effect on firms' marginal costs because their loans are partially dollarized and imports must be paid for in foreign currency. Combined with higher import prices, this leads to an increase in CPI inflation. In contrast, firms' marginal costs in the AE are less affected by currency depreciation because their loans are not dollarized, and less exposed to imports relative to EMEs. As a result, the AE faces weaker inflationary pressures than the EME following a Fed tightening.

The last panel shows that exports in both the EME and AE decline following a Fed tightening. Intuitively, a Fed tightening acts as a negative demand shock in the United States, reducing demand for foreign goods. In the case of the EME, exports to the United States are priced in dollars, so the expenditure-switching channel is muted. In contrast, AE exports are

priced in their own currency, so exchange rate movements help mitigate the decline in demand through expenditure switching. As a result, the contractionary effects of a Fed tightening are more pronounced for the EME than for the AE.

5.3.2 Drivers of policy rate adjustments in emerging economies

To shed light on the mechanisms underlying the policy rate response of the EME to a Fed tightening, we vary, one at a time, the parameters κ^* and ν_H^F, which govern the share of dollarized working capital and the weight assigned by EME households to imports from the United States in the tradeable consumption bundle, respectively. For each parameter, we evaluate how the EME policy rate responds to a 100 basis points Fed tightening across a grid of parameter values, keeping all other parameters fixed.

Figure 8 presents the results of this exercise for κ^* and ν_H^F in Panels A and B, respectively. The red lines show the model-implied response of the EME policy rate as each parameter varies. Teal points report country-specific responses based on the MPM-dataset,[29] and dashed lines correspond to linear regressions of these estimates on the variable shown on the horizontal axis. Overall, both the model and the data indicate that the response of EME central banks increases with the degree of debt dollarization and dollar invoicing.

These results highlight that the positive correlation documented in Table 5 between the intensity of central bank responses and the degree of financial dollarization and dollar invoicing is consistent with a general equilibrium framework. In particular, the model-implied response lines closely match the intercepts and slopes of the empirical regression lines, reinforcing the view that the results in Table 5 are not spurious. Indeed, the fact that two completely different approaches, one reduced-form and the other structural, produce such similar patterns provides strong support for the view that debt dollarization and dollar invoicing are central to the observed heterogeneity in how EME central banks respond to U.S. monetary shocks.

6 Conclusion

We show that EME central banks systematically follow the Fed: at their first policy meeting after an exogenous 100 basis points increase in the Fed funds rate, they raise their policy rate by about 50 basis points. In contrast, AE central banks appear largely neutral to Fed shocks. A

[29] We estimate regression ‘a la (1) for each country. The estimated coefficient β is used as a measure of country-specific responses to a 100 basis points Fed tightening. Because these estimates are subject to substantial measurement error due to differences in sample length across countries, we apply an empirical Bayes shrinkage procedure to reduce the influence of such noise.

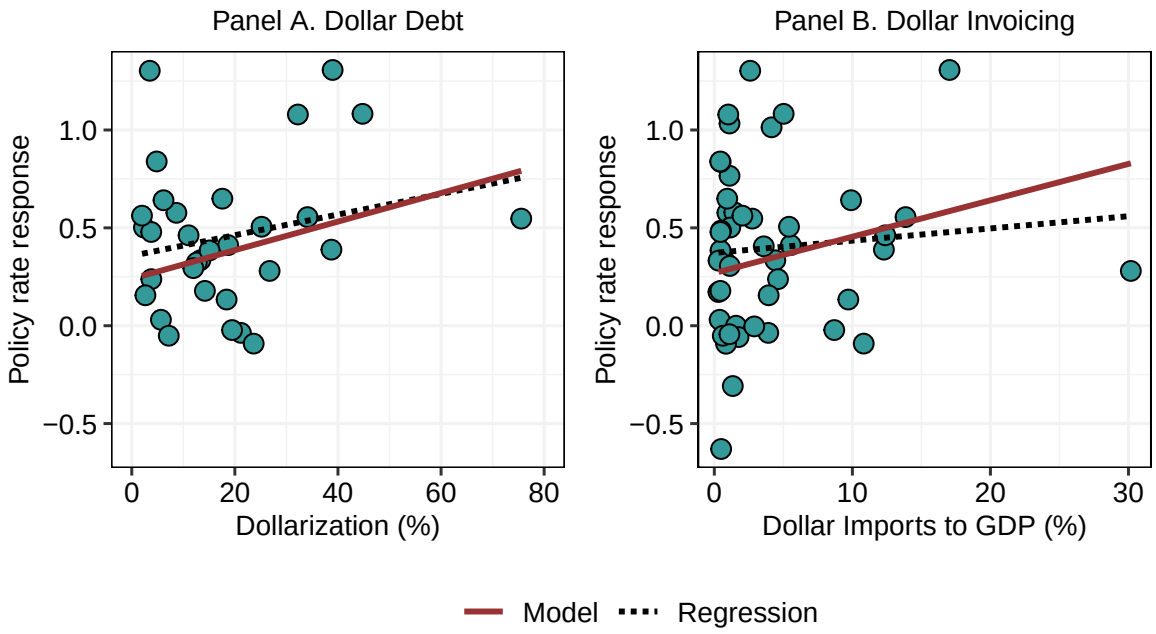


Figure 8. Heterogeneity of EME Responses to a Fed Tightening

Notes: The figure reports the response of the EME policy rate to a 100 basis points Fed tightening as a function of the degree of financial dollarization κ^* (Panel A) and the weight of imports from the United States in the tradeable consumption bundle ν_H^F (Panel B). Red lines correspond to model-implied responses, teal points to country-level estimates based on the MPM-dataset, and dashed lines to linear fits of these estimates.

key element behind identifying this finding is the dataset on monetary policy meetings, which makes it possible to exploit the precise timing of meetings relative to FOMC announcements and thus isolate the causal effect of Fed shocks on central-bank policy decisions.

The causal effects we uncover indicate that, from the perspective of inflation-targeting central banks, a contractionary Fed shock is inflationary. Hence, tightening in response to the Fed contraction constitutes a response consistent with the mandate of price stability. We argue that these empirical patterns align with the structural features of EMEs, where the marginal costs of domestic firms move in tandem with local-currency depreciation —which can themselves be triggered by contractionary Fed shocks. We show that a model with trade and financial linkages in EMEs to the U.S. dollar can replicate the policy responses of inflation-targeting EME central banks following a Fed tightening.

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Appendices

A Appendix to Section 2

In this appendix, we describe the sample coverage of the Monetary Policy Meetings dataset (MPM-dataset) and document the sources of the macroeconomic variables accompanying it. Subsection A.1 describes the range of meetings recorded for each economy in the sample, together with the sources used for inflation and GDP data. Subsection A.2 briefly discusses the series of monetary policy shocks considered in the analyses. Subsection A.3 concludes by describing the assumptions made when merging the policy meetings data with the macroeconomic data and monetary policy shock series.

A.1 Sample Coverage

Policy rates. In Table A1, we present the date of the first and last meeting recorded for each country in our sample (columns *start* and *end*, respectively), along with the total number of meetings for each country (column *meetings*). We also report the income category, classified according to the World Bank's four income groups as of 2023 (second column). The table shows that close to 40% of the meetings recorded are from European countries, while the Americas and Asia and the Pacific each account for approximately 25% of the total. The remaining 10% of meetings come from Semitic and Sub-Saharan countries.

We clarify some important details regarding the meetings included in our sample and the definition of meeting dates. First, we made our best effort to include as many meetings as were available on the public websites of central banks. For countries like Australia, this was possible due to the long historical record posted by their central bank, which allowed us to collect data starting in 1990. Nonetheless, for most countries, data on policy meetings is not publicly available from the moment that central banks began using a policy rate as their instrument. Two clear examples are Japan and Canada, for which policy meeting announcements are available only from 1998 and 2000, respectively, despite it being well known that they had been deciding a policy rate since at least the early 1990s. Second, regarding how we defined the date of a meeting, there were several instances where meetings lasted more than one day (e.g., this was often the case in Iceland). In such cases, we define the meeting date as the last day of the period during which the policy meeting took place.

Table A1. Policy Rates

Country	WB income group	Start	End	Meetings
Americas				2,761
Argentina	Upper middle	2016-08-09	2024-11-01	126
Brazil	Upper middle	1999-03-04	2024-12-11	243
Canada	High	2000-12-05	2024-12-11	197
Chile	High	2001-08-09	2024-12-26	187
Colombia	Upper middle	2007-06-15	2024-12-17	254
Costa Rica	Upper middle	2008-05-28	2024-12-19	382
Dominican Republic	Upper middle	2008-01-31	2024-12-29	197
Guatemala	Upper middle	2005-02-18	2024-11-27	169
Honduras	Lower middle	2015-03-17	2024-12-17	62
Jamaica	Upper middle	2021-03-29	2024-12-19	31
Mexico	Upper middle	2008-02-15	2024-12-19	146
Paraguay	Upper middle	2013-11-13	2024-12-19	138
Peru	Upper middle	2003-11-06	2024-12-12	255
Trinidad and Tobago	High	2011-01-28	2024-12-27	85
United States	High	1995-02-01	2024-12-18	255
Uruguay	High	2020-09-24	2024-12-23	34
Europe				3,634
Albania	Upper middle	2001-08-15	2024-12-18	187
Armenia	Upper middle	2006-05-04	2024-12-10	167
Azerbaijan	Upper middle	2007-04-03	2024-12-18	82
Czech Republic	High	1998-01-22	2024-12-19	262
European Union	High	1999-03-04	2024-12-12	305
Georgia	Upper middle	2008-01-22	2024-12-17	159
Hungary	High	2004-12-20	2024-12-17	239
Iceland	High	2009-03-18	2024-11-19	123
Kazahkstan	Upper middle	2008-03-31	2024-11-29	77
Kyrgyzstan	Lower middle	2014-02-25	2024-10-28	93
Moldova	Upper middle	2010-01-28	2024-12-05	157
North Macedonia	Upper middle	2015-01-13	2024-12-17	107
Norway	High	1999-01-27	2024-12-18	207

Country	WB income group	Start	End	Meetings
Poland	High	2001-01-22	2024-12-04	277
Romania	High	2003-10-06	2024-11-08	62
Russia	High	2013-09-13	2024-12-20	98
Serbia	Upper middle	2006-09-13	2024-12-12	256
Sweden	High	1994-01-22	2024-12-18	185
Tajikistan	Lower middle	2017-01-31	2024-10-25	38
Turkey	Upper middle	2010-05-18	2024-12-26	168
Ukraine	Upper middle	2014-04-14	2024-12-12	88
United Kingdom	High	1997-06-06	2024-12-18	297
Asia and the Pacific				2,356
Australia	High	1990-08-02	2024-12-10	381
China	Upper middle	2019-08-20	2024-12-20	65
India	Upper middle	2008-10-24	2024-12-06	106
Indonesia	Upper middle	2016-04-21	2024-12-18	106
Japan	High	1998-01-16	2024-12-19	355
Malaysia	Upper middle	2004-05-26	2024-11-06	130
Mongolia	Upper middle	2013-01-25	2024-12-13	51
New Zealand	High	1996-12-01	2024-11-27	201
Pakistan	Lower middle	2016-01-30	2024-12-16	64
Philippines	Lower middle	2002-01-17	2024-12-19	213
South Korea	High	1999-12-02	2024-11-28	271
Sri Lanka	Lower middle	2012-05-12	2024-11-26	117
Taiwan	High	2001-06-28	2024-12-19	100
Thailand	Upper middle	2000-05-23	2024-12-18	196
Semitic Countries				682
Egypt	Lower middle	2005-06-02	2024-12-26	165
Israel	High	2002-11-25	2024-11-25	236
Morocco	Lower middle	2006-12-19	2024-12-19	72
Tunisia	Lower middle	2004-01-08	2024-12-28	209
Sub-Saharan Africa				257
Nigeria	Lower middle	2007-02-07	2024-11-26	107
South Africa	Upper middle	2000-11-16	2024-11-21	150

Notes: This table presents the countries in the sample (first column), their income classification according to the

World Bank as of 2023 (second column), the dates of their first and last monetary policy meetings (third and fourth columns), and the total number of meetings recorded (fifth column).

Price indices and GDP growth. We complement the meetings data with information on inflation and GDP growth. Inflation and gross domestic product (GDP) growth are measured using the year-over-year relative changes in the consumer price index (CPI) and GDP, respectively. We obtain the (non-seasonally adjusted) data on prices and GDP mainly from the International Financial Statistics (IFS) database of the International Monetary Fund.[30] When CPI or GDP data for a country were not available in the IFS, we collected the data directly from the websites of central banks or national statistical offices. We summarize the data sources for CPI and GDP in Table A2. The IFS did not include the necessary data for seven out of fifty-eight countries in the case of CPI, and for twenty countries in the case of GDP.

Table A2. Prices and GDP Growth Sources

Country	CPI	GDP
Americas		
Argentina	Statistical Office	IFS
Brazil	IFS	IFS
Canada	IFS	IFS
Chile	IFS	IFS
Colombia	IFS	IFS
Costa Rica	IFS	IFS
Dominican Republic	IFS	Central Bank
Guatemala	IFS	Central Bank
Honduras	IFS	IFS
Jamaica	IFS	Central Bank
Mexico	IFS	IFS
Paraguay	IFS	IFS
Peru	IFS	Central Bank
Trinidad and Tobago	IFS	IFS
United States	IFS	IFS

[30] Using seasonally adjusted data would not materially affect the results, as growth rates are largely insensitive to seasonal adjustments. Typically, the correlation between growth rates computed with seasonally and non-seasonally adjusted data is close to 99%. We decided to use non-seasonally adjusted data because the IFS provides much broader coverage for these variables.

Country	CPI	GDP
Uruguay	IFS	IFS
Europe		
Albania	IFS	Statistical Office
Armenia	IFS	Statistical Office
Azerbaijan	IFS	Statistical Office
Czech Republic	IFS	Statistical Office
European Union	Statistical Office	IFS
Georgia	IFS	Statistical Office
Hungary	IFS	IFS
Iceland	IFS	IFS
Kazahkstan	IFS	IFS
Kyrgyzstan	Statistical Office	CEIC database
Moldova	IFS	Statistical Office
North Macedonia	IFS	IFS
Norway	IFS	IFS
Poland	IFS	IFS
Romania	IFS	IFS
Russia	IFS	IFS
Serbia	IFS	IFS
Sweden	IFS	IFS
Tajikistan	CEIC database	CEIC database
Turkey	IFS	IFS
Ukraine	IFS	IFS
United Kingdom	IFS	IFS
Asia and the Pacific		
Australia	Statistical Office	IFS
China	IFS	IFS
India	IFS	IFS
Indonesia	IFS	IFS
Japan	IFS	IFS
Malaysia	IFS	Ministry of Economy
Mongolia	IFS	Statistical Office
New Zealand	Statistical Office	IFS
Pakistan	IFS	Statistical Office

Country	CPI	GDP
Philippines	IFS	IFS
South Korea	IFS	IFS
Sri Lanka	IFS	IFS
Taiwan	Statistical Office	CEIC database
Thailand	IFS	Statistical Office
Semitic Countries		
Egypt	IFS	Ministry of Planning
Israel	IFS	IFS
Morocco	IFS	Statistical Office
Tunisia	IFS	Statistical Office
Africa		
Nigeria	IFS	IFS
South Africa	IFS	IFS

Notes: This table reports the sources used for consumer price index (CPI) and GDP data for each country. When data were available, we used the International Financial Statistics (IFS) database of the International Monetary Fund. Otherwise, we collected the data directly from the websites of central banks or national statistical offices.

A.2 Monetary Policy Shocks for non-U.S. Economies

Monetary Policy Shocks in the U.S. and European Union. We consider monetary policy shock series for the Fed and for the ECB that are the state-of-the-art series for economic research. For the United States, we use U.S. monetary policy shocks à la Nakamura and Steinsson (2018) and Bauer and Swanson (2023b) (henceforth, NS and BS). Both series are updated by Michael Bauer and Eric Swanson up to December 2023, and are available in Bauer's website. For the European Union, we use the monetary shock series of Jarociński and Karadi (2020), which are updated up to October 2024 in Marek Jarociński's website. The monetary shock series considered are widely used in macroeconomic studies due to their construction from high-frequency data around Federal Open Market Committee (FOMC) announcements, providing a credible measure of monetary surprises.

Monetary Policy Shocks in non-U.S. Economies. Unlike the United States and the European Union, there are no readily available monetary shock series for the central banks of other economies. To overcome this limitation, we use a *"double-filter"* methodology that first filters innovations in policy rates not explained by country-specific inflation and economic growth. In a second step, our methodology takes the innovations obtained in the first step

and filter out common global variation from them. With this procedure, we obtain a measure of monetary movements that is unaccounted by country-specific economic fundamentals and common global factors. Below, we detail the procedure and some characteristics of the resulting shock series.

The first step is in the spirit of Taylor (1993). For each country, except for the United States, we estimate the policy rule

$$i_{c,m} = \alpha + \sum_{i=1}^{I} \phi_i \cdot MPR_{c,m-i} + \beta_\pi \cdot \pi_{c,m-\epsilon} + \beta_{\Delta y} \cdot \Delta y_{c,m-\epsilon} + \beta_{FFR} \cdot i_m^{US} + \sum_{j=0}^{J} \psi_j \cdot u_{c,m-j} \tag{A1}$$

where $i_{c,m-i}$ ($i \geq 0$) denotes the policy rate decided by country c's central bank at meeting $m-i$,$\pi_{c,m-\epsilon}$ is the inflation rate observable by central bank c at meeting m, $\Delta y_{c,m-\epsilon}$ is the GDP annual growth rate observable by central bank c at meeting m, i_m^{US} is the observable Fed funds rate on meeting m, and the $u_{c,m}$'s are the unforecastable deviations from this rule. We estimate the number of lags, I and J, using an automatic ARIMA procedure.

Central banks' meetings happen on different dates and at different frequencies, so we impose a time homogenization assumption. We define a period t as the time window starting at the moment the FOMC made an announcement to the day before the FOMC announcement. Therefore, we redefine the Taylor innovations as the quantity $v_{c,t}$ represented by the sum of the $u_{c,m}$'s such that meeting m in country c happened between FOMC announcement t (inclusive) and FOMC announcement $t+1$ (exclusive). That is, if country c had $M > 1$ meetings between dates t and $t+1$, then all the u's from those meetings are added up to form $v_{c,t}$[31]

Turning to the second step, it might be the case that the v's are spatially correlated. For example, during the Global Financial crisis, most central banks predictably eased their policy rates despite observable inflation and output growth not responding very quickly to the news of a global financial shock. Such central banks likely reacted to the shock in advance because they update their forecast of growth and inflation. In such a case, if we ignore the spatial source of variation, our first step would imply that there was an unusual innovation in the policy rule, and that its magnitude is correlated across countries. Of course, we know that is because of a global shock to which most central banks are reacting to, not because of pure chance. Therefore, the variation in the v's is not fully exogenous variation in the behavior of

[31] An example of such a situation ocurring in our data is Costa Rica. The Fed had two consecutive meetings in February 1, 2017 and March 15, 2017. Between such dates the Central Bank of Costa Rica met four times; hence, we estimate four u's for them and have to add them up to obtain $v_{CR,t}$.

central banks; one needs to address the issue of spatial variation.

To deal with this type of situations we estimate the dynamic factor model

$$\underbrace{\begin{bmatrix} v_{1,t} \\ v_{2,t} \\ \vdots \\ v_{N,t} \end{bmatrix}}_{v_t} = \underbrace{\begin{bmatrix} \lambda_{1,1} & \cdots & \lambda_{1,K} \\ \lambda_{2,1} & \cdots & \lambda_{2,K} \\ \vdots & \vdots & \vdots \\ \lambda_{N,1} & \cdots & \lambda_{N,K} \end{bmatrix}}_{\Lambda} \underbrace{\begin{bmatrix} f_t^1 \\ f_t^2 \\ \vdots \\ f_t^K \end{bmatrix}}_{f_t} + \underbrace{\begin{bmatrix} \xi_{1,t} \\ \xi_{2,t} \\ \vdots \\ \xi_{N,t} \end{bmatrix}}_{\xi_t}, \tag{A2}$$

where f_t is the vector of the K common factors at time t that evolves according to a VAR(p) process, Λ is the matrix of factor loading, and ξ_t is the vector of innovations at time t. By construction, $\xi_{c,t}$ is the component of the surprise in central bank c's Taylor rule that cannot be explained by global sources of variation. Thus, this method identifies common factors across countries, which we can subtract from the v-innovations to obtain a quantity that is closer to what a monetary policy shock should be: a quantity unpredictable by idiosyncratic macroeconomic fundamentals, by its lags, and by common spatial variation. Hence, we interpret ξ_t as the vector of monetary shocks at time t.

To select the number of factors K, we use the information criteria developed by Bai and Ng (2002). They recommend estimating models with $k = 1, \cdots, \bar{K}$ factors to obtain the information criterion statistics $IC(p)$, $p \in \{1, 2, 3\}$, as described in their article, and then choosing the number of factors K corresponding to the model with the lowest $IC(p)$. Panel A of Figure A1 shows that, under all three information criteria of Bai and Ng, it is optimal to select $K = 5$ factors. Panel B complements this by showing the contribution of each factor to the total variance. The first factor alone explains close to 60% of the variation in Taylor innovations, and the first four factors jointly account for 70.5% of the total variation. Interestingly, starting from the fifth factor, the contribution of each factor falls below the average contribution to total variance across all factors (dotted line in Panel B).

Given the number of factors, we next select the number of lags p in the VAR that describes the factor dynamics. Table A3 reports the values of the Bayesian, Akaike, and Hannan–Quinn information criteria for models estimated with $p = 1, \cdots, 10$ lags. Each criterion selects a different number of lags, as the value of p that minimizes each criterion varies. However, all three measures suggest choosing a model with $p \leq 4$ lags. Accordingly, we estimate equation (A2) for $p \in \{1, \cdots, 4\}$ and recover four sets of residuals ξ_t, each representing an alternative monetary shock series. The resulting correlation between any pair of series is at least 95%.

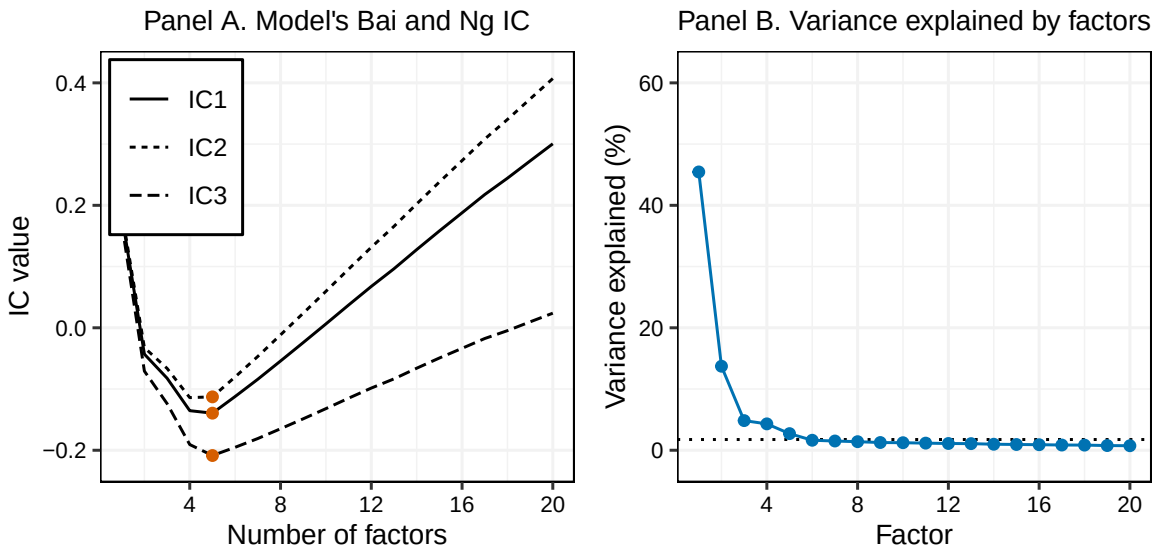


FIGURE A1. SELECTION OF THE NUMBER OF FACTORS

Notes: This figure reports the information criteria values of Bai and Ng (2002) for models with $k = 1, \cdots, 20$ factors (Panel A), and the total contribution to total variance of Taylor innovations v of each factor (Panel B). The orange dots in Panel A are located at the number of factors for which the information criteria values are minimized. The dotted horizontal line in Panel B represents the average contribution to total variance of the Taylor innovations across all factors.

TABLE A3. CRITERIA FOR SELECTING THE NUMBER OF LAGS IN THE FACTOR VAR

Criterium	Lags in factor VAR (p)									
	1	2	3	4	5	6	7	8	9	10
Bayes	2.06	2.09	2.22	2.46	2.83	3.13	3.44	3.82	4.20	4.57
Akaike	1.72	1.50	1.35	1.33	1.44	1.47	1.51	1.63	1.74	1.84
Hannan-Quinn	1.86	1.74	1.70	1.79	2.00	2.14	2.29	2.51	2.74	2.95

Notes: This table reports the information criteria values for models with $p = 1, \cdots, 10$ factors. In the replication code of this article, the Bayes information criteria receives the name of Schwarz information criteria due to how the software is built.

A.3 MPM-Dataset Assembly Process

To assemble the MPM-dataset, we follow a three-step procedure. First, we collect information on monetary policy meetings from the websites of central banks. For each meeting, we record the meeting date, the policy rate in force the day before the meeting, and the policy rate decided at the meeting. This set of variables allows us to track the decisions of monetary

authorities at the time they were made.

The second step is to merge the policy rate data with the information set available in the economy at the time of each meeting. For each local meeting date, we include the most recent figures on the local annual inflation rate and annual GDP growth rate available at the time of the meeting, under the assumption that on month M only inflation and GDP growth information from month $M-1$ and before is available. In addition, for each local meeting date, we incorporate information on the policy rates announced by the Fed in their last three and next three announcements relative to the meeting date. In the final step, we add the monetary policy shocks series that we describe above in section A.2.

A.4 Consensus Economics

We collect the regional forecasts for over 30 countries using the Consensus Economics mean forecasts. Consensus Economics publishes short-term (1 to 2 year ahead forecasts) at a monthly frequency. Coibion and Gorodnichenko (2015) show that professional forecasters' inflation expectations are more aligned with those of central banks compared to households and firms.[32] Moreover, professional forecasters are more accurate and less dispersed than other measures of inflation expectations. Ha, Kose, and Ohnsorge (2019) show this for a large panel of countries. Therefore, using professional forecasters' forecasts enables us to proxy better for the expectations perceived by central bankers when they make policy decisions as well as the future realization of macroeconomic variables.

The list of countries used are presented in Table A4:

Table A4. Policy Rates

Emerging Economies	Advanced Economies
Albania	Australia
Argentina	Canada
Armenia	Switzerland
Azerbaijan	Czech Republic
Brazil	United Kingdom
Chile	Israel
China	Japan
Colombia	South Korea

[32] Coibion and Gorodnichenko (2015) show this for New Zealand and the United States.

Costa Rica	Norway
Dominican Republic	New Zealand
Egypt	Sweden
Georgia	Taiwan
Guatemala	United States
Honduras	
Hungary	
Indonesia	
Kazakhstan	
Sri Lanka	
Moldova	
Mexico	
North Macedonia	
Malaysia	
Nigeria	
Pakistan	
Peru	
Philippines	
Poland	
Paraguay	
Russia	
Serbia	
Thailand	
Turkey	
Ukraine	
Uruguay	
South Africa	

B Appendix to Section 3

B.1 Bunching of Policy Meetings

B.1.1 Prevalence

In the same spirit as Figure 1, Panel A of Figures B2-B6 shows the histograms of the number of days that pass between an FOMC announcement and the next meeting by a central bank, while Panel B of the figures shows the number of days between a local policy meeting and the closest FOMC announcement. Each figure covers the data of one of the five regions considered in this article. Despite some differences in the shape of the histograms, the two facts that we documented in section 3.1 remain:

i) The number of policy decisions made within the first week following an FOMC announcement is notably larger than in each of the subsequent weeks.

ii) Policy meetings concentrate in the week following an FOMC announcement.

Figures B7-B8 and B9 and B10 confirm that the same patterns also hold when dividing economies into high income and middle income, and advanced and emerging economies.

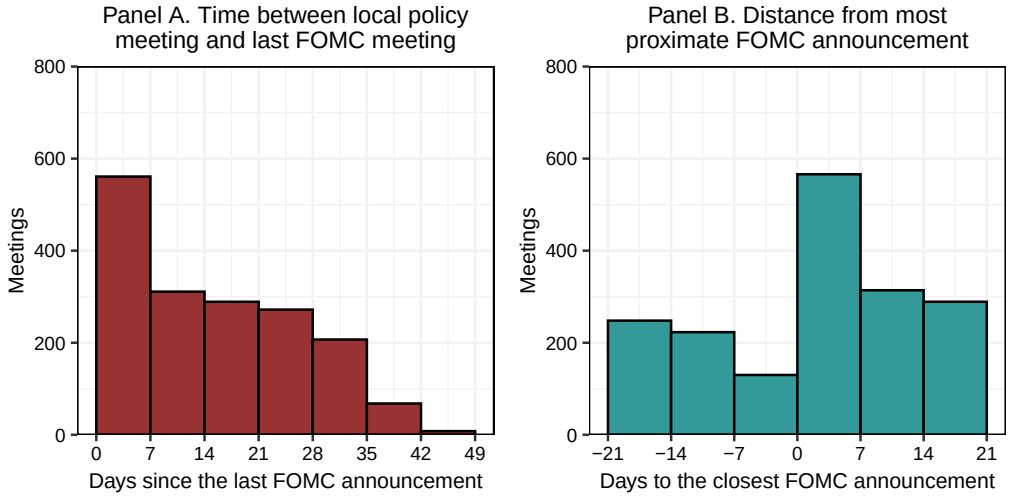


FIGURE B2. BUNCHING OF POLICY MEETINGS: AMERICAS

Notes: This figure shows histograms of the number of days between an FOMC announcement and the next policy meeting of a central bank (Panel A), and of the number of days between a policy meeting and the closest FOMC announcement (Panel B). Panel A includes all meetings, while Panel B restricts the sample to meetings that occurred within three weeks before or after an FOMC announcement. All data is restricted to monetary policy meetings in the Americas.

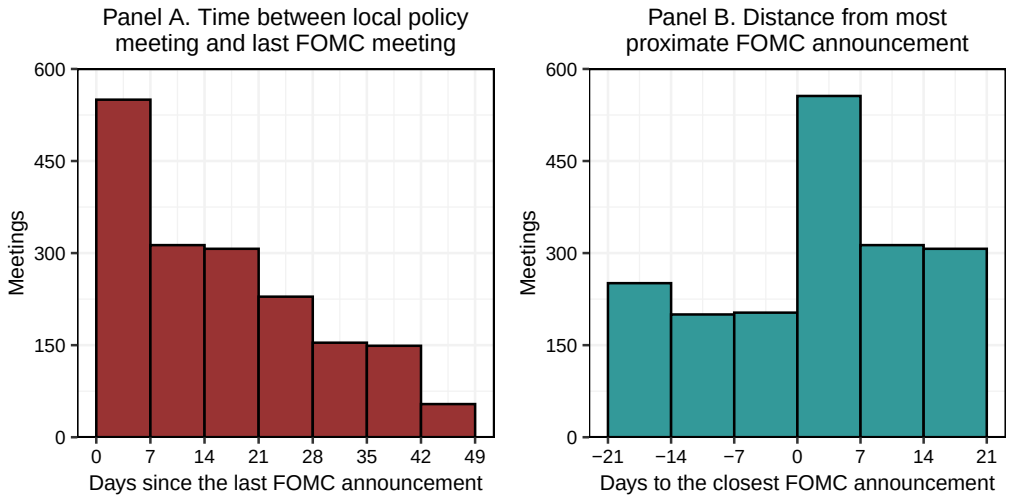


FIGURE B3. BUNCHING OF POLICY MEETINGS: ASIA AND THE PACIFIC

Notes: This figure shows histograms of the number of days between an FOMC announcement and the next policy meeting of a central bank (Panel A), and of the number of days between a policy meeting and the closest FOMC announcement (Panel B). Panel A includes all meetings, while Panel B restricts the sample to meetings that occurred within three weeks before or after an FOMC announcement. All data is restricted to monetary policy meetings in Asia and the Pacific.

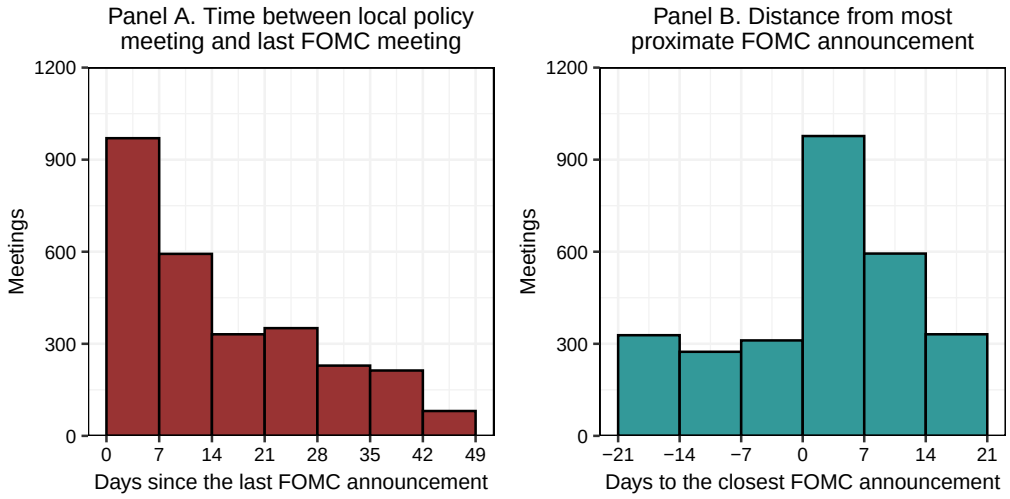


FIGURE B4. BUNCHING OF POLICY MEETINGS: EUROPE

Notes: This figure shows histograms of the number of days between an FOMC announcement and the next policy meeting of a central bank (Panel A), and of the number of days between a policy meeting and the closest FOMC announcement (Panel B). Panel A includes all meetings, while Panel B restricts the sample to meetings that occurred within three weeks before or after an FOMC announcement. All data is restricted to monetary policy meetings in Europe.

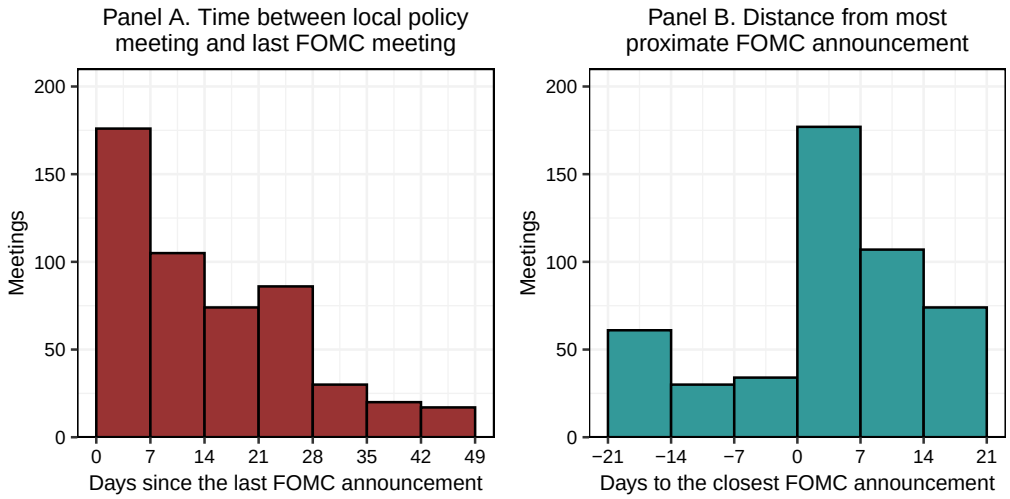


FIGURE B5. BUNCHING OF POLICY MEETINGS: ARAB WORLD

Notes: This figure shows histograms of the number of days between an FOMC announcement and the next policy meeting of a central bank (Panel A), and of the number of days between a policy meeting and the closest FOMC announcement (Panel B). Panel A includes all meetings, while Panel B restricts the sample to meetings that occurred within three weeks before or after an FOMC announcement. All data is restricted to monetary policy meetings in the Arab World.

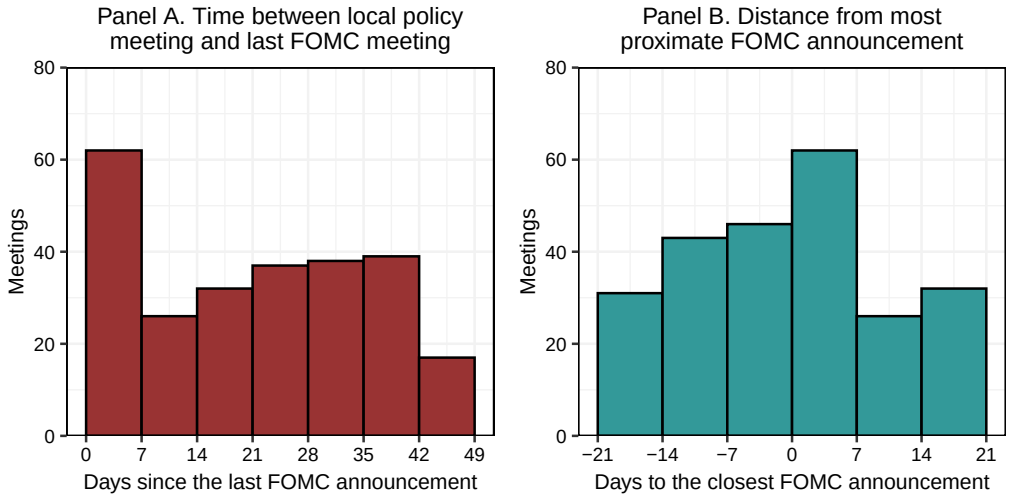


FIGURE B6. BUNCHING OF POLICY MEETINGS: SUB-SAHARAN AFRICA

Notes: This figure shows histograms of the number of days between an FOMC announcement and the next policy meeting of a central bank (Panel A), and of the number of days between a policy meeting and the closest FOMC announcement (Panel B). Panel A includes all meetings, while Panel B restricts the sample to meetings that occurred within three weeks before or after an FOMC announcement. All data is restricted to monetary policy meetings in Sub-Saharan Africa.

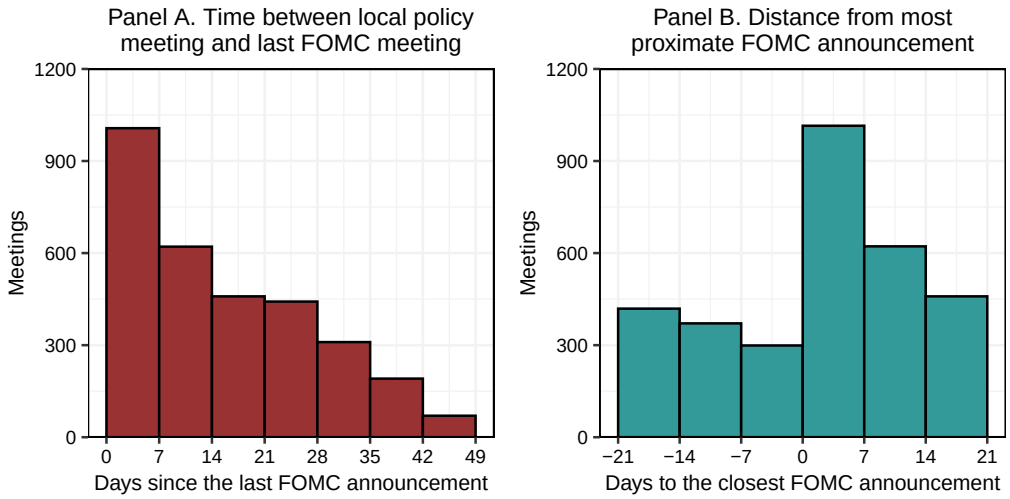


FIGURE B7. BUNCHING OF POLICY MEETINGS: HIGH INCOME ECONOMIES

Notes: This figure shows histograms of the number of days between an FOMC announcement and the next policy meeting of a central bank (Panel A), and of the number of days between a policy meeting and the closest FOMC announcement (Panel B). Panel A includes all meetings, while Panel B restricts the sample to meetings that occurred within three weeks before or after an FOMC announcement. All data is restricted to monetary policy meetings in high income economies.

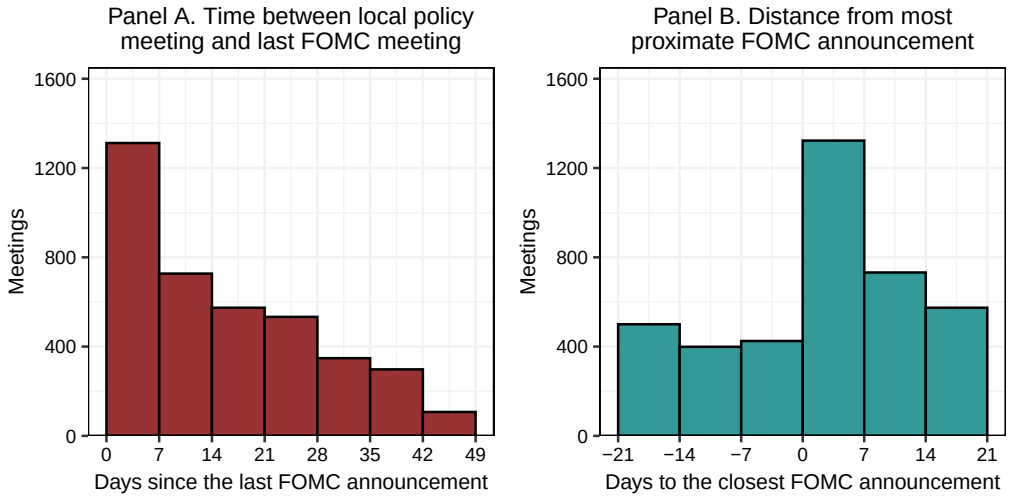


FIGURE B8. BUNCHING OF POLICY MEETINGS: MIDDLE INCOME ECONOMIES

Notes: This figure shows histograms of the number of days between an FOMC announcement and the next policy meeting of a central bank (Panel A), and of the number of days between a policy meeting and the closest FOMC announcement (Panel B). Panel A includes all meetings, while Panel B restricts the sample to meetings that occurred within three weeks before or after an FOMC announcement. All data is restricted to monetary policy meetings in middle income economies.

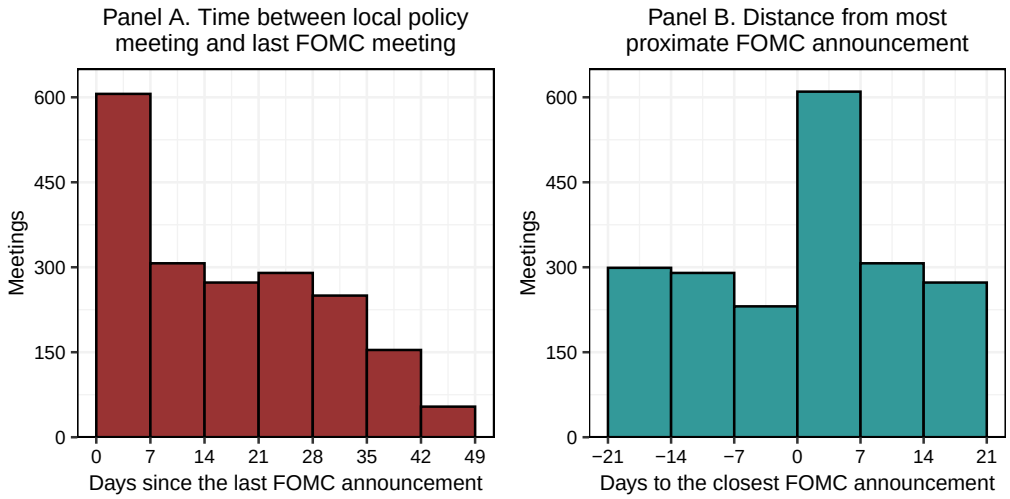


FIGURE B9. BUNCHING OF POLICY MEETINGS: ADVANCED ECONOMIES

Notes: This figure shows histograms of the number of days between an FOMC announcement and the next policy meeting of a central bank (Panel A), and of the number of days between a policy meeting and the closest FOMC announcement (Panel B). Panel A includes all meetings, while Panel B restricts the sample to meetings that occurred within three weeks before or after an FOMC announcement. All data is restricted to monetary policy meetings in advanced economies.

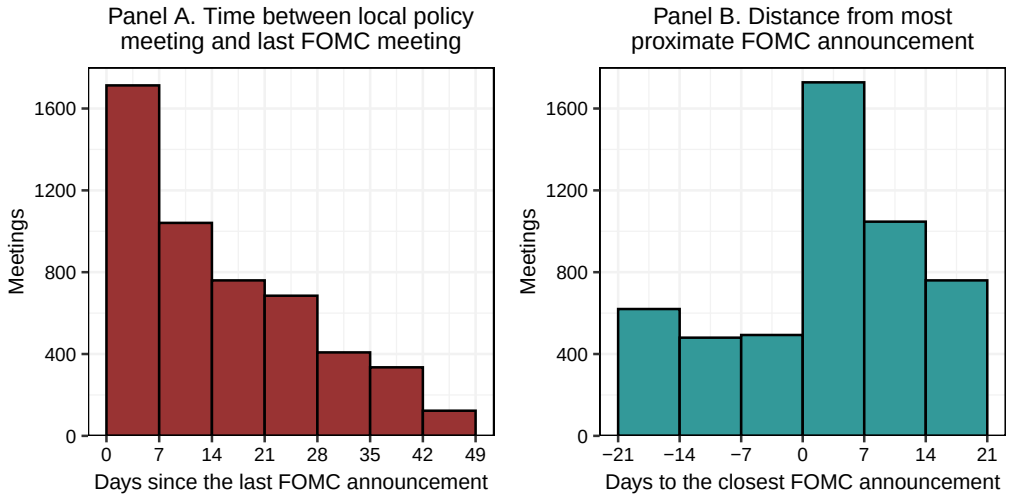


FIGURE B10. BUNCHING OF POLICY MEETINGS: EMERGING ECONOMIES

Notes: This figure shows histograms of the number of days between an FOMC announcement and the next policy meeting of a central bank (Panel A), and of the number of days between a policy meeting and the closest FOMC announcement (Panel B). Panel A includes all meetings, while Panel B restricts the sample to meetings that occurred within three weeks before or after an FOMC announcement. All data is restricted to monetary policy meetings in emerging economies.

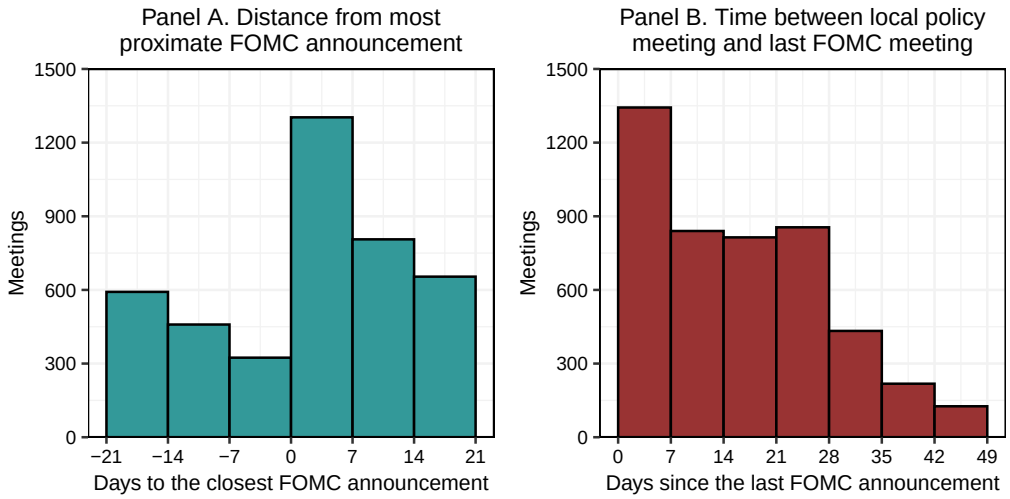


FIGURE B11. BUNCHING OF POLICY MEETINGS: NO MEETINGS AROUND GDP RELEASE DATES

Notes: This figure shows histograms of the number of days between an FOMC announcement and the next policy meeting of a central bank (Panel A), and of the number of days between a policy meeting and the closest FOMC announcement (Panel B). Panel A includes all meetings, while Panel B restricts the sample to meetings that occurred within three weeks before or after an FOMC announcement. The data excludes meetings occurring between three and seven weeks after the end of a quarter.

B.1.2 Heterogeneity

To assess the extent of heterogeneity in the scheduling of policy meetings relative to FOMC announcements, we proceed as follows. For each central bank c, we examine its policy meetings and assign each to one of six intervals based on the number of weeks relative to the nearest FOMC announcement: $I_{-3} \equiv [-3, -2)$, $I_{-2} \equiv [-2, -1)$, $I_{-1} \equiv [-1, 0)$, $I_1 \equiv [0, 1]$, $I_2 \equiv (1, 2]$, and $I_3 \equiv (2, 3]$.[33] We then classify central bank c as scheduling most of its policy meetings $w > 0$ ($w < 0$) weeks after (before) an FOMC announcements if I_w is the interval with highest mass among I_{-3}, ..., I_3. In simple words, if country c is represented by a histogram like the one depicted in Figure 1, we would classify it as scheduling most of its policy meetings in the week following an FOMC announcement.

Figure B12 shows the resulting category for each central bank in the sample. Emerging economies are labeled under their geographic region, while advanced economies are clustered in their own group. A black vertical line separates central banks that tend to schedule their policy meetings before the Fed's decision on the funds rate from those that wait for the Fed. For forty-five out of the fifty-eight non-U.S. central banks (i.e., 78%), the majority of policy

[33] To avoid double counting, we drop meetings that fall simultaneously within three weeks both before and after an FOMC announcement, which leaves us with 88% of all policy meetings.

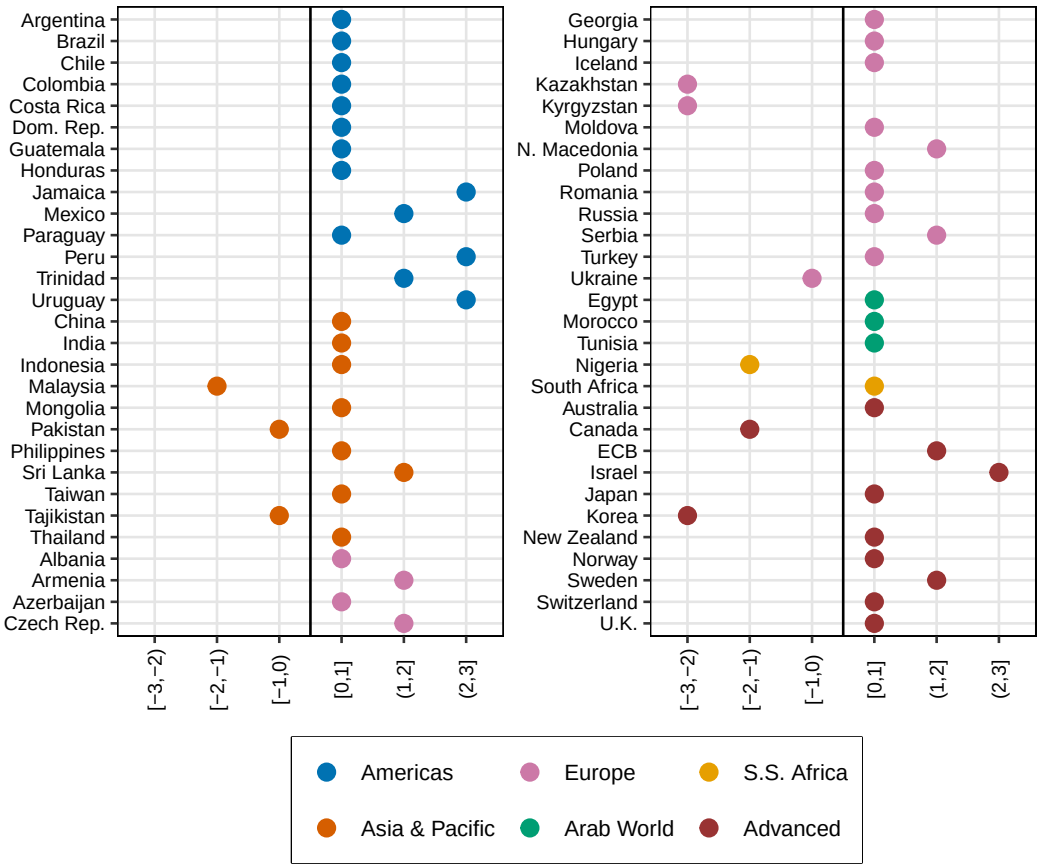


FIGURE B12. RELATIVE DISTANCE TO FOMC ANNOUNCEMENTS

Notes: This figure shows, in weeks relative to FOMC announcements, when most policy meetings are scheduled by each central bank in the MPM dataset. The black vertical line separates central banks that concentrate their meetings before the Fed's decision on the funds rate from those that wait for the Fed.

meetings following a FOMC announcement are within two weeks after the respective FOMC announcement.

B.2 Immediate Adjustments After FOMC Announcements

B.2.1 Aggregation

Why does granularity matter for our results? In the main text we argue that when policy rate data are aggregated at monthly frequency, it will often be the case that the Fed meets in the second half of the month (recall that two-thirds of FOMC meetings during our sample period took place in the second half of the month), so other central banks are less likely to meet in those months. Under this scenario, aggregated data would imply an attenuated contemporaneous effect of U.S. monetary policy on its non-U.S. counterparts.

To provide evidence that this claim is a reasonable explanation for why previous papers using aggregated monthly data do not find a significant contemporaneous causal effect of U.S. monetary surprises on non-U.S. central banks' policy rates, consider the regression equation:

$$\Delta i_{c,t} = \alpha_c + \beta \cdot \Delta i_t^{US} + \gamma' \boldsymbol{X}_{c,t} + e_{c,t},$$

where the notation is analogous to that used in specification (1). That is, $\Delta i_{c,t}$ and Δi_t^{US} are the changes in the end-of-the-month policy rate and Fed funds rate; $\boldsymbol{X}_{c,t}$ is a vector of controls including the change in inflation and GDP growth; and $e_{c,t}$ are the error terms.

If our explanation for why aggregated data deliver imprecise and attenuated estimates for β is reasonable, it must be the case that estimates of β using only the aggregated country-month observations from when the Fed conducted its meetings in the first 10 days of the month should be more precise and less noisy than those using their counterparts from the last 10 days. The rationale is that central banks are more likely to meet in the same month when the FOMC holds a meeting at the start of the month than at the end of the month, so estimates using the first sub-sample must capture a majority of the first policy decisions made by other central banks following the FOMC announcements.

Table B1 shows the estimates of β for the two subsamples using the Bauer and Swanson (2023b) shocks as an instrument for Δi_t^{US}. Panel A shows that when the Fed meets in the first 10 days of the month, using aggregate data delivers a β close to 0.55, as in our baseline results. There is a slight attenuation effect in the coefficient, but it is highly statistically significant. On the other hand, estimates of β using data from months when the Fed conducted its meetings in the last 10 days of the month are extremely noisy and have a negative sign.

B.2.2 Robustness to Identifying Assumption and Outlier Exclusion

We verify that the baseline results presented in Table 2 are not sensitive either to the inclusion of lags of the dependent variable, the identification assumption or to the choice of measuring changes in the log-gross rate as the dependent variable. Specifically, we proceed as follows for each case:

i) *Adding lags of the dependent variable.* We re-estimate equation (1) by including $p \in 1, 3, 6$ lags in the vector of controls. Table B2 reports the corresponding estimates of β. Overall, the estimates are slightly more conservative than in the baseline specification: an exogenous 100 basis points increase in the Fed funds rate is associated with a 40 basis points increase in non-U.S. policy rates. In the remainder of this section, we report estimates of β using six lags of the dependent variable.

TABLE B1. CONTEMPORANEOUS RELATION BETWEEN POLICY RATES AND THE FED FUNDS RATE

	Policy rate decision $\Delta i_{c,\tau+\epsilon}$		
	All	Emerging	Advanced
Panel A. Months with FOMC meetings in days 1–10			
Δi_{τ}^{US}	0.487	0.450***	0.045
	(0.300)	(0.166)	(1.673)
Observations	1,425	1,054	371
F-stat (1st stage)	55.5	112.5	0.844
Panel B. Months with FOMC meetings in days 21—			
Δi_{τ}	-0.626	-0.559	-1.012
	(1.324)	(1.041)	(3.244)
Observations	2,644	1,984	660
F-stat (1st stage)	24.7	23.3	2.20

Notes: The table reports the coefficient β of the monthly version of equation (1) estimated using U.S. monetary policy shocks à la Bauer and Swanson (2023b) as instruments for the Fed funds rate. The vector of controls includes country fixed effects and the most recent change in annual CPI inflation and output growth. Driscoll and Kraay (1998) standard errors are reported in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

ii) *Placebo test.* The key identification assumption of our empirical design is that the monetary shock series used as an instrument for the Fed funds rate is truly exogenous. By construction, the Bauer and Swanson (2023b) monetary shock series is not autocorrelated and removes the predictable component associated with well-known macroeconomic fundamentals. As they argue, their series is therefore very likely to be exogenous. Nevertheless, a potential threat to identification arises if the shocks identified by Bauer and Swanson are not truly exogenous —specifically, if central banks' forecasts outperform market expectations in predicting the Fed's actions. To address this concern, we estimate the following placebo regression:

$$\Delta i_{c,\tau-\epsilon} = \alpha_c + \beta \cdot \Delta i_{\tau}^{US} + \gamma' \boldsymbol{X}_{c,\tau-\epsilon} + e_{c,\tau-\epsilon}, \tag{B3}$$

which is equivalent to equation (1), with the important difference that the local meetings now correspond to the last meetings of central banks before an FOMC announcement (note that we now use $\tau - \epsilon$ instead of $\tau + \epsilon$). If $\beta = 0$, this would indicate that central

TABLE B2. MPR DECISIONS FOLLOWING AN FOMC ANNOUNCEMENT, LAGS INCLUDED

	Policy rate decision $\Delta i_{c,\tau+\epsilon}$		
	All	Emerging	Advanced
Panel A. $p = 1$ *lags*			
Δi_{τ}^{US}	0.284**	0.401**	-0.097
	(0.136)	(0.163)	(0.336)
Observations	6,000	4,575	1,425
F-stat (1st stage)	172.9	126.7	41.7
Panel B. $p = 3$ *lags*			
Δi_{τ}	0.261**	0.387**	-0.159
	(0.129)	(0.166)	(0.380)
Observations	5,931	4,516	1,415
F-stat (1st stage)	167.9	122.3	38.7
Panel C. $p = 6$ *lags*			
Δi_{τ}	0.258*	0.398**	-0.178
	(0.137)	(0.179)	(0.388)
Observations	5,826	4,426	1,400
F-stat (1st stage)	156.4	111.5	38.8

Notes: The table reports the coefficient β of the monthly version of equation (1) estimated using U.S. monetary policy shocks à la Bauer and Swanson (2023b) as instruments for the Fed funds rate. The vector of control variables $\boldsymbol{X}$ includes country fixed effects, the most recent change in annual CPI inflation and output growth, and the policy rate changes announced after the previous $p \in \{1, 3, 6\}$ policy meetings. Driscoll and Kraay (1998) standard errors are reported in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

banks are not anticipating the exogenous changes in the Fed funds rate proxied by the Bauer and Swanson shocks.[34] Table B3 shows that the null hypothesis $\beta = 0$ cannot be rejected for both emerging and advanced economy central banks (p-values around 80%),

[34] An alternative interpretation is that $\beta = 0$ because central banks anticipate, at least partially, upcoming U.S. monetary shocks but choose not to react before the FOMC announcement. However, such a scenario would require an explanation for why central banks decide to act only after the Fed's policy decision is made public.

confirming that central banks are not foreseeing the monetary shocks as measured by the Bauer and Swanson series.

Table B3. Policy Rates Decisions Before an FOMC Announcement

	Policy rate decision $\Delta i_{c,\tau-\epsilon}$		
	All	Emerging	Advanced
Δi_τ	0.069	0.077	-0.027
	(0.124)	(0.165)	(0.238)
Observations	5,781	4,384	1,397
F-stat (1st stage)	134.5	90.6	39.4

Notes: The table reports the coefficient β of equation (B3) estimated using U.S. monetary policy shocks à la Bauer and Swanson (2023b) as instruments for the Fed funds rate. The vector of control variables $\boldsymbol{X}$ includes country fixed effects, the most recent change in annual CPI inflation and output growth, and the policy rate changes announced after the previous six policy meetings. Driscoll and Kraay (1998) standard errors are reported in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

iii) *No trimming of outliers.* For our baseline results, we exclude observations in which the change in the policy rate is below the 1st percentile or above the 99th percentile. This choice helps address outliers. Table B4 reports the estimates of equation (1)'s β including observations for which the policy rate exceeds 20%. The magnitude of these alternative estimates is similar to that of the baseline.

Table B4. MPR Decisions Following an FOMC Announcement, Outliers Included

	Policy rate decision $\Delta i_{c,\tau+\epsilon}$		
	All	Emerging	Advanced
Δi_τ	0.570***	0.732***	0.170
	(0.157)	(0.220)	(0.163)
Observations	6,108	4,678	1,430
F-stat (1st stage)	189.3	137.0	50.1

Notes: The table reports the coefficient β of equation (1) estimated using U.S. monetary policy shocks à la Bauer and Swanson (2023b) as instruments for the Fed funds rate. The vector of control variables $\boldsymbol{X}$ includes country fixed effects, the most recent change in annual CPI inflation and output growth, and the policy rate changes announced after the previous six policy meetings. Driscoll and Kraay (1998) standard errors are reported in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

These results confirm that our findings are not driven by spurious co-movements between U.S. monetary shocks and other shocks, nor by the exclusion of country-meeting pairs for which the policy rate was an outlier.

B.3 Dynamic Response of Policy Rates

B.3.1 Impulse-Response Functions, LP-IV estimates

We estimate the following local projection equation:

$$i_{c,\tau+h} = \alpha_c^h + \beta^h \cdot \Delta i_\tau^{US} + \boldsymbol{\gamma}^h \cdot \boldsymbol{X}'_{c,\tau} + e_{c,\tau+h}, \quad h = -240, ..., 360. \tag{B4}$$

This specification closely mirrors equation (2), with the key difference that the treatment variable is now the change in the Fed funds rate, Δi_τ^{US}, which we instrument using the SB monetary policy shock series.

Figure B13 reports the estimated responses across horizons. Overall, the results are very similar to those obtained under our baseline specification.

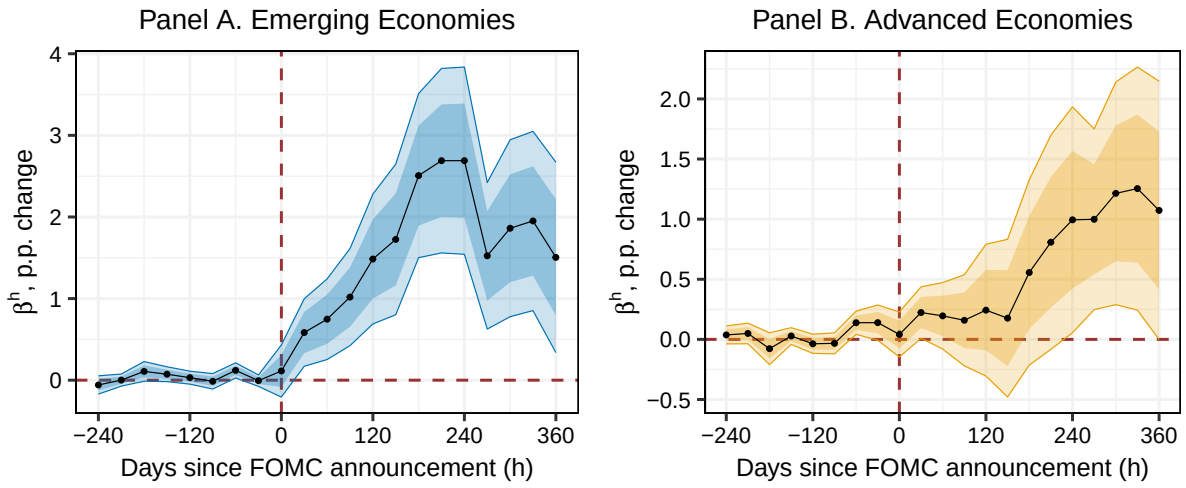


FIGURE B13. RESPONSE OF POLICY RATES TO A U.S. MONETARY SHOCK, LP-IV

Notes: This figure reports the estimates of the β^h in equation (B4) for emerging economies (Panel A) and advanced economies (Panel B). For the estimation procedure we use the Bauer and Swanson (2023b) monetary shock series as an instrument for the Fed funds rate. The light (dark) shaded areas represent the 90% (68%) symmetric confidence intervals, constructed using lag-augmented heteroskedasticity-corrected standard errors.

B.3.2 Impulse-Response Functions with outliers

To understand how central bankers respond to U.S. monetary shocks at all times — including *times of crisis* — we re-estimate equation (2):

$$i_{c,\tau+h} = \alpha_c^h + \beta^h \cdot MPS_\tau^{US} + \boldsymbol{\gamma}^{h\prime} \cdot \boldsymbol{X}_{c,\tau} + e_{c,\tau+h}, \quad h = -120, -90, ..., 360.$$

including observations $i_{c,\tau+h}$ associated with FOMC meetings τ that were held between July 2008–June 2009 and January 2020–December 2020.

We present the results of this exercise in Figure B14. Relative to our baseline, the response function of emerging economy policy rates still indicates that emerging economy central banks adjust rates upward following a U.S. monetary policy shock for at least four months, although the response is weaker. On the other hand, the non-significant response of advanced economy central banks observed in the baseline disappears: once U.S. monetary policy shocks from crises (the Global Financial Crisis and the COVID pandemic) are included in the analysis, advanced economy central banks appear to lean against the Fed.

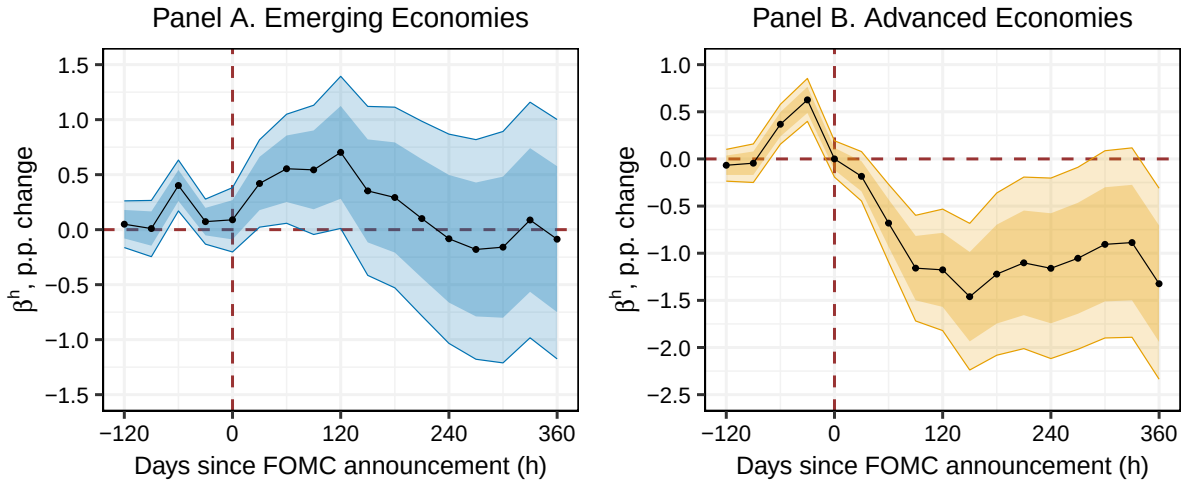


FIGURE B14. RESPONSE OF POLICY RATES TO A U.S. MONETARY SHOCK, ALL DATA

Notes: See text for details. The light (dark) shaded areas represent the 90% (68%) symmetric confidence intervals, constructed using lag-augmented heteroskedasticity-corrected standard errors.

What drives these differences? In short, the Global Financial Crisis. To show this, we re-estimate the local projection equation (2) on two sub-samples: *i)* one excluding the peak of the GFC (July 2008–June 2009); and *ii)* one excluding the start of the COVID pandemic (2020). Excluding the GFC yields results very similar to the baseline, with slightly larger positive responses (see Figure B15), indicating that our baseline results are robust to the influence of the COVID pandemic. In contrast, when the COVID episode is excluded and the

GFC episode is included, emerging economies still raise policy rates after a Fed tightening, although only mildly, whereas advanced economy central banks ease their policy rates (see Figure B16). Overall, the baseline results hold well for emerging economies but not for advanced economies, largely reflecting the influence of the GFC —an outlier event.

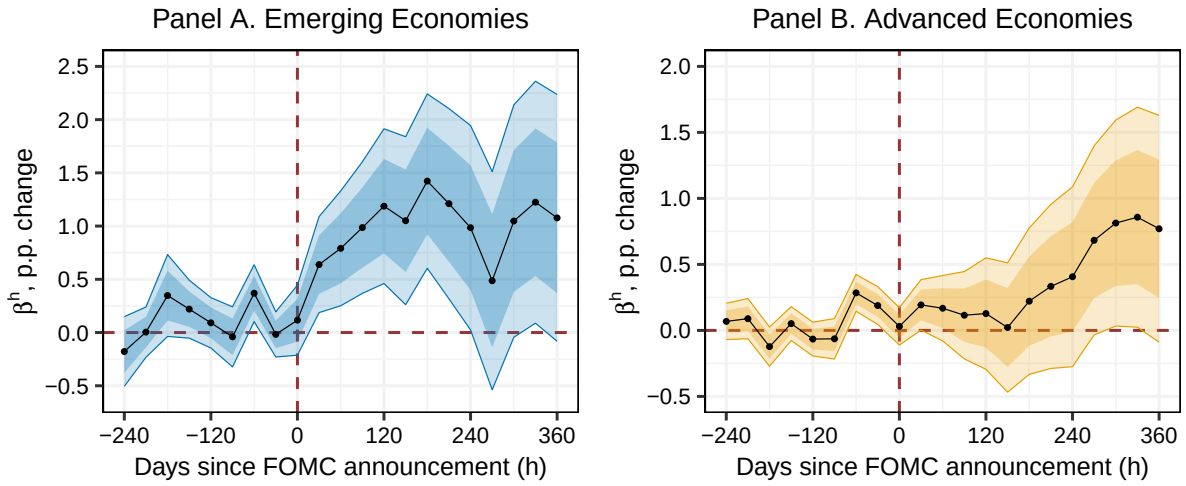


FIGURE B15. RESPONSE OF POLICY RATES TO A U.S. MONETARY SHOCK, NO GFC INFLUENCE

Notes: See text for details. The light (dark) shaded areas represent the 90% (68%) symmetric confidence intervals, constructed using lag-augmented heteroskedasticity-corrected standard errors.

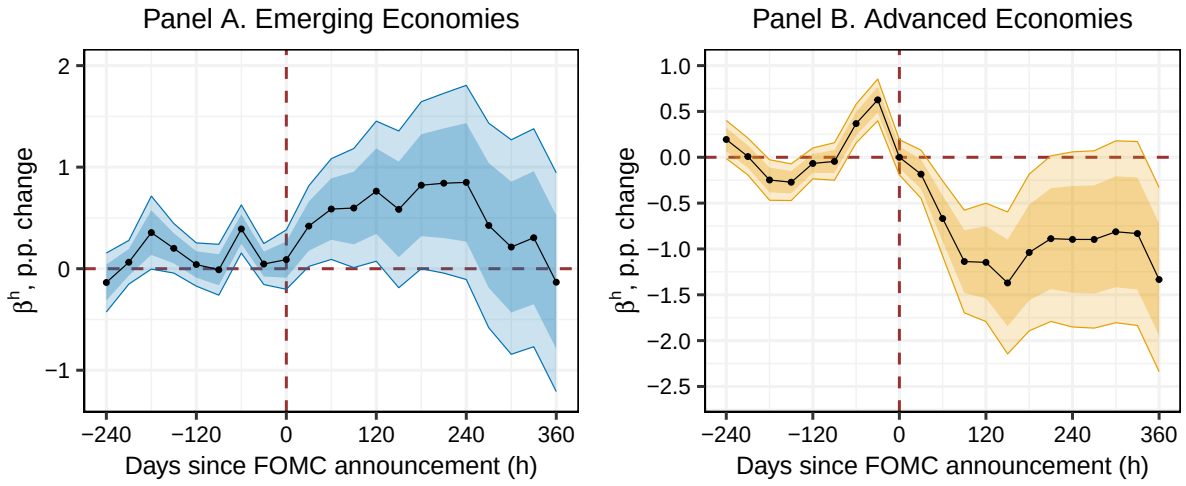


FIGURE B16. RESPONSE OF POLICY RATES TO A U.S. MONETARY SHOCK, NO COVID INFLUENCE

Notes: See text for details. The light (dark) shaded areas represent the 90% (68%) symmetric confidence intervals, constructed using lag-augmented heteroskedasticity-corrected standard errors.

B.3.3 Impulse-Response Functions without ZLB

We estimate equation (2) as in the baseline specification, but exclude observations from the zero lower bound (ZLB) period (January 2009–December 2015). Figure B17 presents the estimated response functions. Overall, the results for both emerging and advanced economies are very similar to the baseline estimates. These findings indicate that the ZLB period does not drive our baseline results.

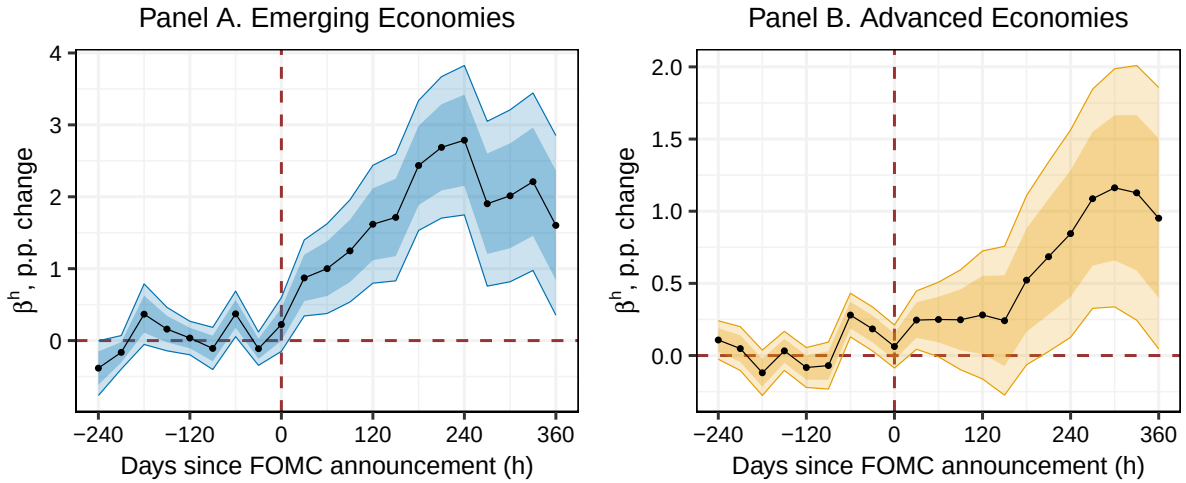


FIGURE B17. RESPONSE OF POLICY RATES TO A U.S. MONETARY SHOCK, NO ZLB

Notes: See text for details. The light (dark) shaded areas represent the 90% (68%) symmetric confidence intervals, constructed using lag-augmented heteroskedasticity-corrected standard errors.

B.4 Additional results

B.4.1 Central Banks' Response to the ECB

Are central banks changing their policy rates in the meeting scheduled after a policy rate announcement of the Government Council of the ECB (GC-ECB)? To provide a question to this question, we do an exercise in the spirit of section 3.2. We estimate regressions of the form

$$\Delta i_{c,\tau+\epsilon} = \alpha_c + \beta \cdot \Delta i_{\tau}^{ECB} + \gamma' \boldsymbol{X}_{c,\tau} + e_{c,\tau+\epsilon}, \tag{B5}$$

which differ from specification (1) in three aspects. First, we use the change in ECB's Deposit Facility Rate (DFR) Δi_{τ}^{ECB} as the treatment variable instead of Δ_{τ}^{US} as the treatment variable. Second, τ now indices the dates of GC-ECB announcements. Third, β now represents the average change in central banks' policy rates decided in the meeting following an announcement by the GC-ECB.

Table B5. Policy Rates Decisions Following an GC-ECB Announcement

	Policy rate decision $\Delta i_{c,\tau+\epsilon}$					
	All		Emerging		Advanced	
	(1)	(2)	(1)	(2)	(1)	(2)
Δi_{τ}^{ECB}	-0.051	-0.114	-0.117	-0.189	0.117	0.129
	(0.178)	(0.164)	(0.236)	(0.209)	(0.177)	(0.184)
Country FE	×	✓	×	✓	×	✓
Controls	×	✓	×	✓	×	✓
Observations	7,286	7,067	5,169	4,980	2,117	2,087
First stage F-stat	210.0	185.4	151.3	139.3	82.9	72.2

Notes. The table reports the coefficient β of equation (1) estimated using ordinary least squares (Panel A) and using ECB monetary policy shocks à la Jarociński and Karadi (2020) instruments for the ECB Deposits Facility Rate (Panels B) for different subsamples and two specifications of the vector of control variables $\boldsymbol{X}$. Specification (1) only includes an intercept, while specification (2) adds the most recent change in annual CPI inflation, output growth, and the first lag of the dependent variable. Heteroskedasticity-corrected standard errors are reported in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

To estimate β we use the monetary shock series of Jarociński and Karadi (2020) as an instrument. Table B5 reports β estimated without controls (specification 1) and with Taylor rule controls (specification 2) using the data of all economies, and from emerging and advanced economies separately. We estimate β to be very close to zero in all three cases and far from being statistically significant. In sum, these estimates indicate that central banks do not react to ECB surprise changes in their policy rates very fast.

B.4.2 Is the Fed Influenced by Other Central Banks?

To assess whether the Fed is influenced by the decisions of other central banks, we estimate the following equation separately for each central bank c:

$$\Delta i_{\tau}^{US} = \alpha_c + \theta_c \cdot \Delta i_{c,\tau-\epsilon} + \gamma_c' \cdot \boldsymbol{X}_{\tau-\epsilon}^{US} + \nu_{c,\tau}, \tag{B6}$$

where θ_c captures the effect of central bank c's most recent policy rate decision ($\Delta i_{c,\tau-\epsilon}$) on the Fed's policy rate change (Δi_{τ}^{US}).[35] This specification mirrors equation (1), with the key difference being that the dependent and explanatory variables are reversed. Another difference is that the control vector now includes changes in U.S. inflation and GDP growth

[35] The variables $\Delta i_{c,\tau-\epsilon}$ are standardized. Since standardization is a linear transformation, it does not affect the interpretation of the statistical significance of θ_c.

observable to the Fed at the time of its policy announcement, as well as a lag of the Fed funds rate.

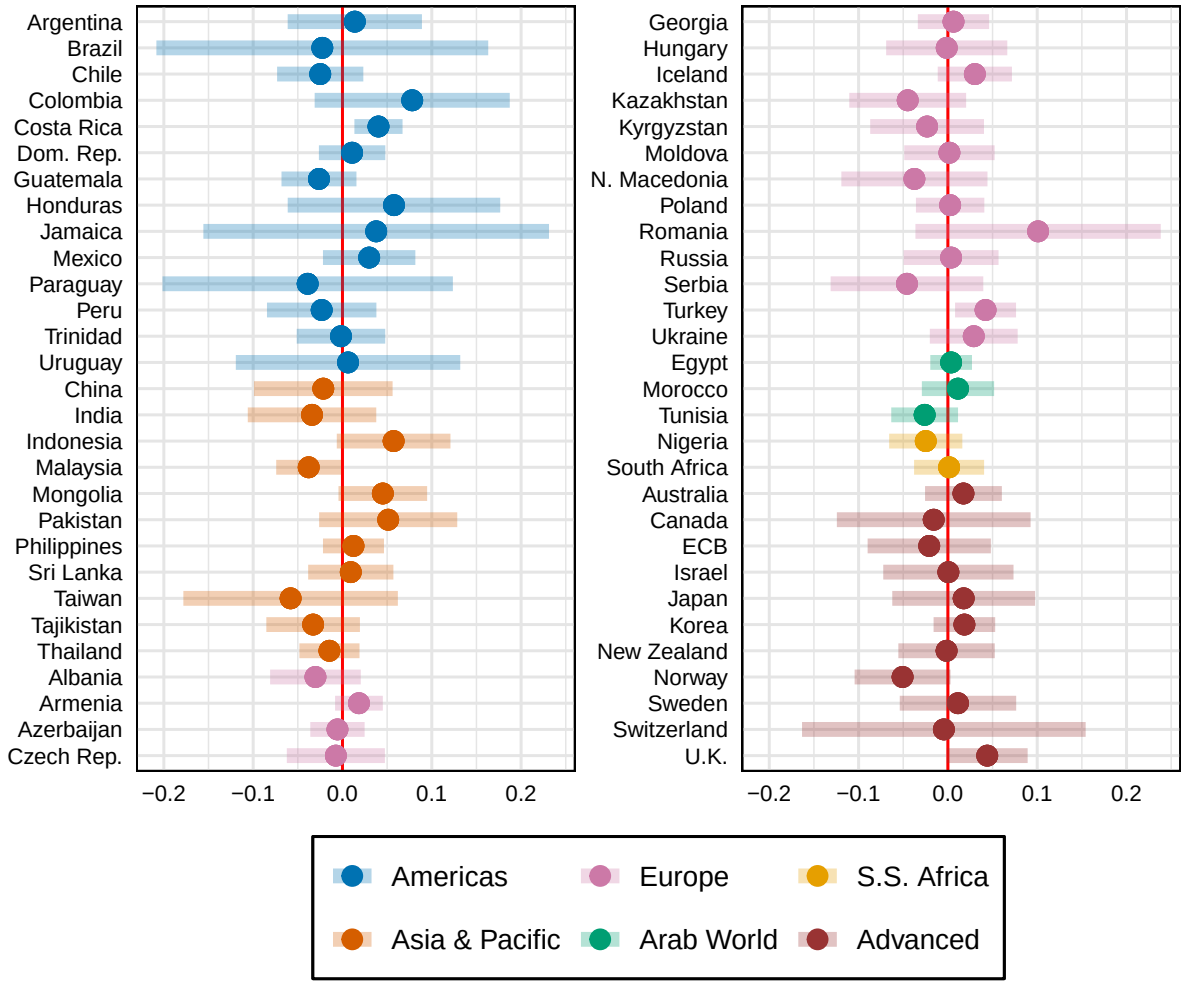


FIGURE B18. FED RESPONSE TO OTHER CENTRAL BANKS

Notes: This figure reports the estimated coefficients θ_c in equation (B6) and their 95% confidence intervals (light shaded areas). The coefficients measure the average effect of a one-standard-deviation change in the policy rate of country c on the policy decision of the Fed.

We compute IV estimates of θ_c using monetary policy shock series for country c as instruments.[36] Figure B18 presents the IV estimates of θ_c and their 95% confidence intervals (light shaded areas). Of the fifty-seven coefficients, 95% (i.e., 54 central banks) are not statistically

[36]For most countries, there is no direct equivalent to the monetary shock series of Nakamura and Steinsson (2018) or Bauer and Swanson (2023b). To address this limitation, we construct a proxy by first estimating residuals from policy rules à la Taylor (1993), and then extracting the common global component from those residuals using a dynamic factor model. We interpret the residuals from this factor model as the country-specific monetary shocks. Full details are provided in Appendix A.2.

significant at the 5% level. For the three cases where θ_c is statistically significant —Costa Rica, Malaysia, and Turkey— it is difficult to think of a plausible explanation for the Fed systematically reacting to these economies. The rejection in these cases is likely due to a Type I error in hypothesis testing; after all, if the null hypothesis is true, a 5% significance level implies a 5% chance of false rejection.

This fact holds broadly and applies to central banks like the ECB, the Bank of England, or the People's Bank of China. This is in line with previous evidence on the limited impact of monetary shocks from these central banks on U.S. macroeconomic variables (Ca'Zorzi et al., 2020; Miranda-Agrippino & Rey, 2022).

B.4.3 Post-Covid Tightening Cycle

During the post-Covid tightening cycle many EMEs have been known to tighten ahead of the Fed, see (Mendoza-Fernández et al., 2026; Nakamura et al., 2025). Figure B19 presents the average policy rates in EMEs and AEs along with the Fed funds rate and various fundamental variables. The figure shows that average EM policy rates broadly followed the U.S. dollar's appreciation against advanced economies, as well as the expected hikes in the Fed funds rate. The oil prices was a less consistent driver of EM hikes in this period. Therefore, EM central banks moved ahead of the Fed in order to tame private sector's inflation expectations.

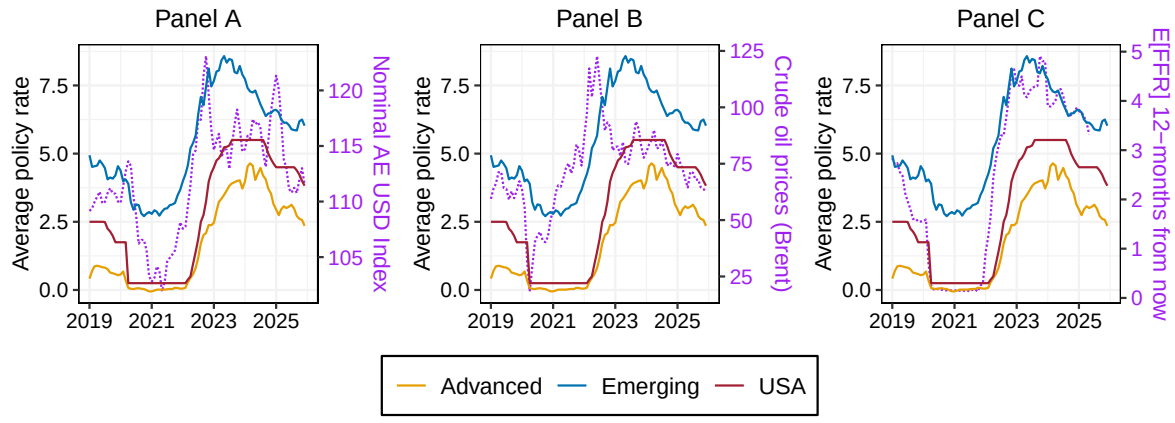


FIGURE B19. POLICY RATES AND FUNDAMENTALS, 2019-2025

Notes: The nominal US dollar index for advanced economies and European Brent oil prices are from the Federal Reserve Economic Data (FRED). Expectations for the federal funds rate are proxied by expectations of short-term interest rates 12 months ahead from Consensus Economics.

The findings from this period is highly consistent with our IV estimates on following the Fed. Monetary policy shocks based on high-frequency identification methods contain news

about both the change in the current target rate as well as the change in the expected policy path. Therefore, monetary policy shocks in 2021 about expected Fed hikes in 2022 can trigger EMEs to follow ahead. Overall, we conclude that this period is a perfect example of EME central bankers following the Fed, both in changes in the current fed funds rate and the expected policy path.

C Appendix to Section 4

C.1 Expectations

C.1.1 Forecast Revisions for the US economy

In section 4, we show that changes in the Fed funds rate significantly change forecasters expectations on non-U.S. economies. One possibility is that changes in US monetary policy has direct effects on expectations on the US economy and demand which can spillover to foreign economies. In Table C1 we present the revisions of US macroeconomic variables around an FOMC announcement. There is no significant change in US expectations following a Fed tightening and therefore we rule out the possibility that our results are driven by the demand channel.

Table C1. Forecast Revisions around FOMC Announcements - USA

	Forecast revisions for:			
	12-month inflation $\mathbb{E}_m[\pi_{US,12}]$	12-month growth $\mathbb{E}_m[\Delta y_{US,12}]$	12-month inflation $\mathbb{E}_m[\pi_{US,12}]$	12-month growth $\Delta\mathbb{E}_m[y_{US,12}]$
	(1)	(2)	(3)	(4)
ΔFFR_m	-0.254	-0.132	0.070	0.013
	(0.267)	(0.218)	(0.129)	(0.134)
Observations	216	222	216	222
F-test (1st stage)	13.4	18.8	38.5	53.6
IV	BS	BS	NS	NS

Notes: The table reports the coefficient β of equation (3) estimated using U.S. monetary policy shocks à la Nakamura and Steinsson (2018) and Bauer and Swanson (2023b) instruments for the Fed funds rate. Robust standard errors are in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

C.1.2 Robustness to alternative monetary series

In section 4, we show that changes in the Fed funds rate significantly change forecasters expectations in EMEs. To confirm such results are not driven by our choice of instrument for monetary policy, we estimate the results also with the instrument of Nakamura and Steinsson (2018). Table C2 presents the results. We confirm that our main results are robust to the alter-

Table C2. Forecast Revisions around FOMC Announcements

	Forecast revisions for:			
	12-month inflation $\mathbb{E}[\pi_{c,12}]$	3-month depreciation $\mathbb{E}[\Delta e_{c,3}]$	12-month depreciation $\mathbb{E}[\Delta e_{c,12}]$	12-month growth $\mathbb{E}[\Delta y_{c,12}]$
	(1)	(2)	(3)	(4)
Δi_{τ}^{US}	0.409***	0.023	-0.300	0.096**
	(0.077)	(0.626)	(0.597)	(0.044)
Observations	6,046	3,413	3,373	6,243
F-test (1st stage)	892.5	640.9	655.0	849.1

Notes: The table reports the coefficient β of equation (3) estimated using U.S. monetary policy shocks à la Nakamura and Steinsson (2018) as the instrument for the Fed funds rate. Lag-augmented heteroskedasticity-corrected standard errors are in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

nate instrument: expectations on inflation and depreciation are revised upwards in emerging economies.

We have additionally shown that changes in the Fed funds rate significantly change forecasters inflation expectations on non-U.S. economies across horizons. Forecasters expect inflation to rise over time following an unexpected Fed tightening. To confirm such results are not driven by our choice of instrumental variable, we instrument for the Fed funds rate with the Nakamura and Steinsson (2018) shocks. Figure C20 presents the results. We confirm that our main results are robust to the alternate instruments.

C.1.3 Robustness to the Inclusion of Outliers

In section 4, we show that changes in the Fed funds rate significantly change forecasters expectations on non-U.S. economies. To confirm such results are not driven by our choice of removal of outliers in the forecast revisions, we re-estimate our main results including all observations. Table C3 presents the results. We confirm that our main results are robust to the inclusion of outliers. In fact, outliers strengthen the magnitude and significance of our results especially with regards to emerging economies.

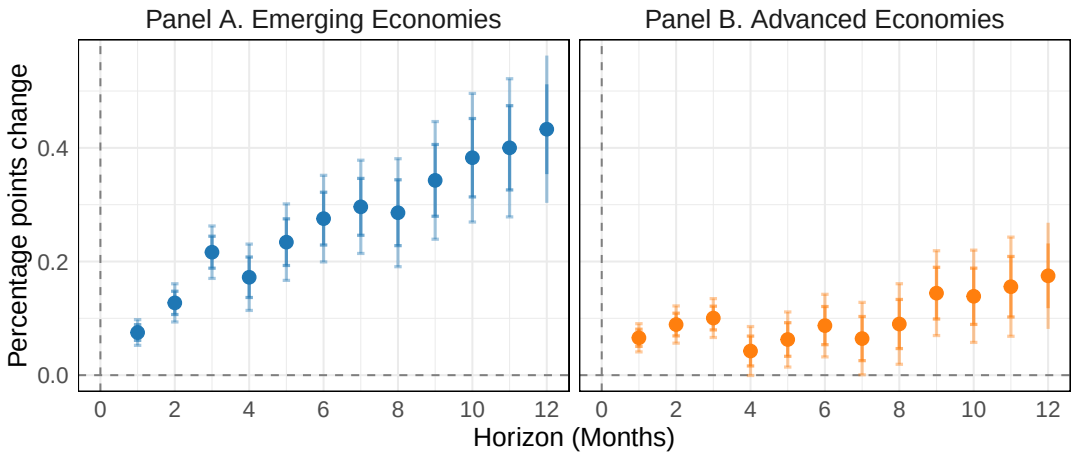


FIGURE C20. PASS-THROUGH TO INFLATION EXPECTATIONS ACROSS HORIZONS

Notes: The pass-through is defined as the estimated coefficient β of equation (3). The light (dark) shaded areas represent the 90% (68%) symmetric confidence intervals, constructed using lag-augmented heteroskedasticity-corrected standard errors. U.S. monetary policy shocks à la Nakamura and Steinsson (2018) as the instrument for the Fed funds rate.

TABLE C3. FORECAST REVISIONS AROUND FOMC ANNOUNCEMENTS

	Forecast revisions for:			
	12-month inflation $\mathbb{E}[\pi_{c,12}]$	3-month depreciation $\mathbb{E}[\Delta e_{c,3}]$	12-month depreciation $\mathbb{E}[\Delta e_{c,12}]$	12-month growth $\mathbb{E}[\Delta y_{c,12}]$
	(1)	(2)	(3)	(4)
Panel A. Emerging Economies				
Δi_τ^{US}	1.052**	4.333***	3.806***	-0.386*
	(0.458)	(1.397)	(1.349)	(0.231)
Observations	6,588	4,287	4,282	6,822
F-stat (1st stage)	215.5	259.7	259.8	237.4
Panel B. Advanced Economies				
Δi_τ^{US}	0.047	4.605***	2.343**	-0.093
	(0.099)	(1.531)	(1.097)	(0.082)
Observations	1,593	1,832	1,832	1,712
F-stat (1st stage)	76.5	115.4	115.8	115.0

Notes: The table reports the coefficient β of equation (3) estimated using U.S. monetary policy shocks à la Bauer and Swanson (2023b) as the instrument for the Fed funds rate. Lag-augmented heteroskedasticity-corrected standard errors are in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

D Appendix to Section 5

D.1 Details of the Calibration

We set the common discount factor across all countries to $\beta = 0.9950$, which corresponds to a steady-state real interest rate of 2 percent per year. The intertemporal elasticity of substitution, σ, and the inverse Frisch elasticity of labor supply, χ, take standard values of 1 and 2, respectively. Using the OECD Inter-Country Input-Output Tables, we observe that intermediate inputs account for roughly 50 percent of total production. Accordingly, we calibrate the capital share and the labor share in gross output to $\alpha = 0.10$ and $\gamma_p = 0.40$, respectively. The remaining firm-level parameters follow Justiniano, Primiceri, and Tambalotti (2010). The steady-state price markup is set to $\theta_p = 0.20$, while the degree of price rigidity, ξ_p, is set to 0.60. The elasticity of intratemporal substitution between domestic and imported tradable consumption goods is set to $\rho_c = 1.01$. The elasticity of substitution between intermediate inputs in production, ρ_m, and the elasticity of substitution between final tradable and non-tradable goods, ρ_s, are set to 0.60 and 0.44, respectively, implying complementarity between these goods. These values are in line with standard calibrations in open-economy macroeconomic models (see, for example, Erceg, Gust, and López-Salido, 2010; Akinci and Queralto, 2023; Schmitt-Grohe and Uribe, and references therein).

Trade openness—both at the final and intermediate good levels—lies at the core of the asymmetries in our model. The calibration of these parameters is guided by evidence from the OECD Inter-Country Input-Output Tables. Within the tradable consumption bundle (i.e., at the final-goods level), the United States is the least open to trade, while EMs are the most open among the three blocs. Moreover, EMs account for the bulk of global exports in final goods. In contrast, at the production stage (i.e., in intermediate-goods trade), the pattern is reversed: the United States is the most open, and EMs are the least open to trade. This difference reflects EMs' role as major global producers of intermediate inputs.

Another important asymmetry across blocs concerns the share of tradable goods in consumption. As emphasized by Akinci et al. (2024), the tradable component of U.S. consumption has been steadily declining, with non-tradable goods accounting for an increasing share —approximately 31% in 2019. We calibrate the relative weight of tradable to non-tradable consumption, denoted by ω_s, to reflect these differences across blocs. Specifically, we assign $\omega_s = 0.30$ for the United States, $\omega_s^* = 0.75$ for AEs, and $\omega_s^{**} = 0.30$ for EMEs, capturing the relatively low tradable share in U.S. consumption and the higher reliance on tradable goods in developing economies. Finally, the relative sizes of the three blocs are set to $n_H = 0.2$, $n_F = 0.5$, and $n_W = 0.3$. Table D1 reports the complete set of calibrated parameters.

Table D1. Parameter Descriptions and Values

Parameter	Description	Value
β	Discount factor	0.9950
σ	Elasticity of substitution	2.00
χ	Inverse Frisch elasticity of labor supply	2.00
α	Capital share of gross output	0.10
γ_p	Labor share of gross output	0.40
θ_p	Price markup	0.20
ξ_p	Price stickiness	0.60
ρ_c	Final tradeable consumption price elasticity	1.01
ρ_m	Intermediate tradeable production price elasticity	0.60
ρ_s	Sectoral price elasticity	0.44
ω_s	Sectoral bias for final tr consumption relative to US ntr good	0.30
ω_s^*	Sectoral bias for final tr consumption relative to EM ntr good	0.30
ω_s^{**}	Sectoral bias for final tr consumption relative to AE ntr good	0.75
θ_w^*	EM working capital ratio	2.50
ν_H^*	US traded bond adjustment cost for EM	0.001
ν_H^{**}	US traded bond adjustment cost for AE	0.001
n_H	US size relative to world	0.20
n_F	EM size relative to world	0.50
n_W	AE size relative to world	0.30
ρ_{mp}	Persistence of monetary shock	0.30

D.2 Firm's Profit Maximization Under Working Capital

In this section, we show how U.S. monetary policy can directly effect EM firms marginal costs and prices in the EM. Solving the foreign firm's profit maximization problem the nominal marginal cost in sector $i = \{T, N\}$ is:

$$MC_t^{i*} = \left(\frac{W_t^*/Z_t^{i*}[1+\mu_t]}{\gamma_p}\right)^{\gamma_p} \left(\frac{{R_{K,t}^i}^*}{\alpha}\right)^{\alpha} \left(\frac{{P_{M,t}^i}^*[1+\mu_t]}{1-\alpha-\gamma_p}\right)^{1-\alpha-\gamma_p} \tag{D7}$$

where $\mu_t = \left[\kappa^* R_{risk,t}^n + (1-\kappa^*)R^{n*}_t - 1\right]\theta_w^*$ represents the interest rate payments associated with the given input being financed through working capital. $R_{risk,t}^n$ is defined as

$$R_{risk,t}^n = R_t^n + Q_t, \text{ where}$$

$$Q_t \equiv \left(\frac{S^H_{F,t}}{S^H_{F,t-1}} - 1 \right)$$

meaning a risk adjusted interest rate on dollar debt, increasing proportionally to exchange rate depreciation of the EM against the U.S. dollar. Taking logs,

$$\log MC^{i*}_t = \gamma_p \left[\log W^*_t - \log Z^{i*}_t + \log(1+\mu_t) - \log \gamma_p \right] \tag{D8}$$

$$+ \alpha \left[\log {R^i_{K,t}}^* - \log \alpha \right] \tag{D9}$$

$$+ (1-\alpha-\gamma_p) \left[\log {P^i_{M,t}}^* + \log(1+\mu_t) - \log(1-\alpha-\gamma_p) \right] \tag{D10}$$

We wish to understand how an increase in the U.S. monetary policy rate (R^n_t) changes the real marginal cost of firms. In order to emphasize the relevant channels, we ignore the *indirect/general equilibrium effect* of R^n_t on wages, hours worked, and rental rate of capital.

Collecting the terms that depend on R^n_t,

$$\log MC^{i*}_t = (\text{other terms} + (1-\alpha)\log(1+\mu_t) + (1-\alpha-\gamma_p)\log P^{i*}_{M,t}. \tag{D11}$$

Define elasticities

$$\varepsilon_M \equiv \frac{\partial \log P^{i*}_{M,t}}{\partial \log R^n_t}, \qquad \varepsilon_Q \equiv \frac{\partial \log Q_t}{\partial \log R^n_t}, \qquad \varepsilon_* \equiv \frac{\partial \log R^{n,*}_t}{\partial \log R^n_t}. \tag{D12}$$

Thus

$$\frac{\partial Q_t}{\partial R^n_t} = \frac{\varepsilon_Q Q_t}{R^n_t}, \qquad \frac{\partial R^{n,*}_t}{\partial R^n_t} = \frac{\varepsilon_* R^{n,*}_t}{R^n_t}.$$

Differentiate μ_t w.r.t. R^n_t:

$$\frac{\partial \mu_t}{\partial R^n_t} = \theta^*_w \frac{\partial}{\partial R^n_t} \Big(\kappa^* (R^n_t + Q_t) + (1-\kappa^*) R^{n,*}_t - 1 \Big) \tag{D13}$$

$$= \theta^*_w \Big(\kappa^* \Big(1 + \frac{\partial Q_t}{\partial R^n_t} \Big) + (1-\kappa^*) \frac{\partial R^{n,*}_t}{\partial R^n_t} \Big) \tag{D14}$$

$$= \theta^*_w \Big(\kappa^* \Big(1 + \frac{\varepsilon_Q Q_t}{R^n_t} \Big) + (1-\kappa^*) \frac{\varepsilon_* R^{n,*}_t}{R^n_t} \Big). \tag{D15}$$

Hence

$$R^n_t \frac{\partial \mu_t}{\partial R^n_t} = \theta^*_w \Big(\kappa^* \big(R^n_t + \varepsilon_Q Q_t \big) + (1-\kappa^*) \varepsilon_* R^{n,*}_t \Big), \tag{D16}$$

and

$$\frac{\partial \log(1+\mu_t)}{\partial \log R^n_t} = \frac{R^n_t}{1+\mu_t} \frac{\partial \mu_t}{\partial R^n_t} = \frac{\theta^*_w}{1+\mu_t} \Big(\kappa^* \big(R^n_t + \varepsilon_Q Q_t \big) + (1-\kappa^*) \varepsilon_* R^{n,*}_t \Big). \tag{D17}$$

Therefore the elasticity of nominal marginal cost w.r.t. the U.S. policy rate is

$$\frac{\partial \log MC_t^{i*}}{\partial \log R_t^n} = \underbrace{(1-\alpha)\frac{\theta_w^*}{1+\mu_t}\Big(\kappa^*\big(R_t^n + \varepsilon_Q Q_t\big) + (1-\kappa^*)\varepsilon_* R_t^{n,*}\Big)}_{\text{Debt channel}} + \underbrace{(1-\alpha-\gamma_p)\,\varepsilon_M}_{\text{Imports Channel}} \tag{D18}$$

Due to import price stickiness and the fact that EM firms have sizable intermediate imports from the rest of the world in foreign currency, exchange rate depreciation caused by a U.S. tightening will directly raise intermediate import prices. In other words, $\epsilon_M > 0$. Similarly, $\epsilon_Q > 0$, due to currency depreciation. At this stage, the response of domestic interest rates remain unclear but we show both in the data and in our calibrated model that $\epsilon_* > 0$.

Are there any re-balancing channels? Yes. So far we have ignored the responses of wages and rental rate of capital. Since $R^n_{risk,t}$, ${P^{i*}}_{M,t}$ go up, wages will fall. The response of the rental rate of capital is ambiguous. We argue that these channels are second order and do not offset the direct effect of interest rate changes in the U.S. will act as a cost push shock. We show this in the main body of the paper for a reasonable calibration of the model.

D.3 Understanding Emerging Economy Taylor Rules

In our baseline model, EM central banks are more hawkish relative to AE central banks yet they observe more inflation relative to AEs, consistent with the findings of Nakamura et al. (2025) for the post-Covid inflationary shocks. We use our model to explore why EME central banks are more hawkish relative to AE central banks by setting the Taylor rule parameters of EMEs to that of the AEs (see Table 6). Specifically, we let $\gamma_r^* = \gamma_r$, $\gamma_y^* = \gamma_y$, and $\gamma_\pi^* = \gamma_\pi$.

Figure D21 presents select EME responses to a 100 basis points increase in the Fed funds rate. When we let the EME central bank use the AE Taylor rule coefficients instead of the baseline hawkish rule, EME inflation jumps by roughly twice as much on impact. When EME central bankers are less hawkish their economy is more vulnerable to larger jumps in inflation. Therefore EME central banks under the dovish rule, despite putting less effective weight on inflation (due to higher interest-rate smoothing), need to raise their policy rate by more relative to the hawkish rule. To offset this large increase in inflation, they need to raise the policy rate even more than the amount they would have under a more hawkish Taylor rule. This mechanism rationalizes why estimated EME Taylor coefficients are more hawkish in the data and helps explain why EME policy rates comove more strongly with the Fed than those of AEs.

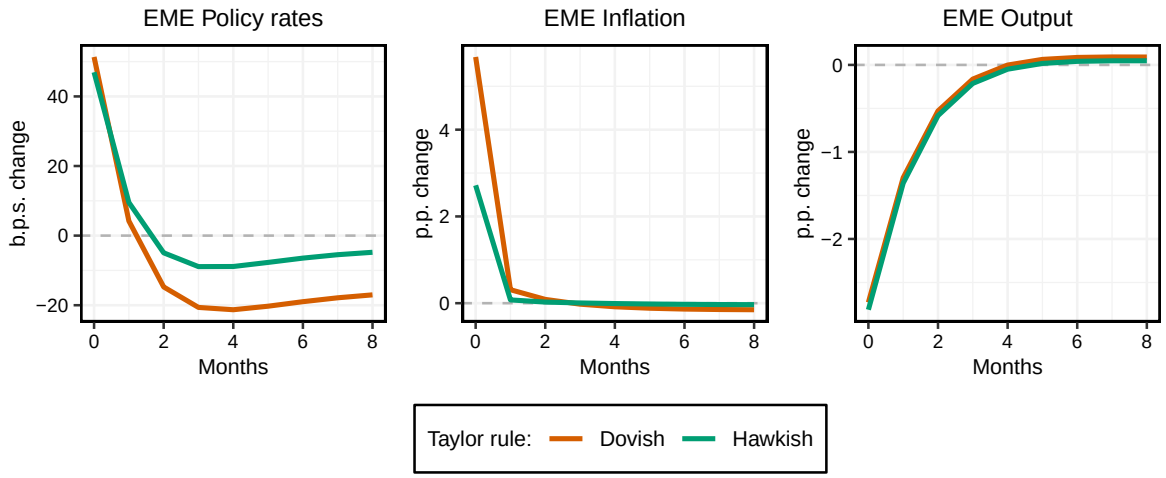


FIGURE D21. EME POLICY RATE RESPONSE, HAWK VS DOVE

Notes: The figure shows the effects of a 100 basis points increase in the Fed funds rate in our model with different Taylor rule parametrization.