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Relative monetary policy and exchange rates

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Non-technical summary

Research Question

Understanding the fundamental drivers of exchange rates remains one of the key challenges in international economics and finance. As the exchange rate is the relative price between two currencies, the relative stance of monetary policy between two currency areas should be an important determinant of its value. Yet, a large macroeconometric literature generally does not view shocks to monetary policy as particularly important for explaining exchange rate behaviour. Although the literature has made significant progress in the precise identification of monetary policy shocks, past work has arguably been fairly narrowly focused – for instance by considering only conventional policy by only one central bank –, preventing to get a full picture of just how important monetary policy is as a driver of the exchange rate.

Contribution

This paper studies the importance of monetary policy for the exchange rate comprehensively. Different from existing work, it takes into account both involved central banks simultaneously, as well as the multifaceted nature of monetary policy, both when it comes to its implementation and channels of transmission to the exchange rate. This is achieved by adapting instrumental variable techniques for shock identification in a structural vector autoregression framework: rather than narrowly focusing on high-frequency changes in interest rates in the immediate aftermath of monetary policy announcements, the paper uses high-frequency changes in the exchange rate itself.

Results

Shocks to the relative stance of monetary policy account the majority of fluctuations in the exchange rate. Specifically, roughly 75 percent of the variation of the EUR-USD exchange rate over a one-month horizon can be explained by shocks to the Fed's monetary policy relative to the Eurosystem's. Similarly high values are obtained also for other exchange rates involving the Japanese yen and British pound. Identifying US and euro area shocks separately reveals that both are roughly equally important for the EUR-USD rate. When it comes to the propagation of the shocks, a relative monetary policy tightening lets the domestic currency appreciate on impact and then gradually depreciate again – i.e., there is no evidence of *delayed overshooting*, just as predicted by standard theory. Yet, sizable jumps in exchange rates are associated with only small changes in interest rate differentials. This points to the importance of transmission channels over and above those operating through risk-free rates.

Nichttechnische Zusammenfassung

Fragestellung

Die Identifizierung der grundlegenden Bestimmungsfaktoren von Wechselkursen bleibt eine der zentralen Herausforderungen in den internationalen Finanz- und Wirtschaftswissenschaften. Da der Wechselkurs der relative Preis zwischen zwei Währungen ist, sollte die relative Ausrichtung der Geldpolitik zweier Währungsräume ein wichtiger Bestimmungsfaktor seines Wertes sein. Dennoch sieht eine umfangreiche empirische Literatur geldpolitische Schocks im Allgemeinen nicht als besonders wichtig für die Erklärung des Wechselkursverhaltens an. Bisherige empirische Studien waren jedoch in Bezug auf die Identifikation geldpolitischer Schocks und ihrer Übertragungskanäle auf den Wechselkurs vergleichsweise eng gefasst, konzentrieren sich also beispielsweise auf konventionelle geldpolitische Maßnahmen nur einer einzelnen Zentralbank.

Beitrag

Dieses Papier untersucht die Bedeutung der Geldpolitik für den Wechselkurs umfassend. Im Gegensatz zu bestehenden Arbeiten berücksichtigt es gleichzeitig beide beteiligten Zentralbanken sowie die vielschichtige Natur der Geldpolitik, sowohl in Bezug auf ihre Implementierung als auch auf die Übertragungskanäle auf den Wechselkurs. Dies wird erreicht, indem Instrumentalvariablen zur Schockidentifikation in einem strukturellen vektor-autoregressiven Modell zur Anwendung kommen. Anstatt sich dafür jedoch eng auf hochfrequente Änderungen von Zinssätzen nach geldpolitischen Ankündigungen zu konzentrieren, verwendet das Papier hochfrequente Änderungen des Wechselkurses selbst.

Ergebnisse

Schocks in der relativen Ausrichtung der Geldpolitik erklären den Großteil der Schwankungen des Wechselkurses. So können etwa 75 Prozent der Variation des EUR-USD-Kurses über einen Ein-Monats-Horizont durch Schocks in der Geldpolitik der Fed im Vergleich zum Eurosystem erklärt werden. Ähnlich hohe Werte werden auch für andere Wechselkurse des japanischen Yen und des britischen Pfund gemessen. Werden Schocks in den USA und im Euroraum separat identifiziert, sind beide ungefähr gleich wichtig für den EUR-USD-Kurs. Was die dynamische Ausbreitung der Schocks betrifft, führt eine relative geldpolitische Straffung dazu, dass die heimische Währung sofort auf- und dann allmählich wieder abwertet. Es gibt also keine Anzeichen für ein *verzögertes Überschießen*, genau wie es die Standardtheorie vorhersagt. Andererseits sind beträchtliche Sprünge in den Wechselkursen mit nur kleinen Änderungen in Zinsdifferenzen verbunden. Dies weist auf die Bedeutung von Übertragungskanälen hin, die über diejenigen hinausgehen, die durch risikofreie Zinssätze wirken.

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Relative Monetary Policy and Exchange Rates

Sören Karau[1]

Abstract

I show that the majority of short-term nominal exchange rate fluctuations among large economies can be explained by changes in the relative stance of their monetary policies. Adapting recently developed instrumental variable techniques for shock identification, I find that monetary policy shocks of the US relative to the euro area account for 76 percent of the short-term fluctuations of the USD-EUR exchange rate over a one-month horizon – substantially more than previously documented. Similar results are obtained for exchange rates involving the British pound and Japanese yen. Relative monetary policy shocks explain a larger fraction of variability of the exchange rate than of interest rate differentials throughout the yield curve, and small changes in risk-free rates are associated with sizable jumps in the exchange rate. Identifying US and euro area shocks separately reveals that both are important for the USD-EUR rate. Taken together, these findings speak to the significance of (not only US) monetary policy in driving frictions in interest parity relations that have recently been found to be crucial for understanding exchange rate behavior from a theoretical perspective.

Keywords: Monetary Policy, Exchange Rates, Proxy VAR

JEL Codes: E44, E52, F31, F41

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1 Introduction

Understanding the fundamental drivers of exchange rates remains one of the key challenges in international economics and finance. There is a long-standing literature documenting the disconnect of exchange rates from fluctuations in macroeconomic variables (Frankel and Rose, 1995; Obstfeld and Rogoff, 2001; Engel, 2014), seemingly indicating that the two must be driven by different sets of shocks. Indeed, addressing various exchange-rate-related puzzles, recent theoretical considerations suggest that exchange rates are primarily driven by financial or currency demand shocks, which shift the demand of investors for home- versus foreign-currency assets (Itskhoki and Mukhin, 2021; Eichenbaum et al., 2021). At the same time, practitioners and the financial press regularly invoke the actions of central banks when discussing key drivers of foreign exchange market activity. In the words of one market participant, "fundamentally, exchange rates are an extension of monetary policy".[1]

This paper revisits the question of how important monetary policy is in driving exchange rate fluctuations. Whereas a large macroeconometric literature confirms that shocks to monetary policy do significantly affect exchange rates (Eichenbaum and Evans, 1995; Bjørnland, 2009; Rogers et al., 2018; Rüth, 2020), they are generally not viewed as the main driver, but only as one among many (Faust and Rogers, 2003). In contrast, this paper provides empirical evidence that the majority of fluctuations in exchange rates can be attributed to changes in the relative stance of monetary policy. I reach this conclusion by adapting recently developed macroeconometric approaches to ensure that the identification of monetary policy shocks is precise, yet comprehensive in scope.

First, different from much of the earlier VAR literature, the paper identifies shocks to monetary policy using instrumental variable techniques. These avoid placing any potentially contestable recursiveness assumptions or a multitude of sign and zero restrictions on the impact effects of shocks.[2] Instead, identification is guided by the high-frequency response of financial market prices around monetary policy announcements, an approach that has gained currency in the estimation of monetary policy effects in macroeconomic VAR models (Gertler and Karadi, 2015; Miranda-Agrippino and Ricco, 2021), also when it comes to exchange rates (Rogers et al., 2018; Rüth, 2020). Here, I employ the same approach not only using macro data but utilize the fact it works well also in weekly and daily models of financial market data, adding to the robustness of the results and allowing for much tighter confidence bands. Importantly, I make sure to control for information effects

[1] See Financial Times (2023).

[2] Sign restriction identification schemes in principle allow to place identifying restrictions very agnostically. However, in practice researchers often need to place many restrictions on the responses of many model variables in order to achieve meaningful inference, increasing the computational burden and the dependence of the results on the exact model specification. Further, Wolf (2020) shows that sign restriction procedures are generally susceptible to a *masquerading* problem, such that the identified shocks of interest can be confounded by combinations of other shocks.

that have been shown to potentially alter responses significantly (Jarociński and Karadi, 2020; Miranda-Agrippino and Ricco, 2021), and again also play a role in the context of exchange rates (Franz, 2020; Gürkaynak et al., 2021; Antolin-Diaz et al., 2023).

Second, the analysis is comprehensive in that it takes into account changes in the monetary policy of both involved central banks simultaneously. The existing literature usually identifies shocks stemming from only one central bank (most often the Fed), and if multiple shocks are identified they are estimated in separate models – preventing to get a full picture of just how much variation in exchange rates monetary policy can explain. In contrast, the analysis in this paper encompasses changes to the relative stance of monetary policy in a single model. This is achieved in two alternative ways. The first approach estimates a single relative monetary policy shock. It hence treats a relative monetary tightening by, say, the Eurosystem equivalently to a relative easing by the Fed, and vice versa. This approach has the advantage that it features especially high instrument relevance and is computationally more convenient. The second approach identifies US and euro area monetary policy shocks separately but simultaneously. Placing restrictions on the correlation structure of the two instruments and shocks then achieves set instead of point identification. Yet, this approach has the advantage that it allows to assess the relative importance of the involved central banks, in this case the Fed and the Eurosystem in affecting the USD-EUR exchange rate. Notably, both approaches lead to very similar results in that they confirm the large (combined) importance of monetary policy for exchange rate behavior.

Third, the analysis is comprehensive in that it accounts for the potentially multi-faceted nature of monetary policy – both in terms of its implementation and its effects on exchange rates via many possible channels of transmission. Following the Great Financial Crisis (GFC), central banks around the globe have engaged in unconventional monetary measures including forward guidance and large scale asset purchases, which are not captured by focusing on short-term interest rate changes alone. Past studies have established that also unconventional policy affects exchange rates (Dedola et al., 2021), often employing high-frequency interest rate responses to monetary policy announcements throughout the yield curve (Rogers et al., 2018; Inoue and Rossi, 2019). Yet, the existing literature usually studies these different aspects of monetary policy one at a time and/or not with the aim of establishing their overall importance for exchange rate behavior. Further, it has been argued that monetary policy has important effects on financial markets that are not captured by focusing on changes in interest rates (alone), no matter their maturity (Boehm and Kroner, 2023). This can either relate to monetary policy affecting risk-taking behavior (Kroencke et al., 2021; Bauer et al. 2023) or liquidity premia and convenience yields (Jiang et al., 2023).

These considerations inform the choice of the instrument used for identification in this paper. Specifically, in the baseline specification for the USD-EUR rate, I construct the instrument from high-frequency changes of the *USD-EUR exchange rate itself* around monetary policy announcements by the Federal Open Market Committee (FOMC) and the ECB Governing Council. Equivalently, the instrument is constructed from exchange rate movements of the pound and yen when extending the analysis to monetary policies by the Bank of England's Monetary Policy Committee (MPC) and the Bank of Japan's Policy Council. Such an instrument naturally lends itself to the identification of a relative monetary policy shock but can also be used for identifying monetary shocks of the two involved central banks separately, as I outline below. The approach ensures that the instrument is both comprehensive yet agnostic with respect to potential transmission channels of monetary policy to exchange rates. Validating the approach, I confirm that the identified shocks produce conventional macro responses. Notwithstanding these considerations, in multiple robustness exercises I also identify monetary policy shocks using the more traditional approach of constructing a broad interest rate-based instrument, with overall similar results.

I obtain four main findings. First, a contractionary (relative) shock to monetary policy leads to an immediate appreciation and a subsequent depreciation of the domestic currency. There is essentially no sign of a delay in overshooting – qualitatively in line with Dornbusch-type predictions, and different from much of the literature on the dynamic impact of monetary policy on exchanges rates.

The second and most central finding relates to the overall importance of monetary policy: shocks to its relative stance account for the majority of exchange rate fluctuations. In the benchmark specification, the share of the explained forecast variance during the first month for the USD-EUR exchange rate is 76 percent. This share is significantly larger than suggested by previous estimates. I obtain similarly large figures when varying the model along many dimensions, with estimates varying between 65 when using noisy daily data to 86 percent when not cleansing the instrument from potential information effects. Shares are also similarly high for exchange rates involving the pound and yen.

The third main finding pertains to the relative importance of the two involved central banks: when separating US from euro area monetary policy, both turn out to be of roughly equal importance for explaining exchange rate movements in the main specification. While US monetary policy can be shown to be a more important driver of international financial conditions more broadly, this out-sized importance does not seem to equally apply to the USD-EUR rate. This finding is corroborated further by the fact that, again, I find those exchange rates not involving the US dollar also to be driven to a large extend by shocks to the relative stance of the respective central banks.

Finally, the fourth finding concerns the relative impact on exchange rates compared to interest rates: the importance of the relative monetary policy shocks for the exchange rate is generally significantly larger than that for yield differentials between the two respective currency areas. Further, relatively small changes in interest rate differentials are associated with sizable jumps in, and subsequent reversals of, exchange rates. Although the reversal is qualitatively in line with textbook uncovered interest parity (UIP), quantitatively I find that investors bear negative excess currency returns when investing in the higher yielding currency, conditional on the shock: following a one-percent appreciation of the exchange rate, the subsequent depreciation results in pecuniary losses of roughly 50bp over a one-year horizon. When investigating this failure of frictionless UIP, I find evidence for the importance of non-pecuniary returns: a rise in convenience yields, as measured by deviations of covered interested parity between government bonds (Du et al., 2018a), compensates investors for part of the negative currency returns. Notably, also here the picture is fairly symmetric: not only does tighter US monetary policy induce a rise in the convenience yield of domestic safe assets – the usual focus in the literature –, but so does euro area monetary policy.

These results have important implications for the ongoing discussion around the drivers of exchange rate fluctuations. Most broadly, the paper adds to the recently renewed optimism regarding the link between economic news and exchange rates (Sarno and Schmeling, 2014; Lilley et al., 2022; Stavrakeva and Tang, 2023; Chahrour et al., 2021; Engel and Wu, 2022). Further, the paper's findings lend support to theories that augment standard UIP relationships with additional terms related to currency risk premia and convenience yields (e.g. Engel, 2016; Engel and Wu, 2022; Jiang et al., 2021; Jiang et al., 2024b). Indeed, the aforementioned work by Itskhoki and Mukhin (2021) shows that only financial or currency demand shocks that introduce wedges in frictionless UIP relationships can account for a number of long-standing puzzles in exchange rate behavior. In their model, shocks to the willingness of risk-averse financial intermediaries to take on exchange rate risk are the main drivers of exchange rate fluctuations. A similar type of shock is present in Eichenbaum et al. (2021) who interpret it as flight to safety or liquidity concerns by international investors. The main finding of this paper – that monetary policy shocks can account for the majority of short-term exchange rate variability – then informs the search for a source of these shocks. Indeed, my findings echo the concluding thoughts in Itskhoki and Mukhin (2021, p.2225) that a "microfoundation of financial shocks is essential, as they may endogenously interact with or arise from monetary policy". Such an interpretation is in line with a growing literature that finds (particularly US) monetary policy to affect risk sentiment both domestically (Bekaert et al., 2013; Bauer et al., 2023) and internationally (Bruno and Shin, 2015; Kalemli-Özcan, 2019; Miranda-Agrippino and Nenova, 2022), and with efforts to create models in which monetary policy affects UIP premia endogenously (Jiang et al., 2023; Akinci and Queralto, 2024).

Related Literature. The paper contributes to various strands of the literature. First and foremost, there is a large empirical literature studying the dynamic effects of monetary policy on exchange rates in a structural VAR framework. Most studies confirm that a monetary tightening leads to an appreciation of the domestic exchange rate, in line with standard theory. Yet, there is a long-running debate over the question of whether there is evidence of *delayed overshooting* in the sense that the peak exchange rate response is reached not on impact but only substantially later in time. While the seminal contribution by Eichenbaum and Evans (1995) and several other papers (Faust et al., 2003; Scholl and Uhlig, 2008; Bouakez and Normandin, 2010; Heinlein and Krolzig, 2012; Linnemann and Schabert, 2015) find evidence often interpreted as delayed overshooting, some more recent contributions find less support for such a delay (Bjørnland, 2009; Bjørnland and Halvorsen, 2014), in particular those that use more recently developed techniques for shock identification (Rogers et al., 2018; Rüth, 2020; Liao et al., 2023). Kim et al. (2017) find delayed overshooting for US policy shocks only in the Volcker era, but not afterwards.[3] While not the main focus of this paper, I find no evidence of delayed overshooting and show that the peak exchange rate response is generally reached immediately, even when using weekly data.

Fewer papers explicitly study how important monetary policy shocks are overall for exchange rate fluctuations. In those that include such an analysis, usually by means of a forecast error variance decomposition (FEVD), the exercise is usually conducted only for Fed policy shocks – omitting the role of the central bank issuing the other currency.[4] And even then estimates vary widely. Earlier studies that find evidence of delayed overshooting tend to find higher explanatory power for US shocks only over medium- to long-run horizons.[5] Conversely, studies using more recently developed identification schemes without evidence of delayed overshooting find an economically small role of US monetary policy.[6] Innovating on shock identification and taking into account both involved central banks, I show that the importance of monetary policy since the early 2000s is highest over short-term horizons and substantially larger than previously documented.

[3] Other VAR studies on monetary policy and exchange rates include Kim and Roubini (2000), Faust et al. (2003), Jääskelä and Jennings (2011), Kim and Lim (2018), Rüth and Van der Veken (2023).

[4] To my knowledge, Dedola et al. (2021) is the only other paper that identifies a relative monetary policy shock of the Fed and the Eurosystem, but the analysis is limited to asset purchase shocks and the authors do not assess their overall importance for USD-EUR variation. Rogers et al. (2018) also study US and euro area monetary policy shocks on the USD-EUR exchange rate, but identify the shocks in separate models and without conducting variance decompositions. Voss and Willard (2009) identify two monetary policy shocks simultaneously but at the cost of imposing implausible timing restrictions using quarterly data.

[5] For instance, Eichenbaum and Evans (1995) report FEVD shares between 20 and 40 percent after three years, Bouakez and Normandin (2010) of below ten to up to 60 percent after one year, depending on the currencies involved. Groshenny and Javed (2023) report values between 3 to 25 percent.

[6] For instance, short-term estimates by Rüth (2020) are between five and 14 percent for the nominal effective USD exchange rate. Estimates in Faust et al. (2003), who also make use of high-frequency asset price information for identification, are only somewhat larger, with wide confidence bands.

2 Monetary Policy Announcements and Exchange Rate Volatility

Isolating the causal impact of monetary policy on asset prices is non-trivial. Financial markets respond to all sorts of developments and news, and monetary policy might not only affect but also endogenously respond to asset price developments. One part of the literature has therefore opted for an event-study approach by measuring changes of asset prices in short time windows around central bank announcements (Gürkaynak et al., 2005). If these windows are short enough, the causal impact of policy can be measured fairly accurately as monetary policy makers decide on policy some time before the announcement (ruling out reverse causality) and no other major news hit markets at the same time (ruling out other confounding factors).

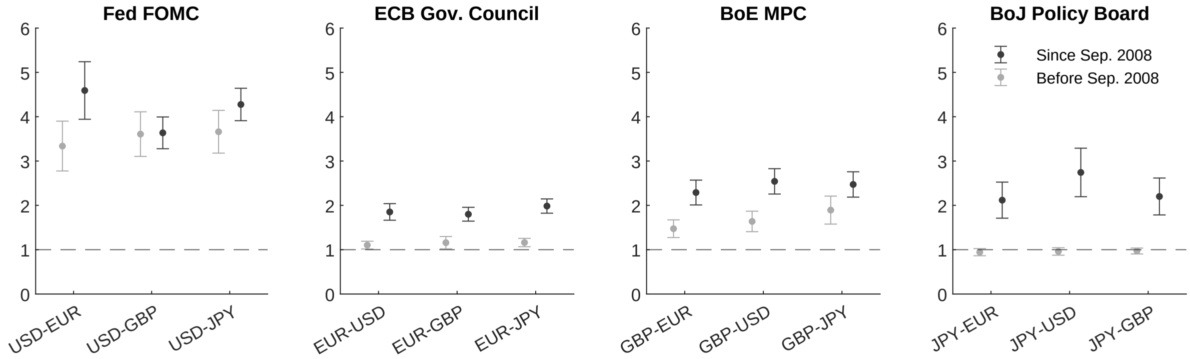
In order to get a first sense of how important monetary policy is for exchange rates, I therefore conduct a high-frequency analysis of return volatilities. To that end I use minute-by-minute data on six exchange rates (USD-EUR, USD-GBP, USD-JPY, EUR-GBP, EUR-JPY, GBP-JPY) obtained from `tickstory.com` from 2003 to 2023. I then compute returns during short monetary policy announcement windows for the respective central banks (FOMC, ECB Governing Council, Bank of England Monetary Policy Committee (MPC) and Bank of Japan Policy Board). Following the literature, the windows start 10 to 15 minutes prior to the press statements and end 15 to 20 minutes afterwards, except when there was a subsequent press conference, in which case the window ends 60 to 75 minutes after the beginning of the press conference.[7] Appendix A provides more details.

In order to compare the sensitivity of returns, I compute volatilities in equivalent time windows for each day without a monetary policy announcement. If monetary policy was an important factor, one would expect exchange rates to be on average more volatile around monetary policy announcements than normally. Figure 1 shows whether this is the case. It depicts mean standard deviations of minute-by-minute returns of the six exchange rates during the announcements windows of the four central banks, expressed relative to their mean values in the absence of any announcements.

A few observations stand out that anticipate some of the results in the VAR analysis below. First, exchange rate volatility is indeed generally significantly higher around monetary policy announcements, with most depicted values statistically significantly above unity. Second, the sensitivity has increased in the period after the financial crisis (dark

[7] The FOMC began to hold press conferences after every second press statement in 2011, and after each one only in 2019. To account for that, the window ends 20 minutes after the beginning of the press statement in case there was no subsequent press conference. Similar considerations apply for the MPC and BoJ Policy Board.

Figure 1: Relative Increase of Exchange Rate Return Volatilities Around Monetary Policy Announcements



Note. Mean standard deviation of exchange rate returns in monetary policy announcement window relative to equivalent time window on days without respective monetary policy announcements. For instance, a value of 2 in the left panel means that the depicted exchange rates were twice as volatile around an FOMC announcement compared to a non-announcement afternoon. Based on minute-by-minute data. Time sample: May 2003 to Aug. 2008 (light, left) and Sep. 2008 to May 2023 (dark, right). Size of bars indicates 95% confidence intervals.

right) compared to before (light left bars). This finding is in line with the results in Ferrari et al. (2021), who link the heightened responsiveness to the low interest rate environment and the introduction of new monetary policy tools following the financial turmoil.[8]

Third, and most notably, the sensitivity to monetary policy news varies with their source: the increase in volatility is much larger for Fed announcements than for those of the other three central banks. Indeed, for the period before the financial crisis, monetary policy annoucements by the Bank of Japan's Policy Council were not generally associated with increased yen volatility at all. The ECB Governing Council and the Bank of England's MPC also before the crisis did affect euro and pound exchange rate volatility significantly, but quantitatively to a more limited extent. Also after the crisis, their impact is significantly smaller than the FOMC's.

While these results hence point to monetary policy as an important potential driver of exchange rate fluctuations, the analysis of high-frequency responses is by construction limited to short time windows. On its own, it can hence not address the question of how important monetary policy is for exchange rate behavior generally. Therefore, the following section describes an econometric framework that employs the high-frequency data used here in order to estimate the overall importance of monetary policy for exchange rate fluctuations also over longer time horizons in a comprehensive fashion.

[8] Indeed, the split date of September 2008 is in reference to Ferrari et al. (2021) who observe a break in the relationship in autumn 2008 for US monetary policy. Splitting the sample according to their identified break date for euro area policy, around 2010-11, yields similar results.

3 The Empirical VAR Approach

This section describes the main empirical methodology based on a structural VAR framework. As will be outlined, relative to the existing literature, particular emphasis is placed on the open-economy context in order to simultaneously account for US and euro area monetary policy, which is crucial when studying effects on the exchange rate. Section 4 will then apply the methodology to the study of the USD-EUR exchange rate, Section 5 will extend it to exchange rates involving the pound and yen.

The empirical analysis is based on the proxy VAR approach using an external instrument for shock identification (Mertens and Ravn, 2013; Stock and Watson, 2018) that has become widely used to estimate the effects of monetary policy (Gertler and Karadi, 2015; Caldara and Herbst, 2019; Miranda-Agrippino and Ricco, 2021). In the following, I briefly lay out the methodology and subsequently provide more details on the construction of the instrument(s) for the application at hand, which is less standard.

Proxy VARs. The structural model is represented by

$$\mathbf{A}_0\mathbf{y}_t = \mathbf{k} + \mathbf{A}_1\mathbf{y}_{t-1} + \ldots + \mathbf{A}_p\mathbf{y}_{t-p} + \epsilon_t, \quad \boldsymbol{\epsilon}_t \sim \mathcal{N}(\mathbf{0}, \mathbf{I}), \tag{1}$$

where $\mathbf{y}_t$ is a $(n \times 1)$ vector of endogenous variables, and $\mathbf{k}$ is a vector of constants. The corresponding reduced-form VAR is:

$$\mathbf{y}_t = \mathbf{c} + \mathbf{B}_1\mathbf{y}_{t-1} + \ldots + \mathbf{B}_p\mathbf{y}_{t-p} + \mathbf{u}_t, \quad \mathbf{u}_t \sim \mathcal{N}(\mathbf{0}, \boldsymbol{\Sigma}), \tag{2}$$

with $\mathbf{c} = \mathbf{A}_0^{-1}\mathbf{k}$ and $\mathbf{B}_i = \mathbf{A}_0^{-1}\mathbf{A}_i$ and $\mathbf{u}_t = \mathbf{A}_0^{-1}\boldsymbol{\epsilon}_t$.[9]

One can partition the shock vectors into those of monetary policy, $\boldsymbol{\epsilon}_t^{p\prime}$, and other shocks, $\boldsymbol{\epsilon}_t^{q\prime}$, with corresponding residual vectors $\mathbf{u}_t = [\mathbf{u}_t^{p\prime}, \mathbf{u}_t^{q\prime}]'$.

Following the recent literature, I use high-frequency market responses (described below) as an external instrument in the proxy VAR to identify the structural innovations $\boldsymbol{\epsilon}_t^{p\prime}$. For these instruments to be valid, the surprise series $\mathbf{z}_t$ needs to be *relevant* and *exogenous* as follows:

$$\mathbb{E}(\mathbf{z}_t\boldsymbol{\epsilon}_t^{p\prime}) = \boldsymbol{\Phi} \neq \mathbf{0}, \tag{3}$$

$$\mathbb{E}(\mathbf{z}_t\boldsymbol{\epsilon}_t^{q\prime}) = \mathbf{0}. \tag{4}$$

In the baseline approach, only one (relative) monetary policy shock is identified, such that Φ and z_t are scalars. In this case, the unknown Φ can be eliminated and structural coefficients of interest can be derived as the ratio of the coefficients of a regression of the

[9] As outlined in Stock and Watson (2018), Miranda-Agrippino and Ricco (2023) and Forni et al. (2023), the original proxy VAR formulation mostly used in the literature implicitly assumes shock invertibility. Related considerations are discussed in Section 4.5.

residuals on the instrument. In the alternative approach, two monetary policy shocks are identified jointly. In this case, $\mathbf{\Phi}$ is a (2×2) matrix and we have:

$$\mathbf{\Phi} = \mathbb{E}(\mathbf{z}_t \boldsymbol{\epsilon}_t^{p\prime}) = \begin{pmatrix} \mathbb{E}(z_t^{us}\epsilon_t^{us}) & \mathbb{E}(z_t^{us}\epsilon_t^{ea}) \\ \mathbb{E}(z_t^{ea}\epsilon_t^{us}) & \mathbb{E}(z_t^{ea}\epsilon_t^{ea}) \end{pmatrix} \tag{5}$$

In order to separate the two shocks, I specify that the non-diagnoal entries in $\mathbf{\Phi}$ are small compared to the diagonal ones, such that each of the two instruments primarily informs its respective shock. Details and derivations for both approaches are described in Appendix B. I verify that instrument relevance is high throughout, with first-stage F statistics exceeding 10, and vastly so in the financial market models in higher-than-monthly frequency.

Instrument: high-frequency asset price responses. Central for identification in the proxy VAR approach is the instrument $\mathbf{z_t}$. In line with the considerations outlined in Section 2, most often the literature uses asset price responses in short windows around monetary policy announcements as instruments. Here, I follow the same approach, yet adapt it to the particular context of the study of exchanges rate effects of monetary policy. Two considerations guide this choice.

First, monetary policy has become increasingly multifaceted. While traditionally changes in the monetary policy stance were implemented and communicated mostly by changing very short-term interest rates, following the GFC central banks have increasingly relied on previously unconventional tools to affect monetary and financial conditions after lowering rates to their effective lower bound. Focusing only on high-frequency responses of short-term interest rates might then for instance fail to capture how monetary policy affects exchange rates via forward guidance or direct asset purchases. Indeed, there is evidence that not only changes in short-term rates but also path and asset purchase shocks affect exchange rates (Dedola et al., 2021; Inoue and Rossi, 2019; Miranda-Agrippino and Nenova, 2022). There is hence some ambiguity over which interest rate to focus on, while the idea is to capture monetary policy innovations by the two involved central banks as comprehensively as possible.

A second consideration for instrument choice concerns the multiplicity of potential transmission channels of monetary policy with regards to the exchange rate. Whereas the original focus of the transmission mechanism was on interest-rate and trade channels, more recently particular emphasis is placed also on how monetary policy shapes risk appetite and risk-bearing capacity (Bekaert et al., 2013; Bauer et al., 2023). This risk-taking channel not only operates domestically but also has an important international dimension (Bruno and Shin, 2015; Kalemli-Özcan, 2019). *Risk-off* episodes are often accompanied by an appreciation of the US dollar, whose value therefore potentially responds to monetary policy over and above what changes in interest rate differentials would predict. Indeed, an important strand of the international finance literature emphasizes risk premia and convenience yields as crucial determinants of exchange rates (Jiang et al., 2021). And the burgeoning literature on the global financial cycle has identified in particular US monetary policy as one main driver of the co-movement of global risky asset prices, including exchange rates (Miranda-Agrippino and Rey, 2020; Miranda-Agrippino and Nenova, 2022). Over and above the question of which interest rate maturity to focus on, these considerations therefore echo Kroencke et al. (2021) in their assessment that "[m]onetary policy surprises extracted from changes in risk-free interest rates alone will necessarily lack an important part of the information contained in monetary policy announcements."[10]

One natural way to address both considerations simultaneously is to use high-frequency changes of the exchange rate, rather than of interest rates, as an input in the construction of the instrument. This approach will capture both the multifaceted nature of monetary policy implementation and at the same time account for all potential channels through which monetary policy might affect the exchange rate. A final advantage of using an exchange rate-based instrument concerns data availability. Whereas established databases of high-frequency asset price movements around monetary policy announcements – for instance those of Gürkaynak et al. (2005), Cieslak and Schrimpf (2019) and Altavilla et al. (2019), which I also rely upon, as described below – often include data on interest rates in several maturities, this is true only for domestic rates.[11] This is problematic for the measurement of relative effects, as European yields are known to be strongly affected by US monetary policy decisions, such that changes in relative interest rates around an FOMC announcement will generally be much smaller than measured changes in US interest rates indicate. In contrast, high-frequency responses of the USD-EUR exchange rates are available (or can be more easily added from available raw data) for both FOMC and Governing Council announcements, thereby enabling a more accurate measurement of the change in the relative stance of monetary policy.

For these reasons, using high-frequency responses of the exchange rate will be relied upon as the main approach in the VAR analysis. I consider the more standard procedure of using high-frequency responses in interest rates as an alternative in Section 4.5.

Instrument: relative and separately identified monetary policy shocks. The most direct way to answer the main question of the paper – how exogenous changes

[10] Boehm and Kroner (2023) make a similar point and argue that "unexplained variation in equities and exchange rates reflects a dimension of monetary policy that is not spanned by changes in the yield curve."

[11] For instance, US yield response data is available for FOMC announcements and euro area yield data for ECB Governing Council news, but not vice versa.

in the relative stance of monetary policy influence the exchange rate – is to identify a single relative monetary policy shock in the spirit of Dedola et al. (2021). Such a shock can be identified using a single instrument that uses high-frequency asset price changes around both, say, FOMC and ECB Governing Council announcements jointly. For instance, any appreciation of the euro against the US dollar around an announcement, whether by the Fed or ECB, would then indicate a monetary tightening by the Eurosystem relative to the Fed or, equivalently, a relative easing by the Fed. This approach has the advantage that a particularly large number of monetary surprises (i.e. roughly twice the amount) can be used to construct the instrument, substantially increasing instrument relevance, particularly in a model using lower-frequency data. What is more, no additional assumptions have to be made regarding how to allocate the information contained in several instruments to different shocks. Identifying a relative shock is therefore more agnostic and the model in practice is easier to estimate.

Focusing on a single relative monetary policy therefore comes with advantages, but the procedure is silent on the relative contribution of, in this case, US and euro area policy, and on potential differences in their propagation. Estimating the shocks separately in two different structural models would be straightforward and allow to study the latter, but not the former: one could assess differences in impulse responses, but the explained share of exchange rate variation – the ultimate goal of the analysis – can only be consistently computed in a single model that identifies the shocks jointly. Therefore, as an alternative to the relative instrument, I construct two instrument series and identify US and euro area shocks separately but simultaneously. In this case, all market responses to FOMC announcements are used for the US instrument and all responses to Governing Council announcements are used for the euro area instrument. As mentioned above (and explained in detail in Appendix B), the procedure then achieves set identification. It allocates the information contained in the two instruments to the identified structural shocks by specifying that each shock has to primarily explain the variation of its instrument rather than the other.

Instrument: cleansing from information effects. In the presence of information asymmetries between the central bank and market participants, high-frequency asset price responses around monetary policy announcements could contain "information" or "signaling" effects (Melosi, 2017; Nakamura and Steinsson, 2018; Miranda-Agrippino and Ricco, 2021; Cieslak and Schrimpf, 2019; Jarociński and Karadi, 2020; Kerssenfischer, 2022).[12] If the researcher then simply used the changes in expected interest rates as an external instrument for monetary policy shock identification, the exogeneity assumption (4) is likely to be violated. The researcher would then measure not the impulse response to an actual exogenous monetary policy shock, but instead that to some combination of fundamental shocks the central bank responds to.

Just as in the case of interest rates, high-frequency responses of exchange rates have also been found to be contaminated by potential information effects (Gürkaynak et al., 2021). Hence, in order to use them to identify pure monetary policy shocks, I adapt the approach in Jarociński and Karadi (2020). Specifically, for the models in levels that make use of two instrument series, I employ a sign restriction procedure to separate monetary policy from central bank information surprises, as for instance in Cieslak and Schrimpf (2019), Kerssenfischer (2022) and Karau (2023). As outlined in Appendix A, monetary policy surprises that lead to an appreciation of the domestic currency are required to be accompanied by a fall in stock prices.

For the relative models instead, I adapt the *poor man's* approach in Jarociński and Karadi (2020) by simply excluding those surprises that are associated with an appreciation of the domestic currency and with a contemporaneous rise in domestic stock prices, and vice versa. This approach is less flexible in that it classifies each announcement in a binary fashion rather than letting it contain both monetary policy and information components. Applying the sign restriction procedure in the relative models, however, would require information on how relative stock prices change around monetary announcements, which is not available to me over the whole sample period.[13]

Instrumented variable. Throughout, I report impulse responses that are normalized to change the exchange rate on impact by one percent. This reflects that the exchange rate is the variable that is instrumented, i.e. it is the residuals of the exchange rate equation in the VAR that are used in the first stage of the instrumental variable procedure. It is important to note that this choice achieves high instrument relevance, but otherwise is made purely for expositional convenience. Indeed, by construction of the proxy VAR estimator, the instrumented variable does not matter for the relative shape of the impulse responses. Hence, one could equally use, say, an interest rate differential to obtain the exact same responses that I report below by simple rescaling (as long as the correlation of the respective residual with the instrument is high enough, i.e. the instrument is not

[12] By announcing a change in policy, the central bank might reveal its assessment of the macroeconomic outlook to the private sector. Market responses to policy news would then potentially reflect the market's assessment of this additional information, rather than a reaction to exogenous monetary policy shocks alone. For instance, were the market to find that a rate hike primarily reflects policy makers' assessment that the economy to likely perform more favorably than previously expected, the price of risky assets like stocks need not necessarily fall in response to the rate hike, and could even increase.

[13] I obtain very similar results when also using the sign restriction in the relative model, i.e. by assuming that changes in domestic stock prices are a good measure changes in relative stock prices as well.

too weak).[14] By the same token, forecast error variance decompositions are unaffected by the decision which variable to instrument. Hence, using the exchange rate as the instrumented variable is an econometric choice but does not entail any economic argument. In particular, it does not in any way imply that the exchange rate – rather than a short-term interest rate – serves as a direct policy instrument.

4 Relative Monetary Policy and the USD-EUR Exchange Rate

This section features the main empirical analysis that studies the importance of US and euro area monetary policy shocks for the USD-EUR exchange rate. Section 5 will extend the baseline approach to exchange rates involving the British pound and Japanese yen and hence include the study of (relative) Japanese and British monetary policies. The choice to focus on the USD-EUR rate in this section reflects the status of the Fed and the Eurosystem as arguably the two most important rate setters in the world. However, it is also based on data availability that enables me to vary the model and identification specifications along many dimensions, which is essential to validate the approach and compare it to more traditional ways of monetary policy shock identification.

After outlining the data in Section 4.1, Section 4.2 studies a relative monetary policy shock, before Section 4.3 considers US and euro area monetary policy separately by identifying two shocks simultaneously. Section 4.4 then analyzes conditional UIP and the role of convenience yields. Finally, Section 4.5 conducts a battery of robustness tests and showcases additional results of interest based on both identification approaches. Throughout, results are presented in the form of impulse response functions (IRFs), but the main emphasis lies on the explanatory power of monetary policy for the exchange rate, which is assessed by means of a forecast error variance decomposition (FEVD).

4.1 Instrument and VAR data

As in Section 2, the high-frequency response data for the US are taken from the shock databases by Gürkaynak et al. (2005) and Cieslak and Schrimpf (2019),[15] extended with high-frequency data from `tickstory.com` and Refinitiv (LSEG). For the euro area, I rely on the monetary event study database by Altavilla et al. (2019),[16] and extend it likewise. This results in high-frequency asset price responses to 497 monetary policy

[14] Indeed, in case I identify two shocks, there needs to be second instrumented variable and for that I use an interest rate that delivers high instrument relevance.

[15] Available at `https://www.dropbox.com/s/7i5l0u4369jov85/JIEreplication.zip?dl=0`.

[16] Available at `https://www.ecb.europa.eu/pub/pdf/annex/Dataset_EA-MPD.xlsx`.

Table 1: Data and VAR model specifications (USD-EUR models)

Variable	Source	Financial market model (1)	Financial market model (2)	Macro model (3)
USD-EUR exchange rate	Refinitiv (LSEG)	•	•	•
S&P500	S&P	•	•	•
EuroStoxx50	Reuters	•	•	•
10y gov. US bond yields	Bloomberg	•	•	
10y gov. DE bond yields	Bloomberg	•	•	
2y gov. US bond yields	Bloomberg	•	•	
2y gov. DE bond yields	Bloomberg	•	•	
1y gov. US bond yields	Bloomberg			•
1y gov. DE bond yields	Refinitiv (LSEG)			•
3m gov. US bond yields	Reuters	•		
3m euro area OIS rate	Bloomberg	•		
VIX	Bloomberg		•	
1y treasury premium (US - DE)	Du et al. (2018a)	(•)	(•)	
US GDP	FRED, OC			•
Euro area GDP	Eurostat, OC			•
US CPI	FRED			•
Euro area HICP	Eurostat			•
Figures: IRFs		2, 8, C.1, C.2	4, 8, C.1	6, 7, C.3, C.4
Figures: FEVD		3, 9, C.2	5, 9	6, C.3

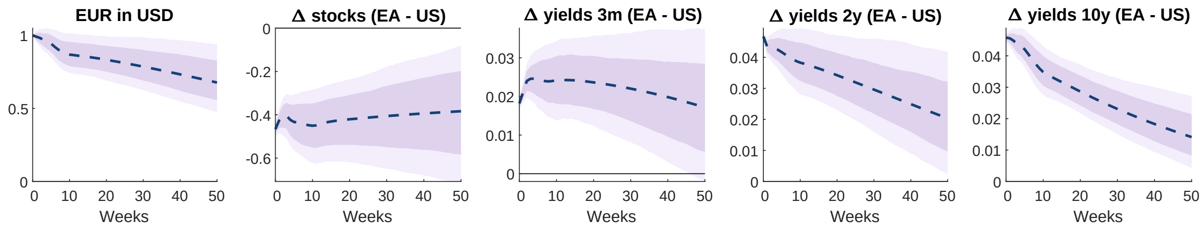
Note. The table lists the variables included in each VAR model, alongside their sources. Columns (1) and (2) refer to financial market models in weekly (and daily) frequency, columns (3) and (4) are estimated in monthly frequency including macro variables. Columns (1) and (3) include all variables (except the exchange rate) in relative terms (euro area minus US), similar to Engel and West (2005), as they identify a single relative monetary policy shock. Column (2) includes all variables in levels (interest rates) or log levels (all other) as it identifies US and euro area monetary policy shocks separately. *OC* denotes *own calculations*, *DE* stands for German. (•) denotes that the respective variable is included only in the models in Section 4.4.

announcements from January 1999 to May 2023. From these, the instrument series are constructed as outlined in Section 3 and described in more detail in Appendix A.[17]

Table 1 provides an overview of the model specifications and time series. The paper puts strong emphasis on the robustness of the results and therefore considers many specifications that vary important aspects of the analysis. The main specification is a VAR model containing financial market variables from both the US and the euro area, estimated in weekly frequency (end-of-week observations). In addition, I estimate monthly models in Section 4.5 including macroeconomic variables from both currency areas. All models include the USD-EUR exchange rate as the main variable of interest, defined as the price of euro in US dollar, such that an increase implies an appreciation of the euro. Additionally, I include data on yields in different maturities and broad stock market indices (S&P500 and EuroStoxx50). For the analysis in Section 4.4 I also add the 1-year treasury premium (deviations from covered interest party of 1-year US treasury relative to German bund yields) obtained from Du et al. (2018a). The monthly macro

[17] In the baseline case, I use time windows that also include central bank press conferences (if there were any), as in Section 2, but verify my main results when using shorter windows that only include press statements.

Figure 2: IRFs to a relative monetary policy shock (weekly data)



Note. Impulse responses to a relative monetary policy shock (relative euro area tightening / US easing), normalized to increase the euro's value vs. the US dollar by 1 percent. Weekly data. Values in percent(age points). Shaded areas denote 68% and 90% confidence bands. Time sample: Jan. 2001 to Mar. 2024.

models contain data on consumer prices (CPI for the US and HICP for the euro area) and monthly GDP estimates obtained from interpolating quarterly series with industrial production indices.

As outlined above, the main specification identifies a single relative monetary policy shock (Section 4.2). In this case, all model variables except the exchange rate (which is already a relative price) need to be specified in relative terms. For instance, interest rates enter in yield differentials, defined as the euro area variable minus its US counterpart. For stock prices, GDP and consumer price indices, I first take first differences of each logged variable, calculate the differential (euro area variable minus US counterpart) and then compute cumulative sums. When instead identifying US and euro area shocks separately (Section 4.3), I include variables in absolute (log) levels.

The main time sample ranges from January 2001 to March 2024. The beginning of the sample is chosen based on data availability and with a view to align the start across all model specifications.[18] The VARs are estimated with 12 lags, equivalent to one quarter in the weekly models and one year in the monthly version.

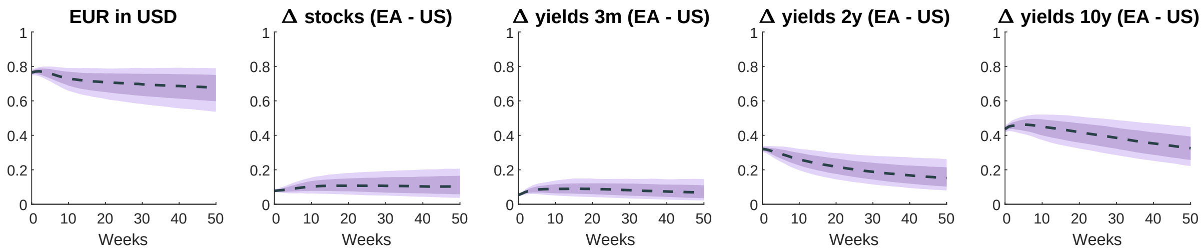
4.2 Relative Monetary Policy Shocks

This section establishes the main result: shocks to the relative stance of monetary policy are the main drivers of short-term nominal exchange rate fluctuations. In addition, I document the absence of any delays in overshooting.

Figure 2 shows impulse response functions (IRFs) to a relative monetary policy shock that increases the euro's value against the US dollar by one percent. It therefore represents a relative euro area monetary tightening or a relative easing of Fed policy. Notably, while the increase in the euro's relative value is persistent, the largest effect is measured directly on impact, with the effect gradually dissipating thereafter. Hence, there is no evidence of a pronounced hump-shaped response. Relative stock market valuations also

[18]The starting date turns out to affect the persistence of the USD-EUR rate response, with implications for the analysis in Section 4.4, but not for the main results, see Section 4.5.

Figure 3: FEVD FOR RELATIVE MONETARY POLICY SHOCK (WEEKLY DATA)



Note. Share of forecast error variance explained by the relative monetary policy shock. Weekly data. Shaded areas denote 68% and 90% confidence bands. Time sample: Jan. 2001 to Mar. 2024.

fall persistently by around half a percent. Interest rate differentials increase throughout all horizons, although the effect is more pronounced at medium to long rather than at the short end of the yield curve. Further, it is noteworthy that the shock increases interest rate differentials despite the fact that high-frequency changes in interest rates in this baseline specification have not been used to construct the instrument used for identification. Yet, changes in relative yields are overall modest (a few basis points) compared to the jump in exchange rates.

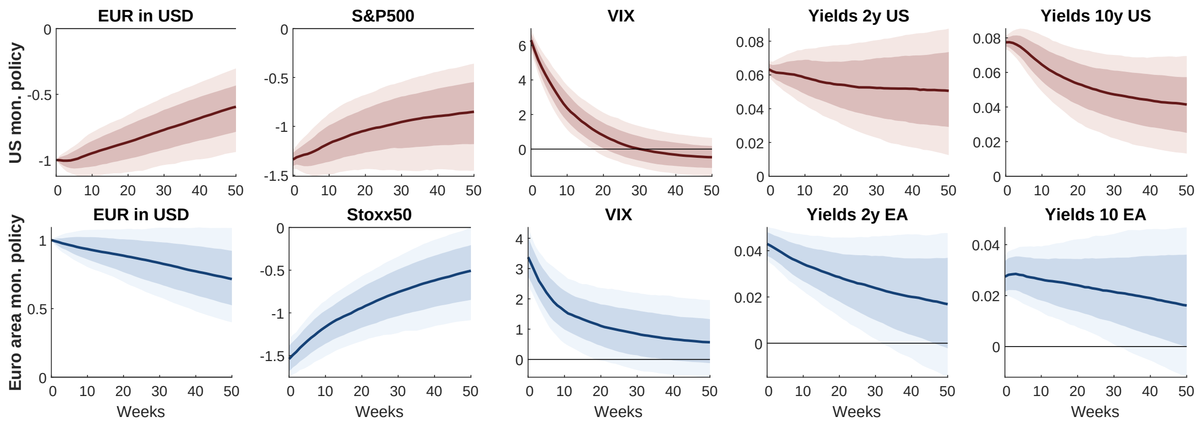
FIGURE 3 reports FEVDs, i.e. it shows how much of the forecast error of each model variable is accounted for by the identified relative monetary policy shock. Strikingly, around three quarters of USD-EUR exchange rate variability can be attributed to shocks to the relative stance of US and euro area monetary policy within the first month. In line with the peak responses on impact in FIGURE 2, the explanatory power is highest during the first few weeks but remains almost as high over a one-year horizon. In contrast, the relative monetary policy shock accounts for much smaller shares of the variance of the relative stock market performance. Notably, even interest rate differentials, through which monetary policy affects the exchange rate via standard, frictionless UIP-based theories, are driven much less by the relative monetary policy shock, with values ranging from below 10 on the short up to 40 percent at the long end of the yield curve.

These findings have at least two important implications. First, they do not support the notion that exchange rates are disconnected from macroeconomic aggregates in the sense that they are driven by a separate set of shocks. Instead, innovations in the relative monetary stance are even more important drivers of the exchange rate than for the relative evolution of other financial market prices commonly found to be affected by monetary policy.[19]

Second, the results suggest that monetary policy affects the exchange rate to an important degree through channels not captured by changes in (relative) risk-free interest rates alone. Indeed, recent theoretical contributions suggest that, in order to account for

[19] SECTION 4.4 will show that this finding extends to macro aggregates like prices and production.

Figure 4: IRFs to US and EA monetary policy shocks (weekly data)



Note. Impulse responses to a US (top, red) and euro area (bottom, blue) contractionary monetary policy shock, normalized to change the euro's value vs. the US dollar by 1 percent. Weekly data. Values in percent(age points). Shaded areas denote 68% and 90% confidence bands. Time sample: Jan. 2001 to Mar. 2024.

a variety of long-standing puzzles, exchange rates must be driven to a large extent by time-varying frictions that introduce wedges in standard UIP relationships (Itskhoki and Mukhin, 2021; Jiang et al., 2021; Eichenbaum et al., 2021; Jiang et al., 2024a). The findings here then indicate that monetary policy must give rise to, or interact with, these frictions to a significant degree. I will revisit this point in Section 4.4.

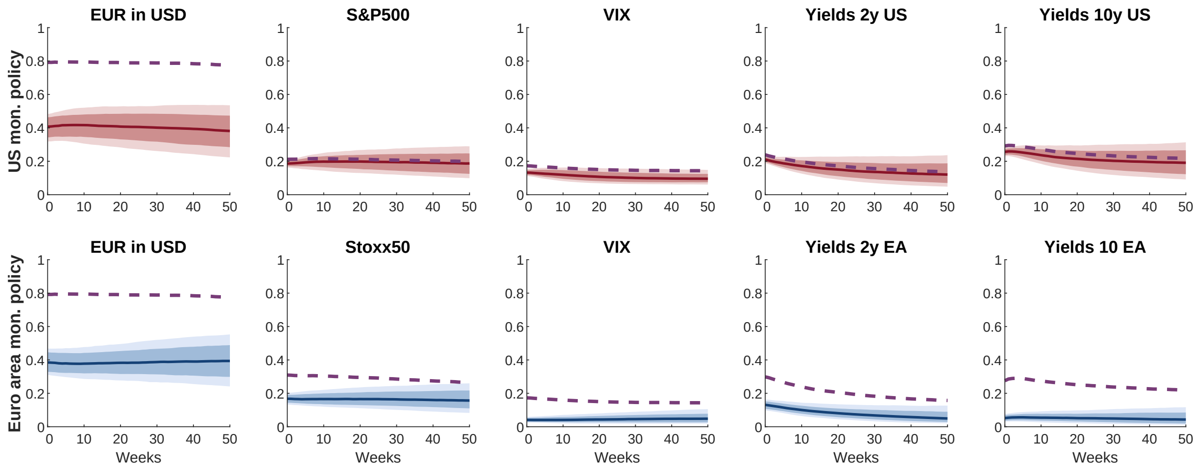
4.3 US and Euro Area Monetary Policy Shocks Separately

In this section I show that both US and euro area monetary policy are important in explaining variation in the USD-EUR rate. I do so by identifying monetary policy shocks in both currency areas separately but simultaneously, as explained in Section 3.

Figure 4 shows results for such an exercise. It depicts impulse responses of selected variables to contractionary US (top) and euro area (bottom) monetary policy shocks. As before, each shock is normalized to change the value of the USD-EUR exchange rate by one percent. Both shocks are contractionary in that they depress the respective stock market index by more than one percent, and increase implied stock market volatility by three to six percent, in line with the recent emphasis in the literature that monetary policy is an important driver of risk perceptions. In each case, the shock is associated with a persistent increase in two- and ten-year interest rates, again in line with responses obtained from standard identification based on interest rate instruments.

Figure 5 reports the corresponding FEVD for both shocks. Each panel shows two lines: the solid lines indicate the explained variance shares of the respective variable due to the US (top, red) and euro area (bottom, blue) shocks individually, while the

Figure 5: FEVD for US and EA monetary policy shocks (weekly data)



Note. Share of forecast error variance explained by the US (top, red) and euro area (bottom, blue) monetary policy shock. Dashed purple line indicates sum of median FEVD share of both shocks. Weekly data. Shaded areas denote 68% and 90% confidence bands. Time sample: Jan. 2001 to Mar. 2024.

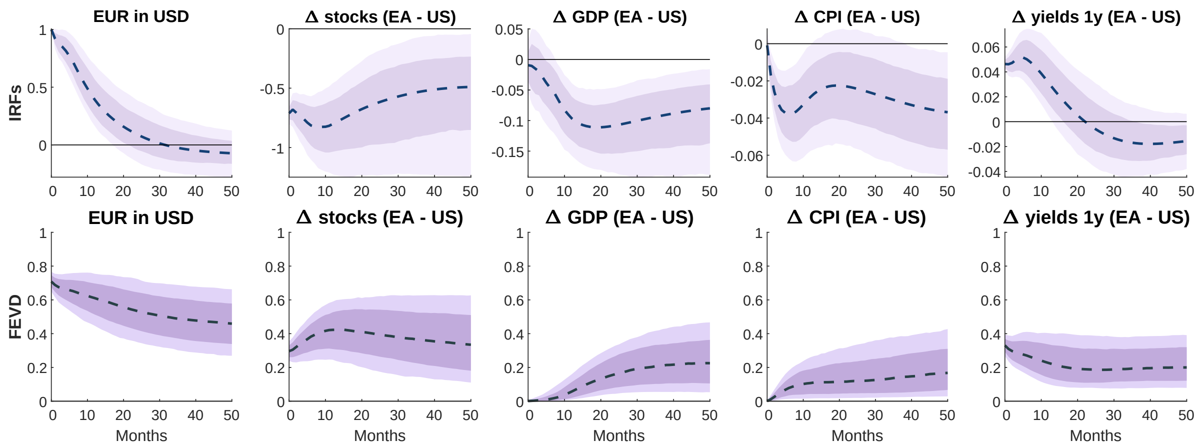
dashed lines in purple report the sum of both shocks.[20] They hence loosely correspond to the dashed lines in Figure 3. And indeed they paint a similar picture: more than three quarters of USD-EUR exchange rate variability can be attributed to both shocks combined, whereas the explained shares are much lower for the other variables, which this time enter in levels. Also the importance of monetary policy for variations in the VIX is quantitatively in line with the other (non-exchange rate) financial variables, as well as with findings in the literature.[21]

However, Figure 5 also reveals that there do not seem to be important differences when it comes to the relative importance of the two shocks for the exchange rate. Over a one-month horizon, shocks to US and euro area monetary policy account for roughly half of the overall variance share. Whereas US policy is generally more important for euro area markets than vice versa (as indicated by the difference between the solid and dashed lines in the other panels), this is not the case for the exchange rate. These findings hence do not indicate that the well-documented out-sized role of US monetary policy in determining conditions in international financial markets (Miranda-Agrippino and Rey, 2020) equally extends to USD-EUR variability. Notably, however, this result depends somewhat more on the exact model specification than the main finding of the large combined importance of both shocks, as I outline in Section 4.5.

[20] Whereas the purple lines in both rows are the same for the USD-EUR exchange rate and the VIX, they may differ for the other variables, as these are specific to each currency area.

[21] See, for instance, Bekaert et al. (2013), Bruno and Shin (2015), Miranda-Agrippino and Rey (2020) for estimates on the importance of US monetary policy for the VIX.

Figure 6: IRFs and FEVD in macro models (monthly data)



Note. Impulse responses (top) and share of forecast error variance explained by the relative monetary policy shock (bottom) in monthly model. Shaded areas denote 68% and 90% confidence bands. Time sample: Jan. 2001 to Mar. 2024.

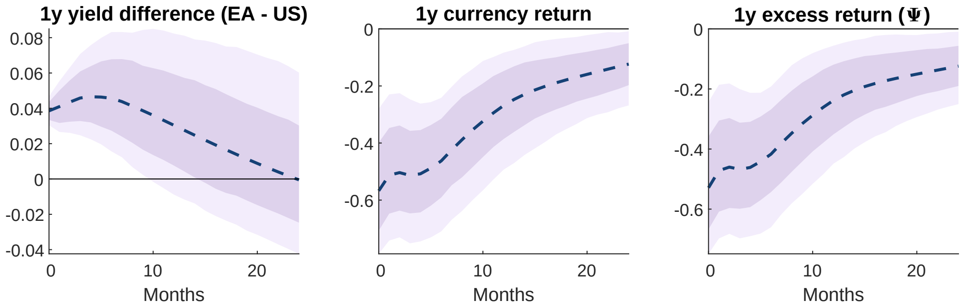
4.4 Conditional UIP and Macro & Long-Run Effects

In this section I extend the analysis in the financial market VAR in weekly frequency to a monthly VAR model including macro variables. This extension serves to study longer-run effects that cannot be captured in models using weekly data, and to verify that the identified shocks have effects on macro variables that are in line with theoretical predictions. More importantly, it also allows to test conditional UIP relations, revisiting the relation between exchange rate and interest rate responses mentioned above. After documenting the failure of frictionless UIP in response to the relative monetary policy shock, I then re-employ the weekly models from Sections 4.2 and 4.3 in order to study the role of convenience yields.

Macro and long-run effects. Figure 6 shows, in the top row, impulse responses to a relative monetary policy shock in the monthly relative model. It is noteworthy that also in monthly frequency, there is no sign of delayed exchange rate overshooting, not even within the first quarter, as for instance found in Rüth (2020). An appreciation of the euro due to a relative tightening by the Eurosystem leads, as before, to a fall in the value of European stocks compared to their US counterparts and to an increase in interest rate differentials. But also GDP and consumer prices fall relative to the US, in line with theoretical predictions that a monetary tightening in the euro area (or, equivalently, a monetary easing in the US) should lower European (stimulate US) economic activity relatively more strongly. These results therefore lend plausibility to the proposed identification scheme using the exchange rate-based instrument.

The bottom row of Figure 6 depicts the corresponding forecast error variance decom-

Figure 7: IRFs of yield differentials and (excess) currency returns (monthly data)



Note. Impulse responses of 1-year yield differentials (left), model-implied 1-year currency returns (middle) and 1-year excess currency returns (right panel) as in equation (6), based on the relative monthly model in Figure 6. Shaded areas denote 68% and 90% confidence bands. Time sample: Jan. 2001 to Mar. 2024.

position. In line with expectations and findings in the literature, monetary policy shocks account for only a fairly small fraction of the variation in (relative) GDP and prices, especially over short horizons. Most importantly, variation in the USD-EUR exchange rate is again accounted for to a large extent by monetary policy, with short-run shares of around 70 percent. Furthermore, even over long horizons the relative monetary policy shock explains almost half of USD-EUR variation. Hence, while other, non-monetary shocks certainly become more important in shaping the exchange rate over time, monetary policy remains a non-negligible driving force also over longer periods of time.

Conditional UIP. Finally, I can use the relative monthly model to conveniently test the validity of conditional UIP based on the estimated impulse response functions. Define currency excess returns from the perspective of a euro area investor as:

$$\Psi_{t,\tau} \equiv (i_t^{\$} - i_t) + (\mathbb{E}_t s_{t+\tau} - s_t), \tag{6}$$

in which $i_t^{\$}$ and i_t are the US and euro area risk-free interest rates from t to $t+\tau$, respectively, and s_t is (100 times) the log of the spot exchange rate of the USD-EUR exchange rate in price quotation, such that an increase corresponds to an appreciation of the US dollar against the euro.

The impulse responses in Figure 6 can then be used to assess whether excess returns respond to the identified shock:[22]

$$\frac{\partial \Psi_{t+h,12}}{\partial \epsilon_t} = \frac{\partial (i_{t+h}^{\$} - i_{t+h})}{\partial \epsilon_t} + \frac{\partial \mathbb{E}_t s_{t+h+12}}{\partial \epsilon_t} - \frac{\partial s_{t+h}}{\partial \epsilon_t}, \tag{7}$$

[22] Here I define one-year excess returns (such that $\tau = 12$ as in Yang et al., 2024), rather than annualized per-period returns (as in Eichenbaum and Evans, 1995 and Bjørnland, 2009), as the latter are more erratic and less precisely estimated.

Absent frictions, conditional UIP would predict that the size of the initial difference in relative risk-free yields corresponds to the depreciation of the domestic currency following the initial appreciation due to the shock, i.e. $\frac{\partial \Psi_{t+h,12}}{\partial \epsilon_t} = 0 \; \forall \; h = 0, 1, ..., H$.

FIGURE 7 shows the corresponding impulse responses, next to its two individual components. Note that, as in the figures throughout, the exchange rate is here defined in price quotation from the US perspective, i.e. it corresponds to the inverse of s_t in equation (7). The left panel depicts responses of the interest rate differential (the same as the right-most panel in FIGURE 6). The middle panel shows conditional currency returns, i.e. the exchange rate impulse response after one year compared to its impact value. The right panel then combines the two components. Clearly, the small increase in the relative yield by only a few basis points is overcompensated by the negative currency return, such that excess returns fall by roughly 50bp on impact, after which they gradually revert back to zero. In other words, investing in the higher yielding currency after the shock is *more than offset* by a subsequent depreciation.

THE ROLE OF CONVENIENCE YIELDS. The finding of conditional negative excess currency returns suggests that (relative) monetary policy shocks could introduce or interact with frictions in UIP. Following Jiang et al. (2021), one can show that excess currency returns Ψ_t can be decomposed into a risk premium and a relative convenience yield component, derived from first principles:

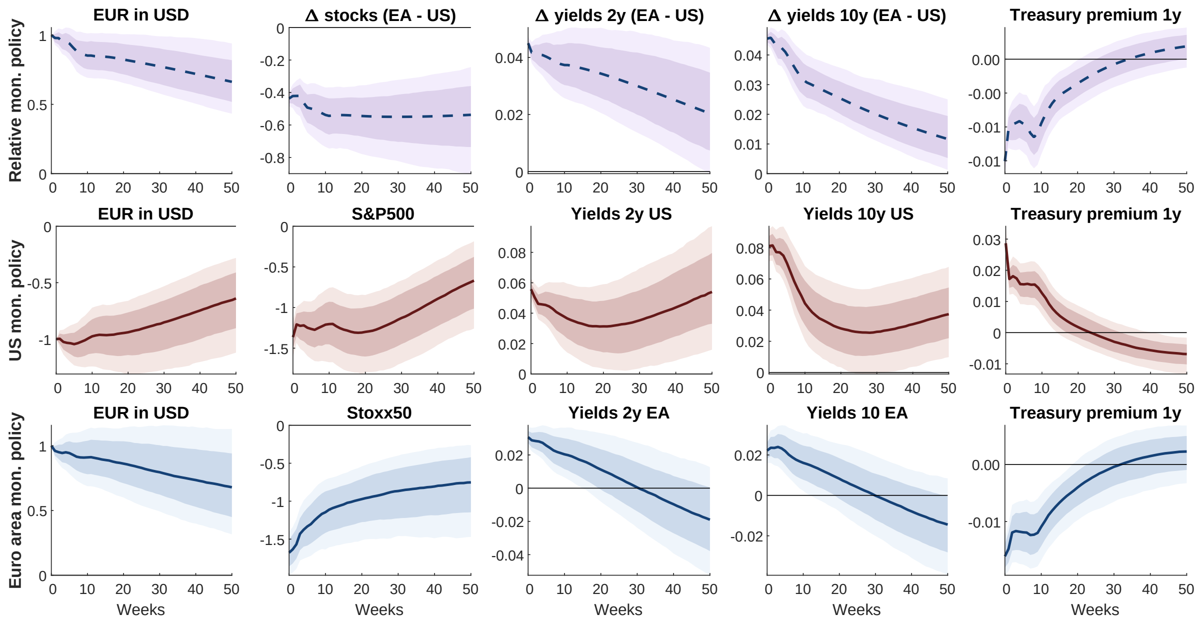
$$\Psi_{t,\tau} = (\lambda_t^{\$} - \lambda_t) - rp_t = (i_t^{\$} - i_t) + (\mathbb{E}_t s_{t+\tau} - s_t), \tag{8}$$

where rp_t is a composite term capturing the variance of the exchange rate and its covariance with the investor's stochastic discount factor, and $(\lambda_t^{\$} - \lambda_t)$ is the relative convenience yield, i.e. a non-pecuniary return that investors receive by investing in safe US dollar rather than non-USD bonds. This relative convenience yield is key for instance in the theories developed in Jiang et al. (2023) and Jiang et al. (2024b). In particular, in Jiang et al. (2023), tighter US monetary policy leads to a reduction in US safe assets, increasing their convenience yield and contributing to US dollar appreciation over and above what the frictionless UIP mechanism via the risk-free interest rate gap would predict.[23]

Not discarding the potential role of currency risk premia, in the following I focus on the convenience yield channel as an explanation of the findings of negative conditional excess currency returns documented above. Jiang et al. (2021) show that relative convenience yields can be measured by the (negative) treasury premium, defined as the deviation from covered interest parity between government bond yields. I therefore add estimates of the

[23]Relatedly, in Akinci and Queralto (2024), tighter US monetary policy also leads to deviations from frictionless UIP and a stronger rise in the dollar's value via financial accelerator mechanisms operating through emerging market balance sheets.

Figure 8: IRFs in models with treasury premium (weekly data)



Note. Impulse responses to a relative monetary policy shock in the relative model (top row, purple), and US (middle, red) and euro area (bottom, blue) contractionary monetary policy shock in the levels model, normalized to change the euro's value vs. the US dollar by 1 percent. Shaded areas denote 68% and 90% confidence bands. Time sample: Jan. 2001 to Mar. 2021.

treasury premium between US and German 1-year government bonds, obtained from Du et al. (2018a), to the weekly VAR models employed earlier.

Figure 8 shows the results. In each row, the first four panels look very similar to the results in Figures 2 and 4.[24] The right-most panels show the response of the treasury premium. Following the relative shock (top row), the premium falls, indicating that a relative monetary policy easing in the US lowers the convenience yield on US government bonds. This is confirmed in the levels model in the middle row: a US monetary policy tightening increases the treasury premium. Interestingly, the opposite is true following the euro area tightening in the bottom row. In this case, the treasury premium falls, indicating that contractionary euro area monetary policy likewise increases convenience yields of safe euro area bonds, leading to a decline relative to its US counterpart.

Qualitatively, the results therefore align with theories that stress the importance of fluctuations in convenience yields for exchange rate behavior. Notably, these mechanisms play a role not only for US dollar assets and US monetary policy but also for their euro area counterparts.[25] When assessing the responses quantitatively, it is important to note

[24]The main reason for the slight differences lies in the fact that the model is estimated only until March 2021, as the treasury premium time series in Du et al. (2018a) is not available afterwards.

[25]See also Du et al. (2018b). This is not to suggest that US monetary policy and the US dollar are not special. Indeed, when identifying the US monetary policy shock in a monthly model, it explains the majority of the short-term forecast error variance of Miranda-Agrippino and Rey's (2020) global financial cycle time series – whereas euro area monetary policy contributes only marginally –, substantiating theories in Jiang et al. (2023) and Georgiadis et al. (2024).

that the treasury premium is a measure of, but is not identical to, the relative convenience yield. Indeed, Jiang et al. (2021) estimate the relation $-x_t = (1-\beta)(\lambda_t^{\$} - \lambda_t)$ and find that β is around 0.9, indicating that the actual convenience yield is roughly ten times larger than the (negative) treasury premium x_t. Applying this estimate to the results in Figure 8 indicates that the relative convenience yields would move on impact by 10bp to 30bp following a monetary policy shock that moves the USD-EUR rate by one percent. They would then – slight differences in model specifications notwithstanding – represent a sizable portion of the roughly 50bp impact response in excess currency returns estimated in Figure 7.

4.5 Additional Results and Robustness

This section presents results when varying model parameters along several dimensions in order to address robustness considerations and discuss further interesting results. The main focus will be on the paper's central finding of the high importance of relative monetary policy shocks for exchange rate variability (Section 4.2). In addition, I will relate to the results in Sections 4.3 and 4.4 where applicable. Some of these exercises will also be useful to ascertain why the importance of (relative) monetary policy shocks is estimated to be so much higher than in previous studies. Short of conducting an actual accounting exercise, I conclude that the differences are a function of i) taking into account monetary policy shocks of both central banks simultaneously, ii) using an exchange rate instrument for identification, or at the very least one that does not only pick up variation in short-term interest rates, and iii) the time sample under consideration.

Table 2 provides an overview of the explained forecast error variance of the USD-EUR exchange rate over a one-month horizon[26] in the relative model when changing one model dimension at a time. Overall, across all specifications the share of the FEVD is high. As gleaned from Figure 3, in the weekly benchmark model the explained fraction is slightly above 75 percent (1st column). Notably, the share also remains high when estimating the model using much more volatile daily (2nd column) or monthly data (3rd column, as in Figure 6 above). In line with the high-frequency analysis in Figure 1, the share is lower when estimating the model only until before the collapse of Lehman Brothers in September 2008 (4th column), after which central banks around the world employed unconventional monetary policy tools. This finding indicates that one of the reasons of the much higher importance of monetary policy in this paper – spanning the period from the early 2000s to 2024 – compared to the previous literature – often going back many decades and not including the post-financial crisis period – can not only be found in the econometric setup but also in the time sample under consideration. The

[26] I choose a one-month horizon in order to be able to compare results across the models with different data frequencies.

Table 2: USD-EUR FEVD IN DIFFERENT MODEL SPECIFICATIONS

Benchmark	**Daily**	**Monthly**	**Pre Sep. 2008**	**Interest rate instr.**	**No instr. cleansing**
76.9	71.2	70.8	57.3	65.5	85.4

Note. The table reports the share of the FEVD of the USD-EUR exchange rate accounted for by the relative monetary policy shock over a one-month horizon when varying various model specifications one at a time. For instance, the first column (benchmark) reports the average of the first four weekly observations in FIGURE 3, the second column reports the equivalent (first 21-day) figure in a financial market variable model using daily data etc. Details on the other models are provided in the main text below.

remaining two specifications in TABLE 2 are discussed in the following in more detail.

INTEREST RATE INSTRUMENT. The main specifications in SECTIONS 4.2 and 4.3 use high-frequency changes in the USD-EUR exchange rate as an instrument for shock identification, rather than changes in interest rates. As argued, this choice reflects the idea that monetary policy does not necessarily affect exchange rates only through changes in risk-free interest rates. Still, it is interesting to compare results to those obtained using the more standard interest rates-based approach in the literature. This will not only again serve to validate the baseline identification but also yield additional results that speak to the relative impact of monetary policy on exchange and interest rates. Hence, as a robustness check I replace the high-frequency responses of the USD-EUR exchange rate with an interest rate series when computing the instrument.[27]

A second reason that motivated the choice of the exchange rate instrument in the benchmark model was the ambiguity with respect to which interest rate maturity to use. Whereas traditionally monetary policy was communicated mostly in terms of setting very short-term interest rates, the increased reliance on forward guidance and asset purchases as policy tools necessitates to also take into account policy effects on other parts of the yield curve. Hence, in order to capture monetary policy surprises as comprehensively as possible, I compute the first principal component of yields at several maturities rather than at a single one.[28] In the baseline case I use changes in two-year and ten-year government bond yields during the monetary policy windows, for which I have data over the entire estimation sample.[29]

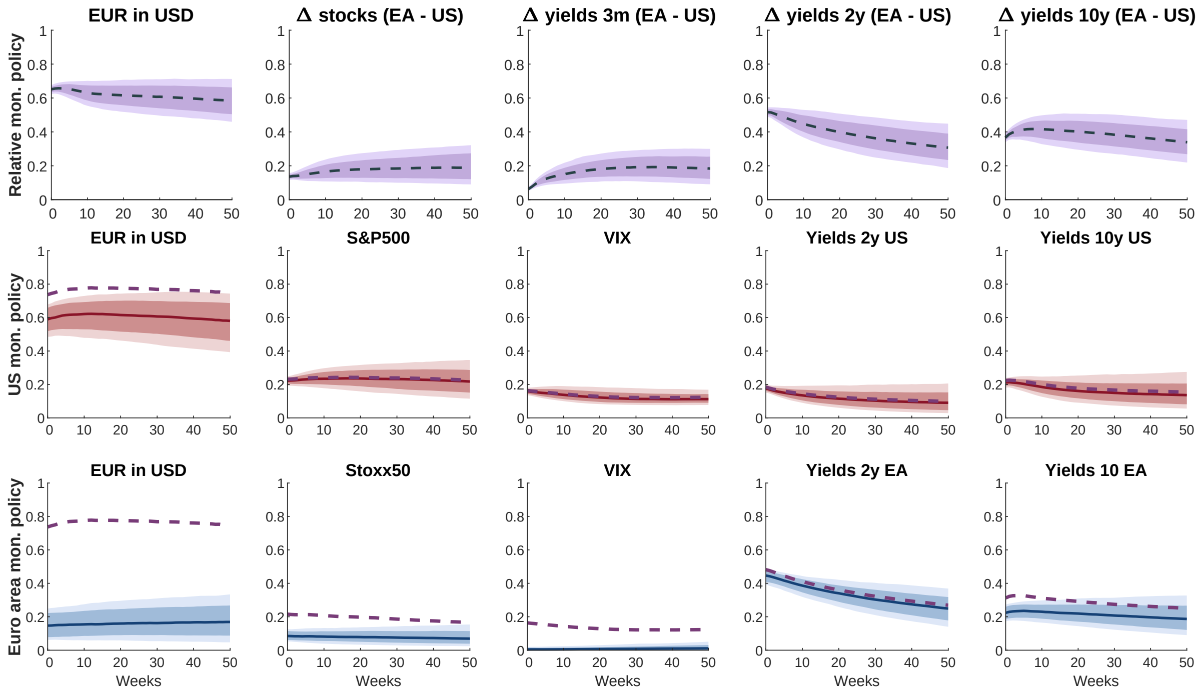
FIGURE C.1 confirms that impulse responses look very similar to before, again for both

[27]As before, I cleanse the series from potential information effects using stock market responses.

[28]See Nakamura and Steinsson (2018) and Bauer et al. (2023) for a similar approach combining information on short-term yields.

[29]The first principal component then explains more than 80 percent of the two individual series in both currency areas.As an additional robustness test, I compute the instrument using additionally high-frequency changes of 3-month interest rates. Due to more limited data availability, in this case I can only compute the instrument until the end of 2019. What is more, for the euro area, data on changes in 3-month German government bond yields in the database of Altavilla et al. (2019) become available only in 2005. For the period before 2005 I therefore use responses in 3-month OIS rates. The first principal component then explains still more than 60 percent of the three individual series (3-month, 2- and 10-year rates). Results are very similar to the baseline case.

Figure 9: FEVD with interest rate instrument (weekly data)



Note. Share of forecast error variance explained by the US (top, red) and euro area (bottom, blue) monetary policy shock. Shocks identified with interest rate instead of exchange rate instrument. Dashed purple line indicates sum of FEVD of both shocks. Weekly data. Shaded areas denote 68% and 90% confidence bands. Time sample: Jan. 2001 to Mar. 2024.

the relative model and the separately identified shocks. More central for the question at hand, Figure 9 shows the corresponding FEVD for the model identified with the interest rate-based instrument. Here, results are more nuanced. First, also in this case (relative) monetary policy shocks account for most of the variance of exchange rate forecast errors. Further, the shocks explain exchange variation again to a higher degree than changes in any section of the yield curve. I verify, however, that these results are less robust than when using the exchange rate instrument. For instance, the combined FEVD share of the two monetary policy shocks fall below 50 percent when using daily or monthly instead of weekly data.

Second, results regarding the relative importance of the two monetary policies are very different from those found in Section 4.3: when using the interest-rate instrument, US monetary policy appears to be roughly three times more important than its euro area counterpart as a driver of the USD-EUR rate.[30] This result is noteworthy as it underscores the finding in Section 4.4: if risk-free interest rate differentials were the only drivers of the response of the exchange rate, one would c.p. expect euro area monetary policy to

[30] Roughly the same is true when using a tight window of -10 to +20 minutes around monetary policy announcements, indicating that accounting for press conferences is especially important for euro area monetary policy.

play a larger, rather than a smaller, role when switching to an interest rate instrument. This is because euro area yields respond to US monetary policy much more than vice versa (see FIGURE C.1). A given surprise rate hike in the euro area then opens up larger gaps in yields across the two currency areas than a comparable rate hike in the US. Euro area monetary policy should then be able to affect the exchange rate more, if the mechanism indeed exclusively ran via interest rate differentials.

NO CLEANSING FROM INFORMATION EFFECTS. In line with recent developments in the literature on the identification of monetary policy shocks, the benchmark identification scheme cleanses the high-frequency instrument from potential information or signalling effects. In the following, I repeat the analyses using an uncleansed instrument. FIGURE C.2 reports IRFs and FEVDs in the weekly relative model when constructing the instrument from high-frequency exchange rate responses alone, i.e. without using stock market responses. Qualitative responses are very similar to those in the benchmark model in FIGURE 2. Interest rate differentials increase throughout the yield curve following the relative euro area tightening. Stock prices fall relatively more strongly in the euro area, although quantitatively the effect is significantly smaller than before. This quantitative difference translates to the FEVDs: compared to the benchmark results in FIGURE 3, also the explained share of the variance of relative stock prices is somewhat smaller. Remarkably, however, the share is even larger for the USD-EUR exchange rate, reaching as much as 85 percent on average during the first month. FIGURE C.3 repeats the exercise for the monthly model, with again very similar results to before, both in terms of IRFs and FEVD.

SHOCK INVERTIBILITY. The empirical results are based on a standard proxy, or 'external instrument' VAR model (Stock and Watson, 2018). As outlined recently by several authors (Plagborg-Møller and Wolf, 2022; Miranda-Agrippino and Ricco, 2023), this standard approach implicitly assumes shock invertibility – a questionable assumption in many empirical applications.[31] It is therefore important to check if the results here are driven by unduly assuming shock invertibility. I conclude that this is unlikely to be the case based on the following observations.

First, Miranda-Agrippino and Ricco (2023) show that non-invertible shocks produce instable impact impulse responses when changing the model specification by removing or adding variables or changing the lag length. Experimenting with changing lags and the set of controls I generally find the impact responses to be highly stable, and especially so in the relative model. Second, both Miranda-Agrippino and Ricco (2023) and Forni et al. (2023) demonstrate how monetary policy shocks identified from instruments that are not

[31] Technically, invertibility means that a shock can be recovered as a linear combination of the present and past values of the VAR variables.

cleansed from information effects are more susceptible to being non-invertible. As outlined above, I make sure to remove potential information effects. Third, Plagborg-Møller and Wolf (2022) suggest, as an alternative to the standard proxy (external instrument) approach the use of an internal instrumental model. This entails to include the instrument as a model variable ordered first and estimate impulse responses recursively. If these responses differed, this would be a sign of non-invertibility of the identified shock using the standard external instrument method. Performing this comparison in the relative model, I obtain almost identical responses, albeit with somewhat wider confidence bands when using the internal method (see Figure C.4). Finally, when performing the formal statistical test proposed by Forni et al. (2023) in the relative benchmark model, I can generally not reject the null hypothesis of shock invertibility.

Time Sample I verify that results are similar when estimating the models only up to the COVID-19 pandemic in early 2020. In particular, results of the monthly macro model are only mildly affected by the pandemic period on account of specifying the variables in relative terms.[32] Other than that, the precise time sample turns out to matter to the extent that the start of the sample heavily determines how persistent the USD-EUR impulse responses are. Indeed, when estimating the model starting in 1999 instead of 2001, responses are more, when starting in 2003 or later they are less persistent than in the baseline case. This does not matter for the impact FEVD shares, yet it does affect the results in Section 4.4. As the size of conditional UIP deviations are primarily driven by the slope of the USD-EUR impulse response, a more persistent response c.p. translates into less negative excess returns. In a model starting in 1999, excess returns then only amount to around -20bp on impact and are statistically significant only when increasing the lag length of the model. On the other hand, excess returns are more significant and fall to more than -50bp when starting the model in 2003 or later.

Time Sample and Further Robustness Considerations. The main result is further robust along a number of dimensions, for instance with respect to the lag lengths in the VAR models or the method of cleansing from information effects (the *poor man's* approach of Jarociński and Karadi (2020) or using sign restrictions). Adding various financial spreads – as emphasized e.g. in Caldara and Herbst (2019) – to the models in levels also does not affect the main result. The same is true when using tighter windows around the monetary policy announcement (-10 to +20 minutes), which then include only the press releases but not the subsequent press conferences. Further, main results hardly change – and monetary policy continues to explain a very high fraction of exchange rate

[32] As macro variables in both the US and the euro area were initially affected to a similar degree by the pandemic, the relative time series to not feature large jumps that have been causing parameter instability in conventional macro VARs in levels, see e.g. Lenza and Primiceri (2022).

variability – when estimating the relative model with non-stationary variables in first differences.

5 Other Exchange Rates: Relative Monetary Policy and the Pound & Yen

This section applies the baseline empirical approach to exchange rates involving two additional currencies: the British pound (GBP) and the Japanese yen (JPY).

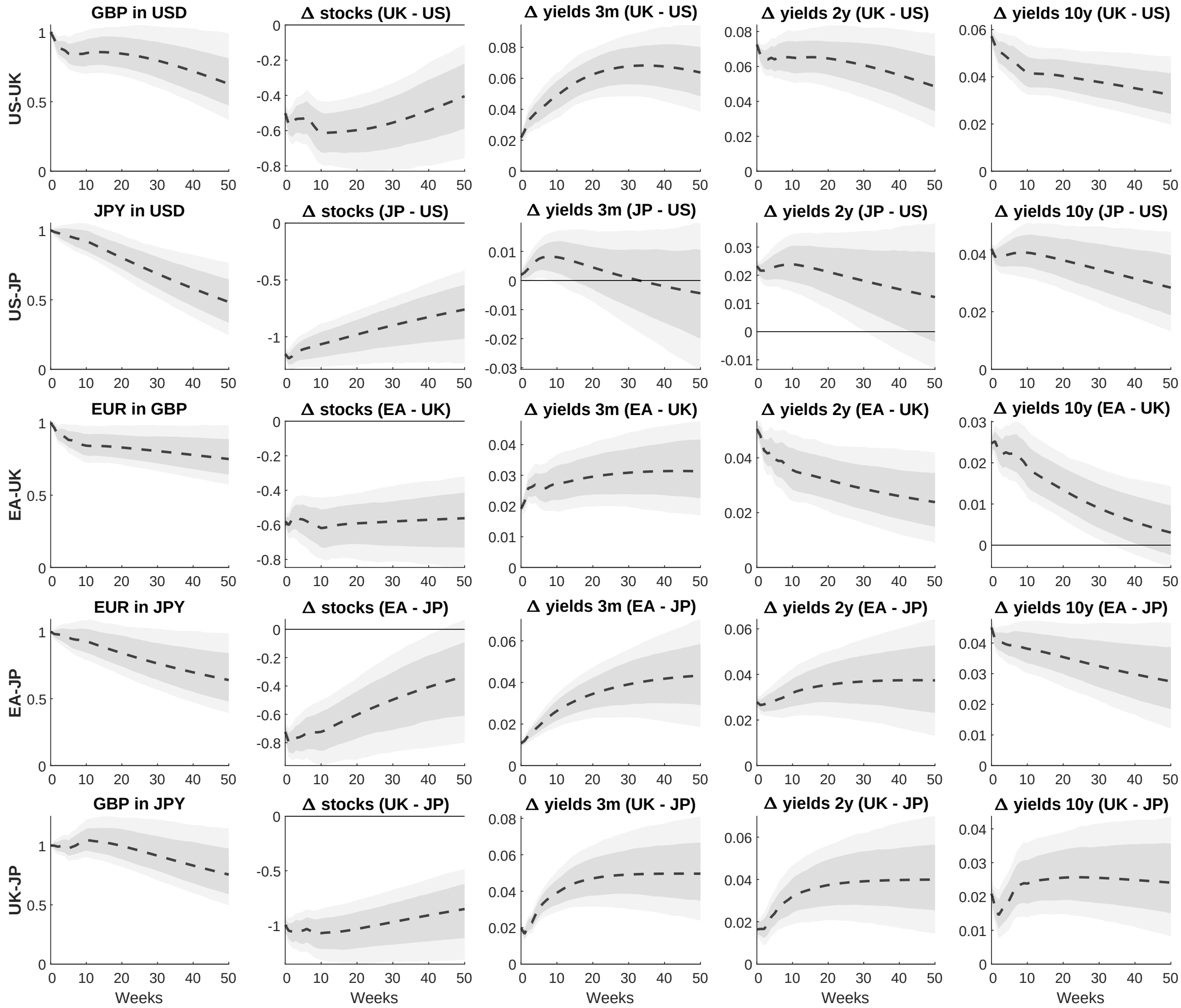
This extension serves two purposes. First, from a methodological point of view, it is useful as another implicit check of the empirical approach. For instance, if it was to yield similar, plausible impulse responses, this would speak to the validity of the methodology. Second, from an economic point of view, it will be interesting to see if the results so far are specific to the USD-EUR exchange rate. In particular, conducting the analysis using exchange rates not involving the US dollar will be informative. If one was to find that relative monetary policies still explain much of exchange rate variation, this would add weight to the notion that not only US policy plays a major role in exchange rate determination.

To that end, I compute the same relative financial market variables, and construct instruments for identification, for each combination of currencies and relative monetary policies.[33] This results in five additional models involving the exchange rates USD-GBP, USD-JPY, EUR-GBP, EUR-JPY and GBP-JPY, for each of which I identify a relative monetary policy shock. Appendix A includes Table A.2 that provides an overview of the model specifications and data sources. It also provides a description of the instrument data, for which I collect, and partly compute myself, high-frequency asset price responses of an additional 458 monetary policy announcements (216 for the Bank of England and 242 for the Bank of Japan).

The instruments are then constructed along the same lines as before: taking into account press conferences, I measure the high-frequency response of the respective exchange rates around monetary policy announcements by the respective central banks, and cleanse the series from information effects and sum daily observations to arrive at instruments in weekly frequency. As high-frequency data availability for announcements by the Bank of Japan and Bank of England is more limited, the instrument series begin

[33] I limit the extension to the baseline approach to preserve space but also for reasons of data availability. For instance, the database of Cieslak and Schrimpf (2019) includes only limited information on how different interest rates respond to monetary policy announcements by the Bank of England and the Bank of Japan. I am however able to extend the database with information on high-frequency exchange rate and (for the post-2017 period) stock price responses. See the discussion in Appendix A for details. The extensive UKMPD database by Braun et al. (2023) will be incorporated in future versions of this paper.

Figure 10: IRFs to relative monetary policy shocks: other currencies (weekly data)



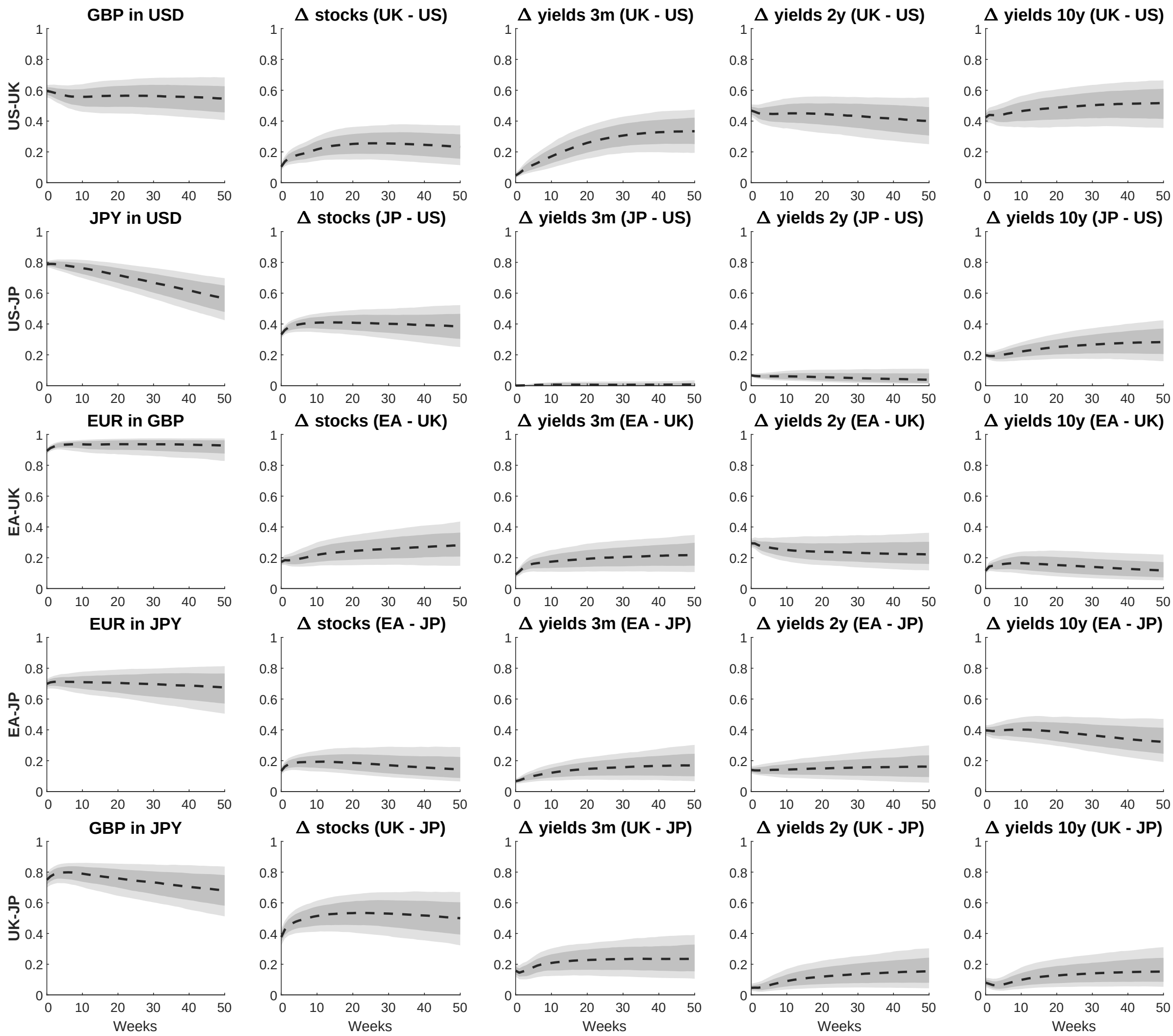
Note. Impulse responses to relative monetary policy shocks in weekly financial market models. Each row reports results for a bilateral model that identifies a relative monetary policy shock involving the two central banks issuing the respective currencies. Remaining details as in Figure 2.

somewhat later such that they cover the period from May 2003 to May 2023. In order to aid comparability to the results so far, however, I still compute all VAR models using time series beginning in 2001.[34]

Figure 10 depicts impulse responses, equivalent to those in Figure 2. Each row represents a different model in which a relative monetary policy shock is identified involving the two central banks that issue the currencies whose exchange rate response is shown in each first panel. Again, responses are normalized to increase the relative value of one of the currencies by one percent. Comparing the results to those for the USD-EUR exchange rate in Figure 2 reveals a robust pattern: in all models, relative stock prices fall by a similar magnitude and interest rate differentials increase throughout the yield

[34]The proxy VAR framework has the advantage that the available instrument time series can be shorter than the time sample covered by the VAR model variables.

Figure 11: FEVDs for relative monetary policy shocks: other currencies (weekly data)



Note. Share of forecast error variance explained by the relative monetary policy shocks in weekly financial market models. Each row reports results for a bilateral model that identifies a relative monetary policy shock involving the two central banks issuing the respective currencies. Remaining details as in Figure 3.

curve, although the yield effects differ somewhat quantitatively and in their persistence. Notably, also in the additional five models here, the response of the exchange rate is free from any delay in overshooting. Further, sizable jumps in the exchange rate are associated with only small changes in relative interest rates. All in all, the empirical method delivers qualitatively identical responses, underlying the robustness and utility of the approach.

Figure 11 shows the corresponding FEVDs. Results are overall again very similar to the findings for the USD-EUR exchange rate. Most importantly, just as in Figure 3, the variation of the exchange rate is explained best, with shares higher than those for yield differences, no matter the maturity. The same is true for relative stock market

Table 3: FEVD FOR DIFFERENT EXCHANGE RATES

	Fed/USD	ECB/EUR	BoE/GBP	BoJ/JPY
Federal Reserve (Fed) / USD		76.9	58.7	78.9
European Central Bank (ECB) / EUR	76.9		91.4	70.8
Bank of England (BoE) / GBP	58.7	91.4		77.5
Bank of Japan (BoJ) / JPY	78.9	70.8	77.5	

Note. The table reports the share of the FEVD of different exchange rates accounted for by relative monetary policy shocks over a one-month horizon in the weekly financial market model. For instance, the first column lists the shares of the USD-EUR, USD-GBP and USD-JPY exchange rate forecast errors that are explained by the relative monetary policy shock of the US and the euro area, the US and the UK, and the US and Japan, respectively. Results are symmetric by construction. Values in percent. Time sample: Jan. 2001 to Mar. 2024.

performances, which relative monetary policy explains to a significantly smaller degree.

TABLE 3 summarizes the FEVD results for all six models (i.e., including the benchmark USD-EUR model). Depending on the currency pair, short-term shares range from more than 50 (GBP-JPY) to around 90 percent (JPY-USD, EUR-GBP), and average to roughly 75 percent. As outlined, these findings have implications for the special status of the US dollar and US monetary policy. Despite the outsized role of US compared in determining international monetary and financial conditions, it is still the case that central banks other than the Fed play an important role in shaping the value of their own relative to other currencies. Moreover, also they seem to affect exchange rates through means other than changes in interest rate differentials alone.

6 Conclusion

In this paper I provide empirical evidence that the majority of short-term exchange rate fluctuations can be attributed to changes in the relative stance of monetary policy. Compared to the existing literature, I identify monetary policy shocks comprehensively by taking into account both involved central banks as well as many potential transmission channels. Throughout I place high emphasis on the robustness of the results. Shocks to monetary policy are found to account for roughly 75 percent of the variation of the USD-EUR exchange rate over a one-month horizon in the baseline specification, with estimates ranging from somewhat above 50 to over 80 percent, also for other exchange rates involving the yen and pound. Strikingly, the importance of monetary policy for exchange rates exceeds that of interest rate differentials, and sizable jumps in exchange rates are associated with only small changes in risk-free interest rate gaps – leading to conditional UIP deviations, indicating that monetary policy must have important effects on the exchange rate over and above what frictionless UIP considerations would suggest.

I interpret my findings in light of recent developments in two strands of the literature. First, various theoretical contributions on exchange rate determination stress the importance of shocks that introduce wedges in otherwise standard UIP relations (Jiang et al., 2021; Itskhoki and Mukhin, 2021; Itskhoki and Mukhin, 2023). Yet the source of these shocks remains largely elusive. The results in this paper indicate that monetary policy must be one important driver of these wedges. Second, this substantiates a recent trend in the empirical literature of monetary policy transmission that extends beyond effects via risk-free interest rates (Kroencke et al., 2021; Bauer et al., 2023; Boehm and Kroner, 2023). Altering the willingness of market participants to bear risk and affecting safe asset demand and convenience yields are important channels of monetary policy transmission to exchange rates, which should be incorporated in open economy models.

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A Data

A.1 High-frequency Data for the Instruments

This section provides some more details on the different instrument series used for identification in the VAR analysis that are constructed from high-frequency asset price responses around monetary policy announcements. Most of the description also applies to the data used in the high-frequency analysis in Section 2.[35]

FOMC (US Federal Reserve). For the Federal Open Market Committee (FOMC), I make use of the (updated) databases by Gürkaynak et al. (2005) and Cieslak and Schrimpf (2019), and extend them manually for the later parts of the sample.

The Fed began in 2011 to regularly hold press conferences to explain the preceding press statements and provide guidance on the policy outlook. Market price responses to this information should ideally be included when constructing the instrument, yet the database by Gürkaynak et al. (2005) only covers half-an-hour windows around the press statement. In contrast, the data by Cieslak and Schrimpf (2019) cover different time windows such that responses can be constructed that include both press statements and press conferences. Yet, it does not cover exchange rate responses. I hence procede as follows. For the time period before 2011, I use high-frequency responses from Gürkaynak et al. (2005). For the time period since 2011, I use Cieslak-Schrimpf data for interest rate and stock price responses and construct exchange rate responses (USD-EUR, USD-GBP, USD-JPY) on my own using minute-by-minute data from `tickstory.com`.[36] Neither of the existing databases covers the period after 2019. For that part of the sample I hence compute all asset prices responses on my own, using data from Refinitiv and again `tickstory.com`. In line with the literature, the FOMC announcement window begins 10 to 15 minutes prior to the press statement. In case there is no subsequent press conference, the window ends 15 to 20 minutes after the press statement, otherwise it ends 60 minutes after the start of the press conference. In total, the instrument data set contains asset price responses to 204 FOMC announcements from January 1999 to May 2023.

ECB Governing Council. For ECB Governing Council announcements things are more straightforward and I rely on the database by Altavilla et al. (2019). It has the ad-

[35] There are some slight differences when it comes to specifying the time windows around monetary policy announcements. These, however, are minor and only due to the fact that in Sections 4 and 5 I generally align the windows with those of the databases I take high-frequency interest rate responses from, for instance starting 15 instead of 10 minutes prior to the event.

[36] To that end, I compute 10-minute intervals around the starting and ending points of the announcement window, and then calculate responses as percentage changes of the median price quote within the two intervals.

vantage that it includes press conferences in the announcement windows from the outset, beginning 15 minutes prior the press statement and ending 75 after the start of the subsequent press conference. The database is also comprehensive in coverage and is regularly updated. The edition used in this paper covers high-frequency responses for interest rates at different maturities, stock prices (EuroStoxx50) and the USD-EUR, GBP-EUR and JPY-EUR exchange rates until the end of 2021. Hence, I use self-computed response data only for the period since 2022, again based on minute-by-minute raw data from Refinitiv and `tickstory.com`. This results in the coverage of high-frequency responses to 295 ECB Governing Council announcements from January 1999 to May 2023.

MPC (Bank of England). For market responses to Monetary Policy Committee (MPC) announcements I extend the database by Cieslak and Schrimpf (2019). It covers 2-year and 10-year interest rate and stock price responses for the period from 2003 to December 2017, but no responses of exchange rates. I hence again use minute-by-minute `tickstory.com` data to compute changes in stock prices (FTSE100, beginning in 2018) and all relevant exchange rates (GBP-USD, GBP-EUR, GBP-JPY, beginning in 2003). As in the case of the Fed, I use half-hour windows around the publication of the monetary policy decision if there was no subsequent press conference, but extend the window to 60 minutes after the beginning of the press conference if there was one. This results in the coverage of 216 MPC announcements from May 2003 to May 2023.

MPM (Bank of Japan). For Monetary Policy Meetings (MPM) by the Policy Board of the Bank of Japan, I again extend the database by Cieslak and Schrimpf (2019). It covers 10-year interest rate and stock price responses until December 2017, but again not for exchange rates. Using minute-by-minute `tickstory.com` data, I therefore compute responses of stock prices (Nikkei225, beginning in 2018) and all relevant exchange rates (JPY-USD, JPY-EUR, JPY-GBP, beginning in 2003). As before, I use half-hour windows in case there was no press conference after the press statement. Taking into account press conferences is complicated by the fact that these usually start only several hours (most often more than three) after the publication of the monetary policy decision. There are two potential ways of how to deal with this issue. One could compute two responses for two time windows (one for the policy decision, one for the press conference) or simply extend the window such that it includes both press statements and press conferences. I opt for the latter approach since there are generally no important data releases between the two events that would result in strong asset price movements unrelated to Japenese monetary policy. I end up with a coverage of asset price responses to 242 policy announcements from May 2003 to May 2023.

Table A.1: Sign Restriction Identification for Proxy VAR Instrument

		Monetary	Information
High-frequency response	Exchange rate	+	+
	Stock market index	−	+

Note. Sign restrictions to create instrument series $\mathbf{z}_t$ (see equations (3) and (4) in the main text) for monetary and central bank information shocks in the model for which two monetary policy shocks are identified separately.

Cleansing from information effects. Having obtained the high-frequency asset price responses, I cleanse them from central bank information effects in the spirit of Jarociński and Karadi (2020). Afterwards, I aggregate the resulting instrument series to weekly and monthly frequency, where applicable, by simply summing the daily values.

Some subtleties emerge depending on whether a relative instrument is constructed or two separate ones. When constructing two separate instrument series, I use exchange rate and domestic stock price surprises for each of the two central banks, and feed these into a sign restriction procedure as indicated by Table A.1.[37]

For the relative instrument, exchange rate and stock price surprises for all monetary policy announcements are taken together, and only one instrument is created. Employing again a sign restriction procedure would, strictly speaking, require to measure how not only exchange rates change around monetary policy announcements, but also *relative* stock prices. However, these cannot be computed as high-frequency response data for foreign stock prices is not available to me for much of the time sample – e.g. for European stocks around FOMC announcements. For that reason, I use the *poor man's* approach of Jarociński and Karadi (2020), in which I simply remove all observations in which the sign of exchange rate response coincides with the sign of the respective domestic stock price response. Notwithstanding these considerations, I obtain very similar results with the sign restriction procedure implicitly assuming no change in non-domestic stock prices.[38]

A.2 VAR Data

Table 1 in the main text lists the variables included in Section 4, alongside their data sources. Financial market variables refer to end-of-week or end-of-month values, respectively, constructed from daily data. In the baseline specifications, all variables

[37]The responses of the exchange rate are replaced by interest rate responses in Section 4.5. In the same section I also conduct a robustness exercise where I do not cleanse from information effects such that no sign restriction procedure is used and z_t is simply the exchange (or interest rate) response.

[38]Indeed, for the later parts of the sample, for which I have data, I verify that the domestic stock market response to domestic monetary policy news is generally much more pronounced than that of foreign stock prices.

enter in levels and are logged, with the exception of the interest rate series. Additionally, as outlined in the main text, the data enter in relative terms in the relative models, with the exception of the exchange rate, which is already a relative price. Data used in Section 5 is constructed equivalently, Table A.2 provides an overview of the time series in each model and the data sources.

Table A.2: DATA AND VAR MODEL SPECIFICATIONS IN SECTION 5

Variable	**Source**	(1)	(2)	(3)	(4)	(5)
USD-GBP exchange rate	Refinitiv (LSEG)	•				
2y gov. US bond yields	Bloomberg	•				
2y gov. UK bond yields	Bloomberg	•				
10y gov. US bond yields	Bloomberg	•				
10y gov. UK bond yields	Bloomberg	•				
3m gov. US bond yields	Reuters	•				
3m gov. UK bond yields	Refinitiv (LSEG)	•				
S&P500	S&P	•				
FTSE100	Reuters	•				
USD-JPY exchange rate	Refinitiv (LSEG)		•			
2y gov. US bond yields	Bloomberg		•			
2y gov. JP bond yields	Reuters		•			
10y gov. US bond yields	Bloomberg		•			
10y gov. JP bond yields	Reuters		•			
3m gov. US bond yields	Reuters		•			
3m gov. JP bond yields	Refinitiv (LSEG)		•			
S&P500	S&P		•			
Nikkei225	Reuters		•			
EUR-GBP exchange rate	Refinitiv (LSEG)			•		
2y gov. DE bond yields	Bloomberg			•		
2y gov. UK bond yields	Bloomberg			•		
10y gov. DE bond yields	Bloomberg			•		
10y gov. UK bond yields	Bloomberg			•		
3m euro area OIS rate	Reuters			•		
3m UK bond yields	Refinitiv (LSEG)			•		
EuroStoxx50	Reuters			•		
FTSE100	Reuters			•		
EUR-JPY exchange rate	Refinitiv (LSEG)				•	
2y gov. DE bond yields	Bloomberg				•	
2y gov. JP bond yields	Reuters				•	
10y gov. DE bond yields	Bloomberg				•	
10y gov. JP bond yields	Reuters				•	
3m euro area OIS rate	Reuters				•	
3m JP bond yields	Refinitiv (LSEG)				•	
EuroStoxx50	Reuters				•	
Nikkei225	Reuters				•	
GBP-JPY exchange rate	Refinitiv (LSEG)					•
2y gov. UK bond yields	Bloomberg					•
2y gov. JP bond yields	Reuters					•
10y gov. UK bond yields	Bloomberg					•
10y gov. JP bond yields	Reuters					•
3m UK bond yields	Refinitiv (LSEG)					•
3m JP bond yields	Refinitiv (LSEG)					•
FTSE100	Reuters					•
Nikkei225	Reuters					•
Figures: IRFs		10	10	10	10	10
Figures: FEVD		11	11	11	11	11

Note. The table lists the variables included in each VAR model in SECTION 5, alongside their sources.

B VAR Model

B.1 Estimation

As is common in the structural VAR literature, I employ Bayesian techniques in order to impose more structure on the estimation of the reduced-form coefficients. I use standard Minnesota priors (as in Litterman, 1986) that are cast in the form of a Normal-Inverse-Wishart prior, which conveniently is the conjugate prior for the likelihood of a VAR with Gaussian innovations (see Miranda-Agrippino and Ricco, 2018).

$$\mathbf{\Sigma} \sim \mathcal{IW}(\underline{\mathbf{S}}, \underline{\nu}), \tag{A.1}$$

$$\beta | \Sigma \sim \mathcal{N}(\underline{\beta}, \Sigma \otimes \underline{\mathbf{\Omega}}). \tag{A.2}$$

$\beta = vec([\mathbf{c}, \mathbf{B}_1, ..., \mathbf{B}_p]')$ are the stacked coefficient matrices and $\underline{\mathbf{S}}$, $\underline{\nu}$, $\underline{\beta}$ and $\underline{\mathbf{\Omega}}$ are hyperparameters. Specifically, $\underline{\mathbf{S}}$ and $\underline{\nu}$ are, respectively, the scale matrix and the degrees of freedom of the prior inverse Wishart distribution. As is standard, I specify $\underline{\mathbf{S}}$ as a diagonal matrix with entries σ_i^2 equal to the residual variance of the regression of each variable onto its own first lag. The degrees of freedom are set to $\underline{\nu} = n + 2$ so as to ensure that the prior variances of the coefficient matrices exist and $\mathbb{E}(\beta) = \underline{\beta}$ and $\mathbb{V}ar(\beta) = \underline{\mathbf{S}} \otimes \underline{\mathbf{\Omega}}$.

The standard "Minnesota"-type prior assumes the coefficient matrices to be independently normally distributed. Specifically, their first two moments are:

$$\mathbb{E}[(\mathbf{B_l})_{\mathbf{i,j}} | \mathbf{\Sigma}] = \begin{cases} \delta_i & i = j, l = 1 \\ 0 & \text{otherwise} \end{cases} \tag{A.3}$$

$$\mathbb{V}ar[(\mathbf{B_l})_{\mathbf{i,j}} | \mathbf{\Sigma}] = \begin{cases} \frac{\lambda^2}{l^2} & i = j, \forall l \\ \frac{\lambda^2}{l^2} \frac{\Sigma_{i,i}}{\sigma_j^2} & i \neq j, \forall l \end{cases} \tag{A.4}$$

where $(B_l)_{i,j}$ is the response of variable i to variable j at lag l. As the VAR is estimated in levels, generally I set $\delta_i = 1$, implying random-walk behavior of the underlying time series.[39] As is common, I formalize the idea that more recent lags of a variable tend to be more informative by specifying l^2 in the variance entries. The hyperparameter λ controls the overall tightness of the Minnesota prior, which is determined optimally in the spirit of hierarchical modelling as in Giannone et al. (2015). The credible sets are then constructed by drawing 3,000 times (and discarding the first 2,000 draws) from the posteriors and for each draw making use of the external instruments approach outlined

[39]As some of the variables could be considered to be a priori stationary – e.g. those that are first-differenced in a robustness check –, I experiment with setting $\delta_i = 0$, as in Banbura et al. (2010), but generally find my results to be hardly affected.

in the main text and described in more detail below.

B.2 Identification

As outlined in the main text, one can partition the shock vectors into those of monetary policy, $\boldsymbol{\epsilon}_t^{p\prime}$, and other shocks, $\boldsymbol{\epsilon}_t^{q\prime}$, with corresponding residual vectors $\mathbf{u}_t = [\mathbf{u}_t^{p\prime}, \mathbf{u}_t^{q\prime}]'$, instruments $\mathbf{z}_t$ need to be *relevant* and *exogenous*:

$$\mathbb{E}(\mathbf{z}_t \boldsymbol{\epsilon}_t^{p\prime}) = \boldsymbol{\Phi} \neq \mathbf{0}, \tag{A.5}$$

$$\mathbb{E}(\mathbf{z}_t \boldsymbol{\epsilon}_t^{q\prime}) = \mathbf{0}. \tag{A.6}$$

One relative monetary policy shock. In the baseline approach, only one (relative) monetary policy shock is identified, such that Φ and z_t are scalars. Although Φ is unknown, structural objects as relevant entries of $\mathbf{A_0^{-1}} \equiv \boldsymbol{\mathcal{A}}$ can still be derived as follows:

$$\begin{aligned}
\mathbb{E}(z_t \mathbf{u}_t) &= \mathbb{E}(z_t[\boldsymbol{\mathcal{A}}\boldsymbol{\epsilon}_t]) \\
&= \mathbb{E}(z_t[\boldsymbol{\mathcal{A}}_1\boldsymbol{\epsilon}_{1t} + \boldsymbol{\mathcal{A}}_2\boldsymbol{\epsilon}_{2t}]) \\
&= \mathbb{E}(z_t[\boldsymbol{\mathcal{A}}_1\boldsymbol{\epsilon}_{1t}]) + \mathbb{E}(z_t[\boldsymbol{\mathcal{A}}_2\boldsymbol{\epsilon}_{2t}]) \\
&= \Phi\boldsymbol{\mathcal{A}}_1
\end{aligned} \tag{A.7}$$

in which the last equality follows from (3) and (4). $\boldsymbol{\mathcal{A}}_1$ is the column in $\boldsymbol{\mathcal{A}}$ that describes the impact response of the endogenous variables to the identified structural shock, whereas $\boldsymbol{\mathcal{A}}_2$ are the columns for all other, not identified shocks. By partitioning $\boldsymbol{\mathcal{A}}_1 = [\alpha_{11}, \boldsymbol{\mathcal{A}}_{21}']'$ and rearranging, the unknown Φ can be eliminated:

$$\boldsymbol{\mathcal{A}}_{21} = \frac{\mathbb{E}(z_t \mathbf{u}_{2t})}{\mathbb{E}(z_t u_{1t})} \mathcal{A}_{11} \tag{A.8}$$

The structural coefficients $\boldsymbol{\mathcal{A}}_{21}$ can hence be derived as the ratio of the coefficients of a regression of the residuals on the instrument, scaled by α_{11}. For all impulse responses in the paper, α_{11} is the impact response of the exchange rate, normalized to one percent.

Two monetary policy shocks. In the alternative case, two monetary policy shocks are identified jointly, using two instruments, $\mathbf{z}_t^{p\prime} = [z_t^{us}, z_t^{ea}]'$. In this case, $\boldsymbol{\Phi}$ is a (2×2) matrix and we have:

$$\boldsymbol{\Phi} = \mathbb{E}(\mathbf{z}_t \boldsymbol{\epsilon}_t^{p\prime}) = \begin{pmatrix} \mathbb{E}(z_t^{us}\epsilon_t^{us}) & \mathbb{E}(z_t^{us}\epsilon_t^{ea}) \\ \mathbb{E}(z_t^{ea}\epsilon_t^{us}) & \mathbb{E}(z_t^{ea}\epsilon_t^{ea}) \end{pmatrix} \tag{A.9}$$

In contrast to the single-shock case above, $\mathbf{\Phi}$ is then not canceled out. Identification of each shock in $\tilde{\varepsilon}_t$ cannot be achieved using one instrument at the time in the general case in which the off-diagonal entries in (A.9) are not zero.

Instead, identification is achieved as follows.[40] Partitioning the matrices $\boldsymbol{\mathcal{A}}$ and $\mathbf{\Sigma}$ as

$$\boldsymbol{\mathcal{A}} = \begin{bmatrix} \boldsymbol{\mathcal{A}}_{11} & \boldsymbol{\mathcal{A}}_{12} \\ \boldsymbol{\mathcal{A}}_{21} & \boldsymbol{\mathcal{A}}_{22} \end{bmatrix}; \quad \mathbf{\Sigma} = \begin{bmatrix} \mathbf{\Sigma}_{11} & \mathbf{\Sigma}_{12} \\ \mathbf{\Sigma}_{21} & \mathbf{\Sigma}_{22} \end{bmatrix}, \tag{A.10}$$

$\boldsymbol{\mathcal{A}}_{11}$ and $\mathbf{\Sigma}_{11}$ have dimension 2×2, $\boldsymbol{\mathcal{A}}_{21}$ and $\mathbf{\Sigma_{21}}$ has dimension $(n-2) \times 2$, $\boldsymbol{\mathcal{A}}_{12}$ and $\mathbf{\Sigma}_{12}$ have dimension $2 \times (n-2)$ and $\boldsymbol{\mathcal{A}}_{22}$ and $\mathbf{\Sigma_{22}}$ have dimension $(n-2) \times (n-2)$.

Equivalently to the single-shock case above, the goal is to estimate:

$$\boldsymbol{\mathcal{A}}_1 = \begin{bmatrix} \boldsymbol{\mathcal{A}}_{11} \\ \boldsymbol{\mathcal{A}}_{21} \end{bmatrix}. \tag{A.11}$$

In order to get there, first estimate from the data $\hat{\mathbf{\Lambda}} = \hat{\boldsymbol{E}}_{21}\hat{\boldsymbol{E}}_{11}^{-1}$ given $\hat{\boldsymbol{E}} = \hat{\mathbb{E}}(\hat{\boldsymbol{u}}_t \boldsymbol{z}_t')$ of dimension $n \times 2$ and $\hat{\boldsymbol{E}}_{11}$ $\hat{\boldsymbol{E}}_{21}$ its partition into the 2×2 upper component and the remaining $(n-2)\times 2$ lower component. $\hat{\mathbf{\Lambda}}$ is then a consistent estimator for the $(n-2)\times 2$ matrix $\mathbf{\Lambda} = \boldsymbol{\mathcal{A}}_{21}\boldsymbol{\mathcal{A}}_{11}^{-1}$ as $\boldsymbol{E}_{11} = \boldsymbol{\mathcal{A}}_{11}\mathbf{\Lambda}$ and $\boldsymbol{E}_{21} = \boldsymbol{\mathcal{A}}_{21}$, such that $\boldsymbol{E}_{21}\boldsymbol{E}_{11}^{-1} = \boldsymbol{\mathcal{A}}_{21}\boldsymbol{\mathcal{A}}_{11}^{-1}$. This procedure hence allows to estimate $\boldsymbol{\mathcal{A}}_1$ as $\hat{\boldsymbol{\mathcal{A}}}_1 = \begin{bmatrix} \hat{\boldsymbol{\mathcal{A}}}_{11} \\ \hat{\mathbf{\Lambda}}\hat{\boldsymbol{\mathcal{A}}}_{11} \end{bmatrix}$.

The next step is to arrive at an estimate of $\boldsymbol{\mathcal{A}}_{11}$, which itself is unknown. However, from the covariance restrictions it follows that $\boldsymbol{\mathcal{A}}_{11}\boldsymbol{\mathcal{A}}_{11}' = \mathbf{\Sigma}_{11} - \boldsymbol{\mathcal{A}}_{12}\boldsymbol{\mathcal{A}}_{12}'$. One can show that

$$\widehat{\boldsymbol{\mathcal{A}}_{12}\boldsymbol{\mathcal{A}}_{12}'} = (\hat{\mathbf{\Sigma}}_{21} - \hat{\mathbf{\Lambda}}\hat{\mathbf{\Sigma}}_{11})'\hat{\mathbf{\Lambda}}^{-1}(\hat{\mathbf{\Sigma}}_{21} - \hat{\mathbf{\Lambda}}\hat{\mathbf{\Sigma}}_{11}); \tag{A.12}$$

$$\hat{\mathbf{\Gamma}} = \hat{\mathbf{\Sigma}}_{22} + \hat{\mathbf{\Lambda}}\hat{\mathbf{\Sigma}}_{11}\hat{\mathbf{\Lambda}}' - \hat{\mathbf{\Lambda}}_{21}\hat{\mathbf{\Lambda}}' - \hat{\mathbf{\Lambda}}\hat{\mathbf{\Sigma}}_{21}'. \tag{A.13}$$

Having obtained these estimates, define $\hat{\boldsymbol{\mathcal{A}}}_{\mathbf{11}}^c$ as the lower Cholesky decomposition of $\hat{\mathbf{\Sigma}}_{11} - \hat{\boldsymbol{\mathcal{A}}}_{12}\hat{\boldsymbol{\mathcal{A}}}_{12}'$. One can then express $\hat{\boldsymbol{\mathcal{A}}}_{11}$ as $\hat{\boldsymbol{\mathcal{A}}}_{11}^c\boldsymbol{Q}$, in which $\boldsymbol{Q}$ is any 2×2 orthogonal matrix. Having thereby obtained an estimate of $\boldsymbol{\mathcal{A}}_{11}$, the remaining part of $\hat{\boldsymbol{\mathcal{A}}}_1$ can then be calculated simply as $\hat{\boldsymbol{\mathcal{A}}}_{21} = \hat{\mathbf{\Lambda}}\hat{\boldsymbol{\mathcal{A}}}_{11}$.

Finally, one can use the relation $\mathbf{u}_t = \mathbf{A}_0^{-1}\boldsymbol{\epsilon}_t$ in order to compute the resulting structural shocks. For that one needs an estimate for the first 2 columns of $\mathbf{A}_0 = \mathcal{A}^{-1}$, call them $\mathbf{A}_{0,1}$, such that $\boldsymbol{\epsilon}_t = \mathbf{A}_{0,1}\mathbf{u}_t$. This can be obtained as

$$\hat{\mathbf{A}}_{0,1} = (\hat{\boldsymbol{\mathcal{A}}}_{11} - \boldsymbol{\mathcal{A}}_{12}\widehat{\boldsymbol{\mathcal{A}}_{21}^{-1}}\boldsymbol{\mathcal{A}}_{21})'[\boldsymbol{I}_k \quad - \widehat{\boldsymbol{\mathcal{A}}_{12}\boldsymbol{\mathcal{A}}_{21}^{-1}}]. \tag{A.14}$$

[40] The following derivations are based on a slightly corrected version of those in the online appendix of Piffer and Podstawski (2018).

with

$$\widehat{\boldsymbol{\mathcal{A}}_{12}\boldsymbol{\mathcal{A}}'_{21}}^{-1} = (\hat{\boldsymbol{\Sigma}}_{21} - \hat{\boldsymbol{\mathcal{A}}}_{21}\hat{\boldsymbol{\mathcal{A}}}'_{11})(\hat{\boldsymbol{\Sigma}}_{22} - \hat{\boldsymbol{\mathcal{A}}}_{21}\hat{\boldsymbol{\mathcal{A}}}'_{21})^{-1}. \tag{A.15}$$

Separate identification of the two shocks is then achieved by drawing many Q matrices and for each check the resulting correlation of the structural shocks with their respective instruments. Specifically, in order to accept a draw I require

$$corr(z_t^{us}, \hat{\epsilon}_t^{us}) > 0,$$

$$corr(z_t^{ea}, \hat{\epsilon}_t^{ea}) > 0,$$

$$c\bar{o}rr(z_t^{us}, \hat{\epsilon}_t^{us}) > c\bar{o}rr(z_t^{ea}, \hat{\epsilon}_t^{us})\psi,$$

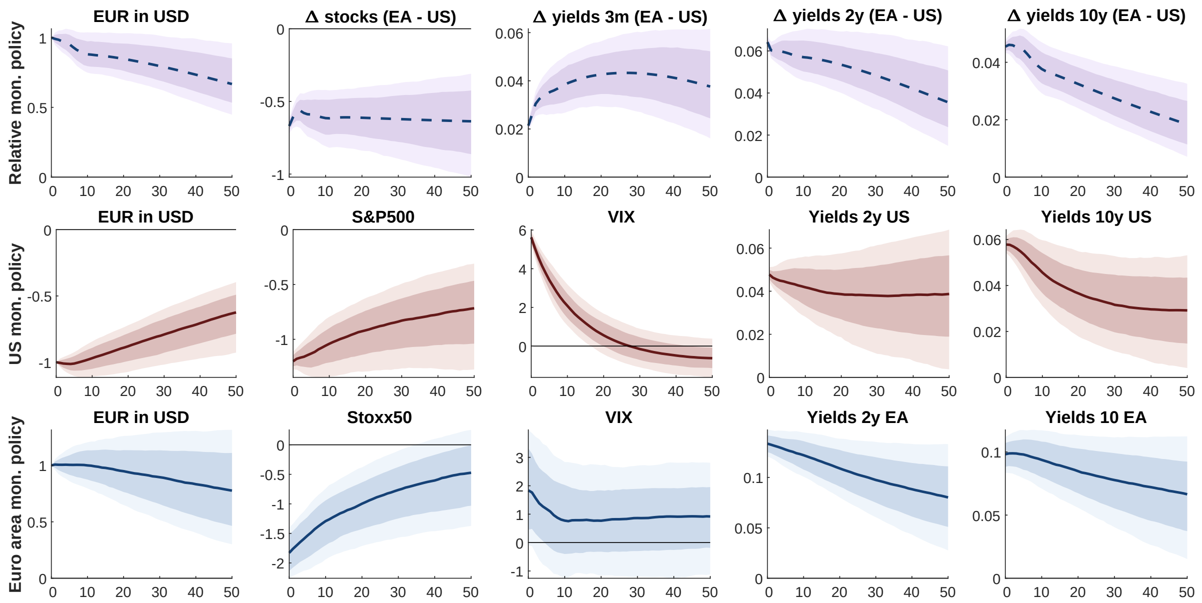
$$c\bar{o}rr(z_t^{ea}, \hat{\epsilon}_t^{ea}) > c\bar{o}rr(z_t^{us}, \hat{\epsilon}_t^{ea})\psi,$$

where $c\bar{o}rr$ is the absolute value of the correlation coefficient and ψ is set to 3, ensuring a sharp differentiation between the shocks.[41] Implementing this procedure into the Bayesian estimation algorithm, I draw Q 100 times for each reduced-form value of β and Σ and choose the resulting median target draw.

[41]Specifying $\psi = 1$ would correspond to Bayesian proxy VAR implementation in Arias et al. (2021), whereas $\psi >> 1$ can be likened to the approach used in Piffer and Podstawski (2018) and generally results in much tighter confidence bands around the relative FEVD shares of the two shocks. I verify that the resulting shocks are correlated with their instruments in the order of 0.25, which would correspond to the high relevance threshold in Arias et al. (2021).

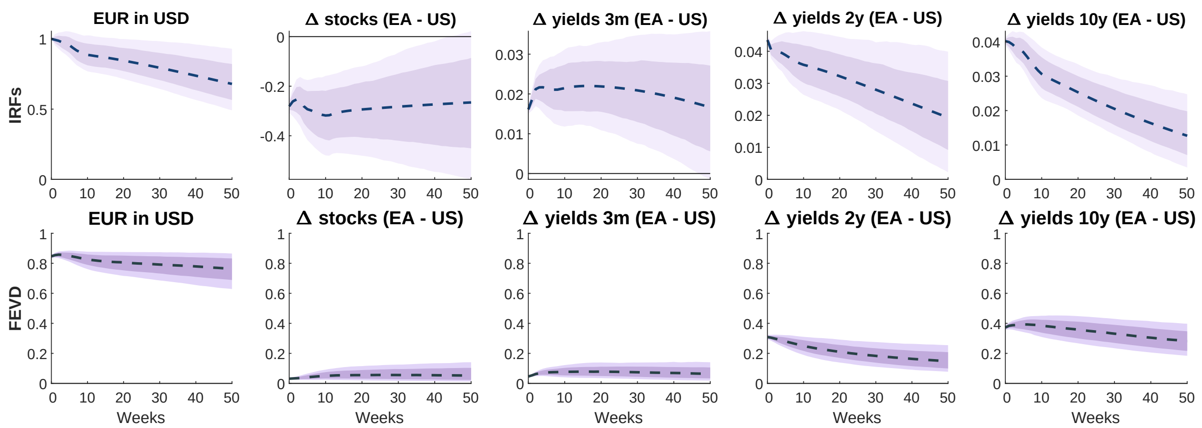
C Additional Results

Figure C.1: IRFs to monetary policy shocks identified with interest rate instrument (weekly data)



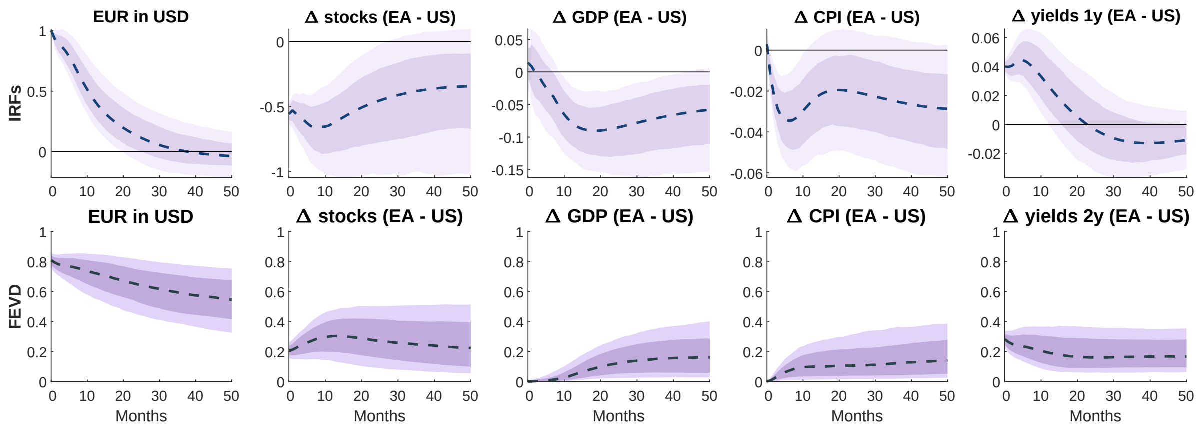
Note. Impulse responses to a relative monetary policy shock in the relative model (top row, purple), and US (middle, red) and euro area (bottom, blue) contractionary monetary policy shock in the levels model, normalized to change the euro's value vs. the US dollar by 1 percent. Shocks identified with interest rate instead of exchange rate instrument. Remaining details as in Figure 4.

Figure C.2: IRFs and FEVD without instrument cleansing (weekly data)



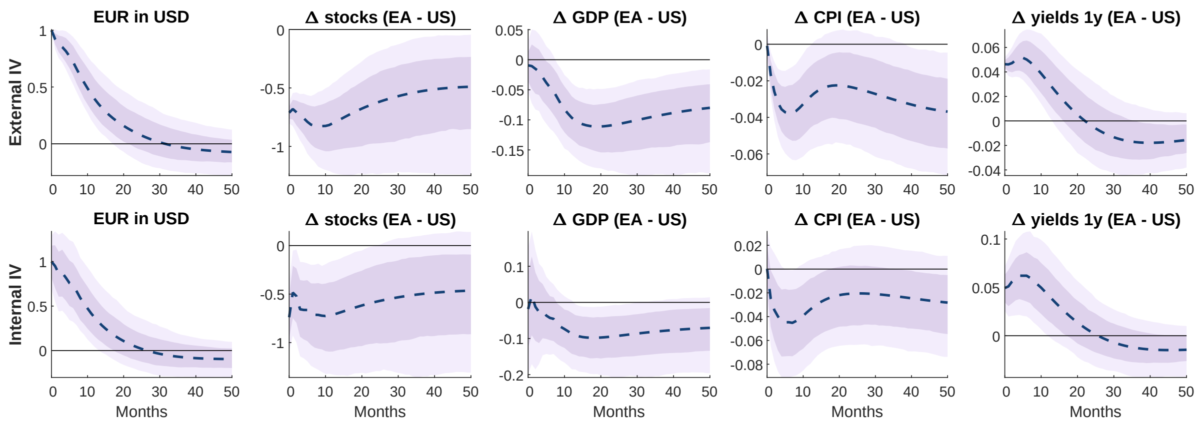
Note. Impulse responses to (top) and share of forecast error variance explained by the relative monetary policy shock (bottom) in weekly relative model. Identification from instrument without cleansing from information effects. Remaining details as in Figures 2 and 3.

Figure C.3: IRFs and FEVD without instrument cleansing (monthly data)



Note. Impulse responses to (top) and share of forecast error variance explained by the relative monetary policy shock (bottom) in monthly relative model. Identification from exchange rate instrument without cleansing from information effects. Remaining details as in Figure 6.

Figure C.4: IRFs, comparison of external with internal IV



Note. Impulse responses to relative monetary policy shock in the monthly relative model, estimated with external IV (= proxy VAR, top) and internal IV (bottom). Remaining details as in Figure 6.