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Hawks vs. Doves: ECB's Monetary Policy in Light of the Fed's Policy Stance

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Non-technical summary

In response to the recent surge in inflation, most major central banks have started tightening their policy stance. However, as macroeconomic fundamentals, financial dynamics or cyclical positions may differ across countries, monetary policies may also diverge internationally. Some central banks — such as the Bank of England or the Federal Reserve (Fed) — started their tightening cycle relatively early, while other central banks — most notably the European Central Bank (ECB) — have been slower to raise their policy rates. This may lead to divergence in monetary policy, opening up potential interaction effects between central banks.

These interaction effects arise because monetary policy often co-moves internationally. However, most models assume that a central bank's reaction function is driven purely by domestic factors, a rather strong assumption given the increased financial and real linkages in the world economy. Even if domestic quantities are somewhat affected by foreign activity through the trade balance, foreign monetary policy is a neglected dimension in these analyses. Therefore, our goal in this paper is to model the ECB's monetary policy as being explicitly dependent on the Fed's policy stance, given the prominent role of the latter in the global financial cycle (Miranda-Agrippino and Rey, 2020).

In this paper, we fill this gap by developing a nonlinear vector autoregression (VAR) model for the EA economy. Our model assumes a nonlinear interaction between its dynamic coefficients and a measure of the US monetary policy stance. In particular, we propose a smooth transition VAR (ST-VAR) that takes the Fed's policy stance as the signal variable determining the regime. We measure the Fed's monetary policy stance by considering the deviation of the effective federal funds rate from the US equilibrium real interest rate (Laubach and Williams, 2003), representing the rate at which an economy operates at its potential output while maintaining stable inflation. Consequently, the deviation from this rate captures the actual stance of monetary policy. ECB monetary policy shocks are identified through high-frequency reactions of financial markets in a narrow window around scheduled events associated with monetary policy announcements (Altavilla et al., 2019). This enables us to flexibly track how the effects of ECB policy surprises change based on the Fed's stance.

Our analysis looks at the financial side of the economy and includes weekly data on EA bond yields of different maturities and inflation-linked swaps to proxy marked-based inflation expectations. Since the number of different maturities is large, we rely on a term structure model in the Nelson-Siegel tradition (Diebold and Li, 2006). This allows us to capture the dynamic responses of real interest rates over a wide range of maturities. Compared to existing studies, we provide a comprehensive analysis that includes not only the full term structure but also explicit conditions on the Fed's stance in a high-frequency setting.

Our findings indicate that the Fed had a rather restrictive stance until the onset of the financial crisis and after 2016. The period in between, however, was characterized by expansionary periods punctuated by several restrictive and transitory policy regimes. With respect to the contractionary monetary surprises in the EA, our findings suggest that the dynamic responses of EA financial markets strongly depend on the respective Fed stance, providing appreciable evidence of nonlinear interaction effects. Moreover, the relationship between impulse responses and US monetary policy is heterogeneous with respect to shape and magnitude over the term structure. At the short end of the curve, government bond yields increase, with a consecutive decrease only in the case of an expansionary Fed stance. The longer end of the curve shows declining yields after one week when the Fed is hawkish. Inflation-linked swaps initially fall, but then rise after one week when the Fed is expansionary. However, the differences in responses between expansionary and contractionary US policy phases are not significant over the entire term structure.

One key advantage of our approach is that it allows us to examine how the term structure of *real* rates changes in response to EA monetary policy shocks conditional on the prevailing US monetary policy regime. This analysis reveals that restrictive ECB surprises trigger stronger positive reactions if the Fed is hawkish as opposed to a situation where monetary policy diverges.

Hawks vs. Doves: ECB's Monetary Policy in Light of the Fed's Policy Stance

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Abstract

The secular increase in globalization led to a substantial increase in the interconnectedness of global financial markets. This has important implications for the conduct of monetary policy, as central bank policies may diverge across countries, potentially affecting key transmission channels of domestic policy actions. In this paper, we use a non-linear multivariate time series model to shed light on how the US monetary policy stance affects the conduct of monetary policy in the euro area. We assume that the dynamic coefficients implicitly depend on a measure of the Federal Reserve's policy stance through a smooth transition function. This assumption allows us to examine how the dynamic responses of financial market quantities such as government bond yields and inflation swaps to euro area monetary policy shocks change with the US policy stance. Scenario-specific impulse responses show that the transmission of euro area monetary policy through financial markets does indeed depend on the prevailing monetary policy regime of the Federal Reserve and has significant effects on a variety of euro area variables.

Keywords: Monetary Policy Transmission, Financial Markets, Real Rates, High-Frequency Data, Smooth Transition VAR

JEL Codes: E43, E52, F42, G12

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1 Introduction

In response to the recent surge in inflation, most major central banks have started tightening their policy stance. However, as macroeconomic fundamentals, financial dynamics or cyclical positions may differ across countries, monetary policies may also diverge internationally. Some central banks — such as the Bank of England or the Federal Reserve (Fed) — started their tightening cycle relatively early, while other central banks — most notably the European Central Bank (ECB) — have been slower to raise their policy rates. This may lead to divergence in monetary policy, opening up potential interaction effects between central banks.

These interaction effects arise because monetary policy often co-moves internationally. However, most models assume that a central bank's reaction function is driven purely by domestic factors (see, e.g., Del Negro *et al.*, 2013, for a recent state-of-the-art dynamic stochastic general equilibrium model that is used for policy analysis), a rather strong assumption given the increased financial and real linkages in the world economy. Even if domestic quantities are somewhat affected by foreign activity through the trade balance, foreign monetary policy is a neglected dimension in these analyses. Therefore, our goal in this paper is to model the ECB's monetary policy as being explicitly dependent on the Fed's policy stance, given the prominent role of the latter in the global financial cycle (Miranda-Agrippino and Rey, 2020).

One way to capture international co-movement in monetary policy (and financial markets more generally) is through the use of large-scale, multi-country models. The literature on empirical global macro models (see, among many others, Dees *et al.*, 2007; Feldkircher and Huber, 2016; Georgiadis, 2017; Koop and Korobilis, 2016; Burriel and Galesi, 2018; Dées and Galesi, 2021) typically uses large-scale models that represent the structure of the economy and are able to capture cross-country relationships. These models often assume that countries are dynamically linked and that shocks originating in one country have immediate and lagged effects not only on itself but also on other countries through spillovers. A key limitation, however, is that these papers implicitly assume monetary policy reaction functions that imply an endogenous reaction of foreign central banks to domestic monetary policy shocks. However, by definition, these models do not allow for an analysis of whether *domestic* transmission channels to monetary shocks change when the foreign central bank changes its behavior. This question, however, is of paramount importance if policymakers want to adequately assess the effects of their own policies in a globalized world. Hence, the main focus of this paper is to empirically analyze how domestic monetary policy effects change in the euro area (EA) *conditional* on a potentially divergent foreign policy stance of the globally most important central bank — the Fed.

More generally, several recent papers have focused on how monetary policy effects are attenuated by various influences. One strand of the literature addresses this question by looking at economic and policy spillovers to other economies. These empirical papers focus mainly on US monetary policy spillovers and spillbacks (Kim, 2001; Canova, 2005; Bluwstein and Canova, 2016; Crespo Cuaresma *et al.*, 2019; Ca'Zorzi *et al.*, 2020; Breitenlechner *et al.*, 2022; Ca'Zorzi *et al.*, 2023; Georgiadis and Jarocinski, 2023) and find strong effects on other economies.

In general, the empirically identified channels through which foreign monetary policy may affect domestic macroeconomic quantities consider both real and financial markets. In this paper, we focus on the channel operating through financial markets. Countries are exposed to changing financial conditions, exchange rate dynamics, and collateral valuation effects with potential implications for bank balance sheets. Here, the US as a supplier of safe assets (Treasuries), its impact on the global financial cycle (Miranda-Agrippino and Rey, 2020), and the special role of the US dollar (Gopinath *et al.*, 2020) are central parts of our analysis. Thus, changes in the demand for US assets triggered by US monetary policy actions can change the demand for EA assets. As a consequence, policy actions of other central banks, especially if not aligned with the Fed's, may affect financial variables with potentially offsetting effects. As recently shown by Ca'Zorzi *et al.* (2023), while US monetary policy shocks have a noticeable impact on several EA macroeconomic and financial indicators, the spillover effects of EA monetary policy on the US economy are rather limited, if not absent. Interestingly, the study also found that the effects on real economic activity and, in particular, financial markets were much larger. To summarize, most empirical studies have focused on the direct effects of monetary policy surprises, i.e., whether identified exogenous shocks have a direct impact on the domestic economy. However, there is a notable lack of research on whether the effects of domestic monetary policy shocks depend on the policy stance of a foreign bank.

In this paper, we aim to fill this gap by developing a nonlinear vector autoregression (VAR) model for the EA economy. Our model assumes a nonlinear interaction between its dynamic coefficients and a measure of the US monetary policy stance. In particular, we propose a smooth transition VAR (ST-VAR) that takes the Fed's policy stance as the signal variable determining the regime. We measure the Fed's monetary policy stance by considering the deviation of the effective federal funds rate from the US equilibrium real interest rate Laubach and Williams (2003), representing the rate at which an economy operates at its potential output while maintaining stable inflation. Consequently, the deviation from this rate captures the actual stance of monetary policy. ECB monetary policy shocks are identified through high-frequency reactions of financial markets in a narrow window around scheduled events associated with monetary policy announcements Altavilla *et al.* (2019). This enables us to flexibly track how the effects of ECB policy surprises change based on the Fed's stance.

Our analysis looks at the financial side of the economy and includes weekly data on EA bond yields of different maturities and inflation-linked swaps to proxy marked-based inflation expectations. Since the number of different maturities is large, we rely on a term structure model in the Nelson-Siegel tradition (Diebold and Li, 2006). This allows us to capture the dynamic responses of real interest rates over a wide range of maturities. Compared to existing studies, we provide a comprehensive analysis that includes not only the full term structure but also explicit conditions on the Fed's stance in a high-frequency setting. This helps to shed light on how international linkages shape financial market reactions to monetary policy.

The empirical results indicate that the Fed had a rather restrictive stance until the onset of the financial crisis and after 2016. The period in between, however, was characterized by expansionary periods punctuated by several restrictive and transitory policy regimes. With respect to the contractionary monetary surprises in the EA, our findings suggest that the dynamic responses of EA financial markets strongly depend on the respective Fed stance, providing appreciable evidence of nonlinear interaction effects. Moreover, the relationship between impulse responses and US monetary policy is heterogeneous with respect to shape and magnitude over the term structure. At the short end of the curve, government bond yields increase, with a consecutive decrease only in the case of an expansionary Fed stance. The longer end of the curve shows declining yields after one week when the Fed is hawkish. Inflation-linked swaps initially fall, but then rise after one week when the Fed is expansionary. However, the differences in responses between expansionary and contractionary US policy phases are not significant over the entire term structure.

One key advantage of our approach is that it allows us to examine how the term structure of *real* rates changes in response to EA monetary policy shocks conditional on the prevailing US monetary policy regime. This analysis reveals that restrictive ECB surprises trigger stronger positive reactions if the Fed is hawkish as opposed to a situation where monetary policy diverges.

The rest of the paper is organized as follows. In the next section, we introduce our econometric model, while in Section 3 we discuss the dataset, perform univariate analyses to set the stage, and show the model-implied US monetary policy regimes. Section 4 presents the main results based on our ST-VAR as well as a robustness exercise. The last section summarizes and concludes the paper.

2 Econometric Framework

Our objective is to investigate whether the effects of the ECB's monetary policy depend on the Federal Reserve's policy stance. In order to account for higher frequency dynamics in the financial sector and potential effects arising from inflation expectations, we use a multivariate time series model of weekly data on government bond yields, inflation-linked swaps, and a set of macro-financial variables. In this section, we explain the term structure model used to obtain the yield and inflation swap curve factors. We then introduce our smooth transition VAR model.

2.1 Nelson-Siegel term-structure model

Our main aim is to obtain a complete picture of how the transmission of monetary policy changes, especially across different maturities of yields and inflation swaps. This implies that the number of time series to be modeled is potentially very large, and overfitting problems arise. As a remedy, we propose a parsimonious term structure model in the Nelson-Siegel (NS) tradition (Nelson and Siegel, 1987; Diebold and Li, 2006; Borağan Aruoba, 2020) for both bond yields and inflation swaps.

Let $j \in \{r, \pi\}$, where $r_t(\tau)$ denotes the weekly EA government bond yield and $\pi_t(\tau)$ the weekly inflation-linked swap, both with maturity τ. Assuming that both the EA yield curve and the respective inflation swap curve feature a factor structure, yields

$$j_t(\tau) = L_{jt} + S_{jt}\left(\frac{1-\exp\left(-\zeta_j\tau\right)}{\zeta_j\tau}\right) + C_{jt}\left(\frac{1-\exp\left(-\zeta_j\tau\right)}{\zeta_j\tau} - \exp\left(-\zeta_j\tau\right)\right), \tag{2.1}$$

with L_{jt}, S_{jt} and C_{jt} denoting a level, slope and curvature factor, which summarizes movements along the yield curve. The corresponding factor loadings depend on a parameter ζ_j, implying an exponential decay rate. The factors feature a convenient economic meaning. While the level factor defines the behavior at the long end of the yield curve (long-run level of interest rates, maturities greater than 15 years), the slope factor determines the relationship between short and long-run interest rates. Finally, the curvature factor defines the behavior in the medium segment of the curve and can be interpreted as the relationship of bond prices to changes in yields.

According to Borağan Aruoba (2020) the resulting three factors for the inflation-linked swap curve have a similar convenient interpretation. The level factor, $L_{\pi t}$ simply captures the long-term inflation expectations, while the slope factor, $S_{\pi t}$ denotes the difference between long- and short-term expectations. Finally, the curvature factor, $C_{\pi t}$ is a measure of medium-term expectations relative to short- and long-term expectations.

Following Diebold and Li (2006), we set $\zeta_r = \zeta_\pi = 0.0609 \times 12$, rendering the loadings as known and deterministic. This feature enables us to obtain the six latent factors in $\boldsymbol{f}_t = (L_{rt}, S_{rt}, C_{rt}, L_{\pi t}, S_{\pi t}, C_{\pi t})'$ with simple period-specific cross-sectional Ordinary Least Square (OLS) regressions on the yield and the inflation swap data, respectively.

2.2 A nonlinear model of monetary policy spillovers

To model the EA economy in a nonlinear fashion, we assume that $\boldsymbol{f}_t$, and additional macro-financial quantities are stored in an $M \times 1$ vector $\boldsymbol{y}_t$ and evolve according to a smooth transition VAR model (Teräsvirta and Anderson, 1992; Weise, 1999; Gefang and Strachan, 2009; Hubrich and Teräsvirta, 2013):

$$\begin{aligned}\boldsymbol{y}_t = &\left(\boldsymbol{c}_1 + \boldsymbol{\gamma}_1 z_t + \sum_{p=1}^{P} \boldsymbol{A}_{1p} \boldsymbol{y}_{t-p}\right) \times S_t + \\ &\left(\boldsymbol{c}_0 + \boldsymbol{\gamma}_0 z_t + \sum_{p=1}^{P} \boldsymbol{A}_{0p} \boldsymbol{y}_{t-p}\right) \times (1 - S_t) + \boldsymbol{\varepsilon}_t,\end{aligned} \tag{2.2}$$

with $\boldsymbol{c}_k$ denoting a regime specific intercept (for $k \in \{0, 1\}$), $\boldsymbol{\gamma}_k$ is an M-dimensional vector measuring the impact of a monetary policy surprise z_t in the domestic economy, $\boldsymbol{A}_{ip}$ are $M \times M$-dimensional coefficient matrices with lag $p = 1, ..., P$ while $\boldsymbol{\varepsilon}_t$ denotes an M-dimensional vector of Gaussian shocks with zero mean and $M \times M$-dimensional variance-covariance matrix $\boldsymbol{\Sigma}$. In a nutshell, the proposed model is a combination of two linear VARs, where each regime is associated with either an expansionary or a contractionary Fed policy stance.

The scalar S_t, which governs the transition between the two regimes, is defined by a nonlinear function g that is bounded between zero and unity and includes the monetary policy stance abroad as an input variable, $S_t = g(z_t^*)$. We assume that this transition process is smooth and defined by a first-order logistic function:

$$g(z_t^*) = \frac{1}{1 + \exp\{-\rho(z_t^* - \mu)\}}, \tag{2.3}$$

where the adjustment process depends on a latent threshold parameter μ and a non-negative speed of adjustment parameter ρ. Note that the input variable z_t^* enters the specification in Eq. (2.3) contemporaneously, which supports the notion that the Fed responds only to domestic factors. Thus, the Fed's monetary policy stance can be considered exogenous to the ECB's monetary policy, following the findings of Ca'Zorzi *et al.* (2020) and the literature on the global role of US monetary policy (Miranda-Agrippino and Rey, 2020).

This model is quite flexible and allows for a nonlinear interaction between domestic dynamics in $\boldsymbol{y}_t$ and the foreign monetary policy stance in z_t^*. We standardize z_t^* to feature zero mean and unit variance. Our choice of the switching function is motivated by its broad empirical support (see, e.g., Teräsvirta, 1994) but also by the fact that it nests models with constant parameters as well as models with abrupt regime shifts. Note that as the parameter ρ becomes large, the function g approaches a heavy-side function and the model reduces to a threshold specification that jumps between regimes when z_t^* exceeds a threshold μ. If ρ is small, it implies a smooth and slowly varying adjustment process across regimes. Thus, not only does it allow us to be agnostic a priori about the exact specification (i.e., threshold vs. smooth transition dynamics), but it also provides a convenient economic interpretation. Since the function is monotonically increasing in z_t^*, the two regimes correspond to loose and tight US monetary policy, respectively. In contrast to models with a clear threshold, where each regime has distinct dynamics, our model is a mixture of the two regimes. More precisely, we model the dynamics of the euro area economy conditional on a mixture of tight and loose US monetary policies with weights depending on z_t^*.

Since our model is heavily parameterized and has significant nonlinearities, we opt for a Bayesian approach to estimation. Therefore, we specify appropriate priors for all parameters of the model based on extensive support from the literature on nonlinear time series models. For the sake of brevity, we simply state the corresponding priors verbally and refer to the technical appendix Section D.1 for more details. For the VAR coefficients, we use a Horseshoe prior (Carvalho *et al.*, 2010), which shrinks irrelevant coefficients towards zero and thus significantly reduces the estimation uncertainty. The prior specification of the variance-covariance matrix follows an inverted Wishart distribution and is therefore standard. Moreover, the prior choice of the parameters governing the state transition ensures a sensible state allocation, with both regimes clearly separated. That is, the threshold parameter follows a weakly informative uniform distribution, while we impose a Gamma prior on the speed of adjustment parameter. The hyperparameters for the latter are set so that the transition between the two states is relatively fast. Applying Bayes' theorem and combining the prior with the model's likelihood function yields the posterior distributions. Since the full posterior distribution is of non-standard form, we use a Markov Chain Monte Carlo (MCMC) algorithm, described in detail in Section D.2, to simulate from it.

3 Data and Model Specification

To examine high-frequency financial market reactions, we measure the impact of ECB monetary policy at a weekly frequency. This approach minimizes the noise typically present in daily time series, while still providing a sufficient number of observations for reliable inference beyond standard monthly frequencies. Our time series data span from 2005:W1 to 2019:W52, covering several monetary policy episodes in the US and the euro area. Due to the unprecedented impact of the COVID-19 pandemic, we have limited our sample to the end of 2019. Our ST-VAR model has a lag length of $p = 4$, which captures dynamic dependencies up to one month in the past. We start with the specific composition of the vector $\boldsymbol{y}_t$. For details on variable definitions, sources, and transformations, see Appendix A. We continue with an elaboration of the yield- and inflation-linked swap data and the respective estimated factors, and proceed with a description of the ECB's policy surprise measure and the Fed's policy stance.

3.1 Estimation of the Yield and Inflation Swap Curve

In our empirical work, we focus on macro-financial developments and monetary policy effects in the euro area, thus treating it as our home economy. Given the dominant role of US monetary policy in the global financial cycle, we condition (Miranda-Agrippino and Rey, 2020; Dées and Galesi, 2021) on the monetary policy stance of the Fed. Thus, the vector of observables $\boldsymbol{y}_t = (z_{EA,t}, \boldsymbol{f}_t, Stocks_t, CISS_t, FX_t)$ includes the monetary policy surprise instrument for EA ($z_{EA,t}$), the yield, indexed with r and the inflation swap curve indexed with π factors, $\boldsymbol{f}_t = (L_{rt}, S_{rt}, C_{rt}, L_{\pi t}, S_{\pi t}, C_{\pi t})$, and macro-financial time series for the EA. The latter consists of the EuroStoxx50 ($Stocks_t$), the Composite Indicator of Systemic Stress ($CISS_t$), and the USD-EUR exchange rate (FX_t). Our signal variable (z_t^*) is a measure of the US monetary policy stance.

To construct the Nelson-Siegel factors of the yield curve, our weekly data set includes EA government bond yields (measured by yield curve instantaneous forward rates and labeled as `YLD`) for 15 maturities, with $\tau \in \{1 : 10, 12, 15, 20, 25, 30\}$.[1] More specifically, we use only data compiled from AAA-rated bonds, as we want to avoid yield curve dynamics stemming from idiosyncratic country risks. During the euro area sovereign debt crisis, especially Greek, Italian, and Spanish government bond dynamics were subject to strong sovereign shocks. Hence, an inclusion of all EA-bonds may impede the identification of monetary policy shocks as it becomes harder to distinguish policy from country-specific shocks.

[1] In principle there are more maturities available within this continuum. However, to match the available maturities for the inflation swaps, we omit these. Including all accessible maturities for the government bond yields leaves the results virtually unchanged.

To construct the Nelson-Siegel factors of the yield curve, our weekly dataset includes EA government bond yields (measured by the yield curve's instantaneous forward rates and labeled `YLD`) for 15 maturities, with $\tau \in \{1 : 10, 12, 15, 20, 25, 30\}$.[2] More specifically, we use only data compiled from AAA-rated bonds, as we want to avoid yield curve dynamics driven by idiosyncratic country risks. During the euro area sovereign debt crisis, the dynamics of Greek, Italian, and Spanish government bonds in particular were subject to strong sovereign shocks. Therefore, including all EA bonds may complicate the dynamic analysis by making it more difficult to distinguish policy from country-specific shocks.

For obtaining the Nelson-Siegel factors of the inflation-linked swap curve, we consider EA inflation swaps (labeled as `INFSWP`) for 15 maturities matching those we used to construct the Nelson-Siegel factors above. More precisely, these are zero-coupon inflation swaps linked to the harmonized consumer price index (HVPI) excluding tobacco for the euro area. They depend on realized and expected inflation, and can thus be employed as a hedging tool for future inflation. For our purpose they carry valuable information about marked-based expected future inflation as reported in Borağan Aruoba (2020) and discussed above. It is important to note that inflation and inflation expectations have a significant impact on the real return that an investor can accumulate over time. In particular, future inflation represents a significant upside risk to real returns, against which investors can use ILS to protect against future inflation. Consequently, we use these swaps as a market-based proxy for otherwise latent inflation expectations. However, ILS also contain an *inflation risk premium*, which investors invested in nominal assets demand for the arising inflation risks as well as a liquidity premium. In principle, to obtain the "pure" expectational component out of the swaps, one has to estimate a model and purge these swaps from other factors. However, for our application, we are interested in the composite, i.e., the inflation-linked swap per se, which includes a measure for uncertainty. Moreover, ILS provide better information than other market-based measures (e.g., inflation-indexed Treasury yields), as shown by Haubrich *et al.* (2012), and are similar to professional forecasters' expectations but less similar to households' inflation expectations. Since we are interested in high-frequency dynamics, survey-based expectations, typically at monthly or quarterly frequency, cannot be used. Therefore, we do not opt for an approach to disentangle "pure" expectations from the risk premium in market-based measures of inflation expectations.

[2] In principle, there are more maturities available along this continuum. However, in order to match the available maturities for the inflation swaps, we omit them. Including all available maturities for government bond yields leaves the results virtually unchanged.

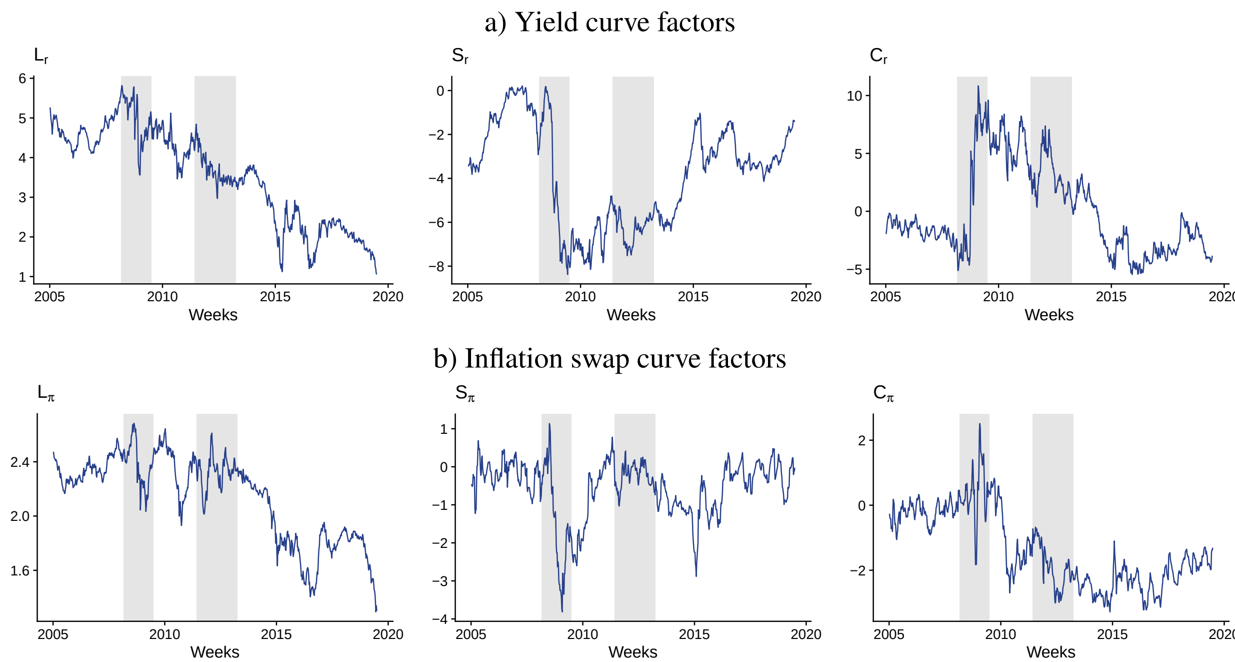
To obtain the Nelson-Siegel factors of the inflation-linked swap curve, we consider EA inflation swaps (labeled INFSWP) for 15 maturities that match those used to construct the Nelson-Siegel factors above. More specifically, they are zero-coupon inflation swaps linked to the Harmonized Index of Consumer Prices (HICP) excluding tobacco for the euro area. They depend on realized and expected inflation and can therefore be used as a hedging instrument against future inflation. For our purposes, they carry valuable information about mark-based expected future inflation as reported in Borağan Aruoba (2020) and discussed above. However, inflation-linked swaps do not necessarily reveal market participants' expectations directly. In general, they also include an *inflation risk premium* that investors in nominal assets demand for the resulting inflation risks, as well as potentially a liquidity premium. In principle, to obtain the "pure" expectation component out of the swaps, one would have to estimate a model and purge these swaps of these other factors. For our application, however, we are interested in the composite, i.e., the inflation-linked swap per se, which includes a measure of uncertainty.

In addition to government bond yields and inflation swaps, we consider the nominal EUR-USD exchange rate (FX, in percentage changes) and the Euro Stoxx 50 index as a stock market indicator (Stocks, in percentage changes). Thus, a depreciation of the exchange rate corresponds to an appreciation of the euro. We also include the EA composite indicator of systemic stress (CISS, Kremer *et al.*, 2012). This index captures 15 indicators of stress in financial markets, and thus should be able to capture financial stability issues associated with monetary policy shocks. Finally, we differenced the data to ensure stationarity.

This dataset covers the financial side of the euro area economy in a high-frequency fashion. A key contribution of our analysis is thus the modeling of the full yield curve, which allows us to trace the effects of monetary policy on different (sub)segments of the EA yield curve. Moreover, since these quantities are nominal, the inclusion of the full inflation swap curve allows us to compute the corresponding real interest rate responses. This setup allows us to gain unique insights into how monetary policy affects not only nominal quantities, but also real quantities and inflation expectations. The inclusion of the CISS and the Eurostoxx index allows us to track movements in financial conditions. Finally, to capture the exchange rate-related channel of monetary policy, we also include the bilateral nominal exchange rate.

Next, we discuss the estimated yield curve and inflation swap curve factors. Panels (a) and (b) in Figure 1 show the respective Nelson-Siegel factors of the yields and inflation swaps in levels. As shown in panel (a), the variation in the curvature factor, C_{rt}, is considerably higher than that of the slope, S_{rt}, and level factors, L_{rt}, which is a common feature of yield curve factors. The negative slope factor throughout most of the sample reflects the typical upward pattern of the EA yield curve, except for short periods prior to the financial crisis, and surprisingly shows a negative correlation with the curvature factor. Towards the end of the sample, the increasing (but still negative) slope factor indicates a flatter curve. In addition, we observe that the level factor fluctuates around a long-run level of about 5% before the financial crisis, but falls sharply in the middle of the crisis, with a steady decrease in the trend of long-term nominal interest rates in subsequent periods and in the aftermath of the Great Recession. This reflects a secular decline in long-term yields, driven in part by unconventional monetary policy in the EA and lower levels of trend inflation. The slope and curvature factors tell a similar story, and the movements can again be linked to monetary policy.

Figure 1: Yield curve and inflation swap factors for the EA.



Notes: The factors (shown in levels) are obtained with a three-factor Nelson Siegel model, where L_{jt} can be interpreted as the level, S_{jt} as the slope, and C_{jt} the curvature factor for j being either the associated yield curve (r) or the inflation swap curve (π) factors. Gray shaded areas indicate EA recessions based on the Euro Area Business Cycle Dating Committee.

Turning to the estimated factors associated with the inflation swap curve, other interesting findings emerge. First, we find a substantial and persistent decline in the level of inflation expectations. Given the level of the yield curve in panel (a) of Fig. 1, this suggests that real interest rates were positive and then approached negative territory toward the end of the sample. Both the slope and the curvature of the swap curve show a sharp decline during the global financial crisis. This implies that short-run inflation expectations declined, while long-run inflation expectations either remained constant or increased. Interestingly, a second sharp decline in the slope of the swap curve is observed during the introduction of the Asset Purchase Program (APP) in mid-2014. This and subsequent purchase programs were designed to support the ECB's policy in the face of the effective lower bound. However, Burban *et al.* (2021) provide a decomposition of the ILS into its expectation and risk premium components. Their estimates suggest a change in the sign of the risk premium around 2013/2014. Therefore, this dynamic could also be reflected in our results. After another significant decline in early 2015 to 2016, possibly caused by another round of asset purchases, the slope fluctuates around zero, implying rather stable inflation expectations over different horizons. To ensure stationarity, the respective factors are differenced in the estimation.

3.2 Measuring Monetary Policy

In order to detect monetary policy interactions between the US and the euro area, we rely on a unique measure of ECB monetary policy shocks, which renders a separate identification obsolete. Moreover, we use an instrument that is purged of various types of shocks and thus carries only information about monetary policy itself. As shown in Jarociński and Karadi (2020), central bank communication conveys information not only about monetary policy but also about the state of the economy. This so-called information effect can cause adverse reactions on financial market variables. Therefore, for our purpose of analyzing the effects of monetary policy, we need to make sure that the instruments used are clearly related to monetary policy surprises.

A convenient measure of EA monetary policy shocks (i.e., our domestic shocks, $z_{EA,t}$) that satisfies the above requirements has been proposed in Altavilla *et al.* (2019). This measure is based on the Euro Area Monetary Policy Event-Study Database and exploits the high-frequency reactions of financial markets in a narrow window (ten minutes before and after the event) around identified events associated with monetary policy announcements.[3] The extracted factors consist of the target, the timing, the forward guidance, and the quantitative easing factor, respectively. Each of these factors has an impact on different segments of the yield curve. By taking the sum of these factors, we obtain a broad summary measure of monetary policy surprises. To avoid a contamination of this unified measure with information effects, we use the approach introduced in Jarociński and Karadi (2020), called "poor-man's sign restrictions". This is achieved by restricting the stock market reaction to policy shocks to be negative in sign. This restriction ensures that we have a measure that

[3] The full Event-Study Database can be found via https://www.ecb.europa.eu/pub/pdf/annex/Dataset_EA-MPD.xlsx and a detailed description of the methodology and data in Altavilla *et al.* (2019).

only captures monetary policy surprises. Moreover, as mentioned in Altavilla *et al.* (2019), Appendix A, we omit three specific outliers (August 31, 2006, October 8, and November 6, 2008) where no press conference was held after the Governing Council meeting.

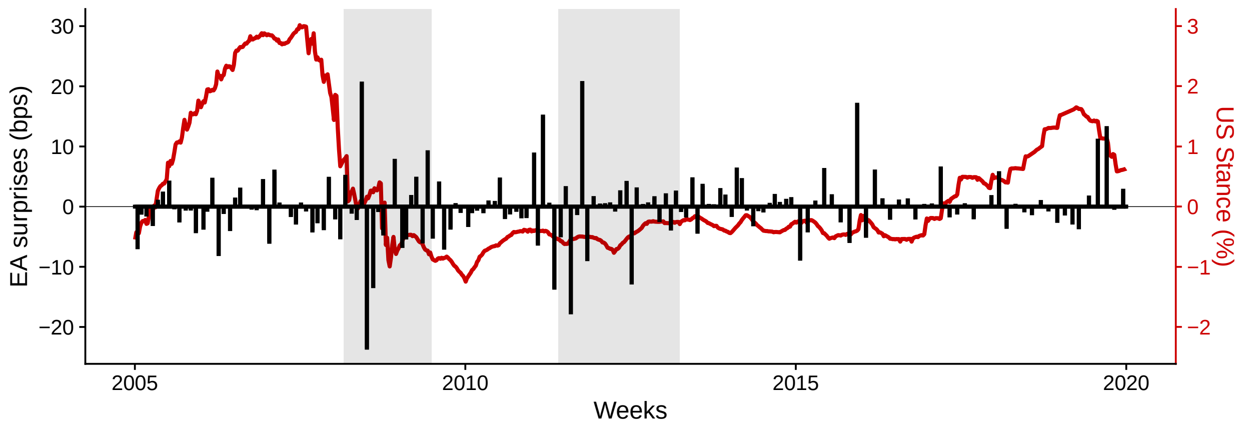
To measure the corresponding US monetary policy stance, denoted z_t^*, we use the effective federal funds rate, DFF_t, and subtract the (one-sided) natural rate estimate, r^*, obtained using the model proposed in Laubach and Williams (2003). Since the latter estimate is only available at the quarterly frequency, we use a linear interpolation to adjust it to the weekly frequency. The effective rate thus serves as a fast-moving component of the stance, while the natural rate captures slower trends. This measure allows us to interpret values above (below) zero as a tight (loose) US monetary policy stance.

In Figure 2 the monetary policy surprise instruments, $z_{EA,t}$, and the US monetary policy stance, z_t^*, are shown. Several features stand out. From the beginning of our sample until the onset of the financial crisis, the Fed pursued a restrictive policy, reflected in a tight stance. To cope with the effects of the crisis, it cut policy rates sharply and introduced additional unconventional measures to counteract the downturn. Most importantly, these measures were designed to support the severely impaired interbank market to maintain liquidity in times of severe financial turmoil and to overcome the constraints imposed by the zero lower bound. The Fed continued its unconventional measures in the form of large-scale bond purchases. As a result, the stance remained rather accommodative until the first quarter of 2017, after which it shifted to a tighter stance given the intention to normalize the balance sheet. The ECB's policy surprises, on the other hand, reveal different information. While the period of the financial crisis was characterized by high upside and downside surprises, in early 2011 the ECB tightened its stance in two consecutive quarters. The benchmark interest rate (on the main refinancing operations) was set at 1.5% in July, leading to large tightening surprises. Shortly thereafter, the euro area entered a recession, prompting the ECB to adopt a more expansionary policy, with expansionary surprises. In subsequent periods, the surprises have been mixed. Interestingly, at the end of the sample, starting in early 2019, the figure shows the more hawkish stance of the Fed and less pronounced surprises from the ECB.

3.3 The interaction effects of two monetary policies

A simple first way to provide empirical evidence on the conditional effects of monetary policy is shown in Figure 3. Here we use the policy surprise measure in basis points for the ECB (Altavilla *et al.*, 2019) and the Fed's policy stance as discussed in the previous section. We first split our sample according to whether

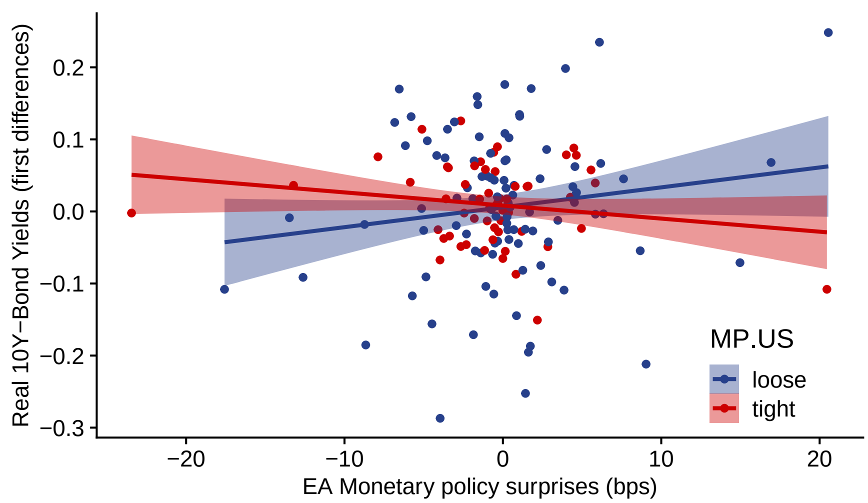
Figure 2: EA Monetary policy surprises (black) and US monetary policy stance (red).



Notes: The black bars denote the EA monetary policy surprises of Altavilla *et al.* (2019), while the red line refers to the US monetary policy stance ($DFF_t - r^*$). Gray shaded areas indicate EA recessions based on the Euro Area Business Cycle Dating Committee.

the Fed's stance (MP.US) is tight or loose, i.e., greater or less than zero. Within each subsample, we then regress the dependent variable – the differential real yield on 10-year AAA-rated European government bonds (computed by subtracting the 10-year inflation swap from the corresponding nominal yield) – on the euro area monetary policy surprise. This gives us a first rough measure of the impact of the ECB's monetary policy *conditional* on the respective US policy stance. Looking at the reactions when the US is in a loose monetary

Figure 3: Effect of an ECB monetary policy shock on real yields *conditional* on the Fed's policy stance.



Notes: The dependent variable is the differenced real yield of 10-years EA government bonds (computed by subtracting the 10-year inflation swap from the corresponding nominal yield). Monetary policy surprises are the sum of the monetary instruments of Altavilla *et al.* (2019) in basis points (bps) and MP.US denotes the Fed's policy stance computed by subtracting an r^* measure from the effective Fed funds rate. Shaded areas show the 90%-confidence intervals and red (blue) dots show ECB surprises during tight (loose) US monetary policy episodes.

policy regime (shown in blue), we observe that unexpected rate hikes in the euro area increase real yields, while rate cuts reduce them. On the other hand, when the Fed's stance is more hawkish (shown in red), the effect of an ECB tightening triggers a negative reaction in euro area real bond yields. We observe a decline in real yields following a surprise rate hike. To extend this picture, we further analyze the *full* term structure of yields and inflation expectations by using the extracted factors of Eq. (2.1). We decompose the picture into nominal yields and inflation expectations and check the relevance of a contemporaneous interaction effect.

Specifically, we use our constructed Nelson-Siegel factors of the nominal yield and inflation swap curve (see Figure 1) and estimate

$$y_{it} = \beta_{i0} + \beta_{i1}\text{MP.EA}_t + \beta_{i2}\Delta\text{MP.US}_t + \gamma_i(\text{MP.EA}_t \times \Delta\text{MP.US}_t) + \epsilon_{it} \tag{3.1}$$

where $i \in \{L, S, C\}$ denotes the level, slope, and curvature factor of the yields and swaps in their level values, MP.EA is the EA monetary instruments used, ΔMP.US is the change in the Fed's policy stance, and ϵ_{it} is an error term. The interaction effect $\text{MP.EA}_t \times \Delta\text{MP.US}_t$ is quantified by the coefficient γ. We report the results of this exercise in Table 1.

We observe that the constant is highly significant in all specifications, while euro area policy surprises do not have a statistically significant effect on any factor. Here, however, the effects of changes in the US policy stance are quite pronounced on both the level factors and the swap curve curvature factor. When the Fed turns hawkish, the level of the yield curve compresses. The slope and curvature factor of the yield curve are not affected at all, while the curvature of the swap curve also decreases. For the level of the swap curve, we also observe a pronounced compression during a tighter US stance. This could be a first indicator of the amplification of EA monetary policy during a tight US stance.

Although this univariate analysis suffers from several shortcomings (such as low explanatory power and lack of dynamics), it motivates further research through a unified analysis. The exercise has revealed some contemporaneous interaction effects of US monetary policy on euro area policy transmission, but has so far remained silent on the dynamics. To gain a deeper understanding, we therefore provide evidence based on our ST-VAR in the following sections.

Table 1: Yield and inflation swap curve factors and monetary policy interactions

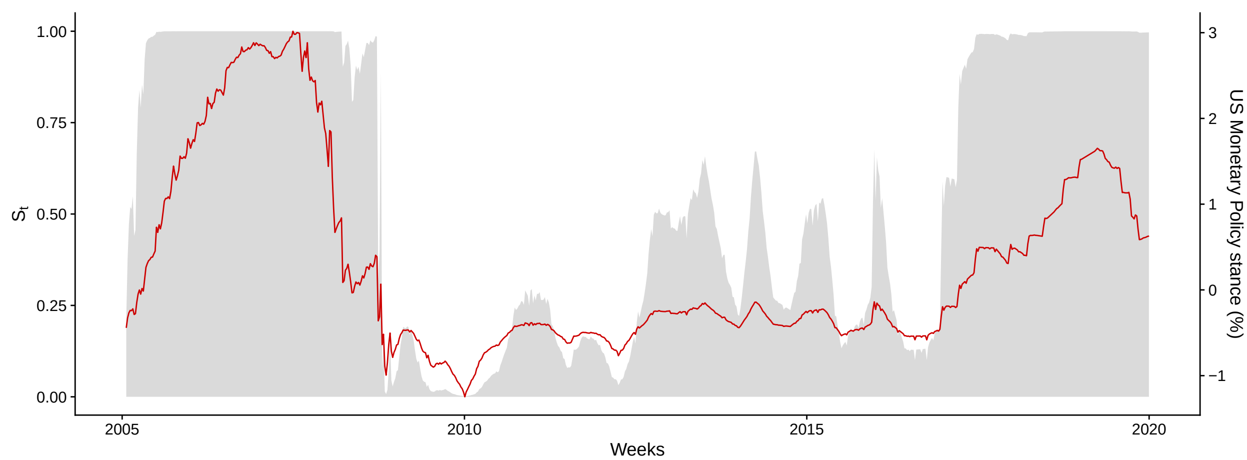
	Dependent variable:					
	Yield Curve factor			Inflation Swap Factor		
	L_r	S_r	C_r	L_π	S_π	C_π
Constant	3.535^{***}	-3.828^{***}	0.007	2.137^{***}	-0.573^{***}	-1.375^{***}
	(0.044)	(0.086)	(0.142)	(0.011)	(0.028)	(0.042)
MP.EA	−0.019	−0.027	0.006	−0.006	0.002	−0.031
	(0.019)	(0.042)	(0.064)	(0.005)	(0.014)	(0.021)
ΔMP.US	-1.312^{**}	−0.625	−0.200	-0.268^{**}	0.522	-1.200^{**}
	(0.595)	(0.952)	(1.865)	(0.111)	(0.336)	(0.557)
MP.EA×ΔMP.US	-0.957^{***}	−0.452	−1.080	-0.190^{**}	0.578^{**}	-1.340^{***}
	(0.316)	(0.728)	(1.279)	(0.074)	(0.282)	(0.324)
Observations	747	747	747	747	747	747
R^2	0.014	0.002	0.001	0.010	0.008	0.023

Notes: Ordinary least squares regression of the respective Nelson-Siegel factor on a constant, the monetary policy surprises for the euro area (MP.EA), the first differences of the US monetary policy stance (ΔMP.US) and the interaction effect with the EA surprises (MP.EA×ΔMP.US). Heteroskedasticity-consistent standard errors are reported in parentheses, *, ** and *** denote significance on the 10%, 5% and 1% level.

3.4 Monetary Policy Regimes

According to our specification in Eq. (2.2), the interaction between the US monetary policy stance and the coefficients of our VAR system is encoded by the function g in Eq. (2.3). An important intermediate result is thus reported in Figure 4, showing the posterior mean of S_t (gray area, left scale) alongside the threshold variable over time (red line, right scale). Thus, it provides information about how $S_t = g(z_t^*)$ evolves over time. In principle, values of S_t close to zero indicate periods characterized by a predominantly expansionary (i.e., negative) Fed stance, while values close to unity indicate periods characterized by a restrictive (i.e., positive) Fed stance. Prior to the global financial crisis, S_t was close to unity in almost all periods. This indicates that the corresponding VAR coefficients were close to $\boldsymbol{A}_{1j}$. During the financial crisis, the Fed lowered its policy rate rapidly and introduced unconventional monetary policy measures to overcome the restrictions of the zero lower bound. This is reflected in a persistent decline in S_t to values close to zero. After the easing periods associated with the financial crisis, we observe a generally expansionary stance until late 2017, but with some signs of tightening.

Figure 4: State allocation for the EA with the US monetary policy stance as threshold variable.



Notes: Gray areas denote the posterior mean of S_t, the transition function, where 0 (1) indicates an expansive (restrictive) regime. The red line depicts the constructed US monetary policy indicator in percent, as defined in Section 3.2.

Given the gradual improvement in US economic conditions from 2013, the Fed began to scale back some of the expansionary measures, most notably through the tapering of the third large-scale asset purchase program, which began in January and ended in October 2014. The Fed's balance sheet remained at the same level until the fall of 2017, when the Fed began a slow unwinding. In addition, the December 2015 FOMC meeting marked a period where the Fed decided to gradually raise interest rates. This led to a gradual tightening with one hike in 2016, three hikes in 2017, and four hikes in 2018, each by 25 basis points. Again, these developments are captured by increasing values of S_t. Thereafter, our model overwhelmingly points to a hawkish stance by the end of the sample. Equipped with this information, we now proceed with the dynamic analysis.

4 Dynamic Responses to a Monetary Policy Shock in the Euro Area

In this section, we discuss the responses to a restrictive monetary policy shock in the euro area. These responses are computed while explicitly conditioning on the US monetary policy stance. This exercise allows us to answer the question of how the monetary policy effect might change in the face of a divergent Fed policy. We therefore report and analyze impulse response functions (IRFs) based on our framework outlined in detail in Section 2.

To simplify the exposition, we provide two sets of results. The first analyzes how a restrictive EA shock affects our quantities of interest when US monetary policy is *tight*, which is achieved by setting the threshold variable z_t^* equal to the 95% percentile. The second is characterized by the 5% percentile of the threshold variable, reflecting a more *loose* US monetary policy stance. This is shown in the top panel of the following figures. In addition, and shown in the middle panel, we report for each IRF the difference in responses between the scenarios over the full horizon. To investigate whether the IRFs change smoothly with z_t^* or whether there are possible abrupt shifts in the transmission mechanisms, we focus on the short-run (one week ahead) responses for different quantiles of z_t^*, shown in the bottom panel.

The next section begins with a discussion of the responses of government bond yields and inflation swaps. These responses can then be combined to examine the responses of real yields. Finally, we briefly discuss the remaining reactions of macrofinancial variables and conclude the section with a robustness exercise.

4.1 Reactions of the term structure of bonds and inflation swaps

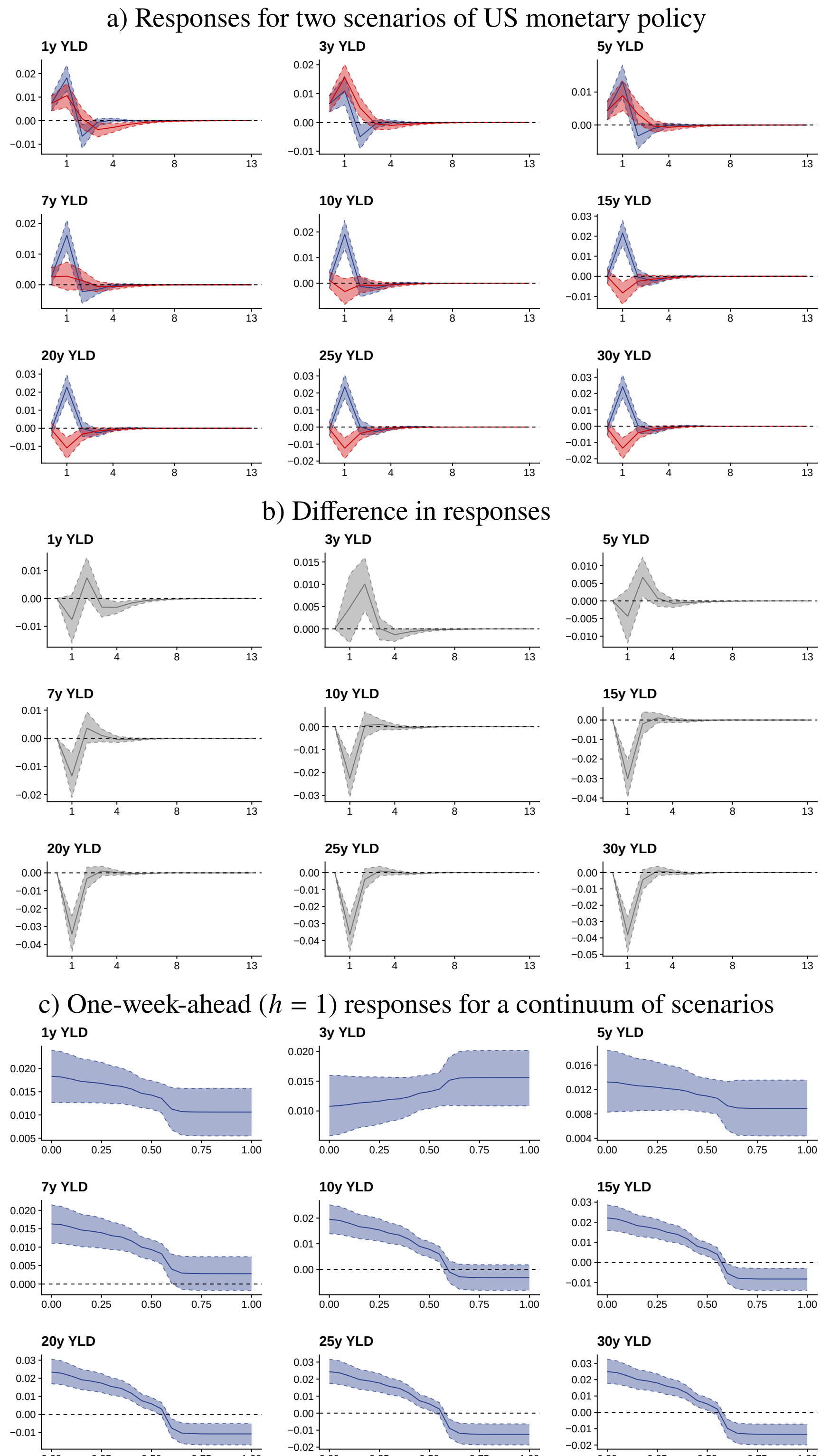
We begin by considering the dynamic responses of the EA term structure of government bonds to a restrictive ECB monetary policy shock in Figure 5. Panel (a) of the figure shows the impulse response functions to a one-standard deviation restrictive EA monetary policy shock under *restrictive* US monetary policy (red shaded areas) and under *expansive* monetary policy (blue shaded area), respectively. Thus, the first case corresponds to a euro area policy shock that falls under a restrictive Fed monetary policy stance, while the second case shows the results under an expansionary monetary policy stance. Note that the chosen percentiles of the US stance correspond to rather strong stances. This allows us to highlight the differences more clearly.

In general, government bond yields rise after a restrictive monetary policy shock, but to a lesser extent at longer maturities. While this observation holds across our two sets of results, the subsequent dynamic response depends strongly on the respective stance of the Fed. Starting with panel (a), by comparing the blue and red lines, we immediately observe that yields react more strongly when the restrictive surprise falls into an expansive US policy stance. For longer maturities, the differences between the posterior distributions of the IRFs (shown in panel (b)) are substantial, statistically significant, and mostly negative for up to a single week after the shock hits the system. Moreover, as maturities increase, the difference also increases from 2 to 4 bps. This implies that euro area yield reactions are much stronger when the Fed's monetary policy is characterized by a (strong) expansionary stance. After one week, however, this picture is reversed for almost all maturities. While medium- and long-term yields return to steady state within a month in both cases, short-term bond yield reactions reverse their contemporaneous increases and eventually turn slightly negative in the case of an expansionary US stance. We conjecture that the stronger reaction of short-term bonds may be due to a simple carry trade argument. If the ECB surprises on the upside (i.e., signals a more hawkish stance) while the Fed is currently in a more accommodative regime, yield differentials will naturally widen. This leads to an increased attractiveness of euro area assets for international investors and generates stronger positive reactions in the bond markets. After about two weeks, this increased demand for EA bonds is reflected in rising bond prices and falling yields across all segments of the yield curve. As shown in panel (b), spreads at the short end of the curve are somewhat less affected, but exhibit some overshooting.

Since this discussion is based on considering two extremes, we now discuss how the one-week-ahead responses change for different values (in terms of quantiles) of z_t^*. The reason we focus on the one-week-ahead response is that it is at this horizon that the differences between the two scenarios are greatest, underscoring the rapid reaction of financial markets and the need to look at higher frequency responses.[4] In summary, these figures confirm the above findings that restrictive EA monetary policy is associated with larger yield reactions in the first week after the shock when the Fed's stance is not restrictive, and vice versa. In addition, this figure provides some information about the specific shape of the g function. Note that the steep part of the logistic function lies in the interval $[0.25, 0.60]$. Thus, when the Fed's policy stance is large and negative (i.e., between the minimum and the 25% percentile, which is expansionary), one-week-ahead reactions are

[4] More results and details about the peak responses of all variables under scrutiny can be found in Appendix B.

Figure 5: Impulse responses of a restrictive EA monetary policy shock for selected government bond yields.



Notes: Panel (a) plots the responses to a contractionary monetary policy shock in the EA conditional on two scenarios: A restrictive US monetary policy stance (95% percentile of the threshold, colored in red) and an expansionary US monetary policy stance (5% percentile of the threshold, colored in blue). Panel (b) plots the differences between the two US regimes (restrictive minus expansionary stance). In both panels, the horizontal axis indicates the weeks following the shock. Panel (c) reports responses for a continuum of scenarios one week ahead ($h = 1$), where the horizontal axis denotes the respective percentile of the US MP stance. The shaded areas denote the 68% posterior credible set and the solid line denotes the posterior median.

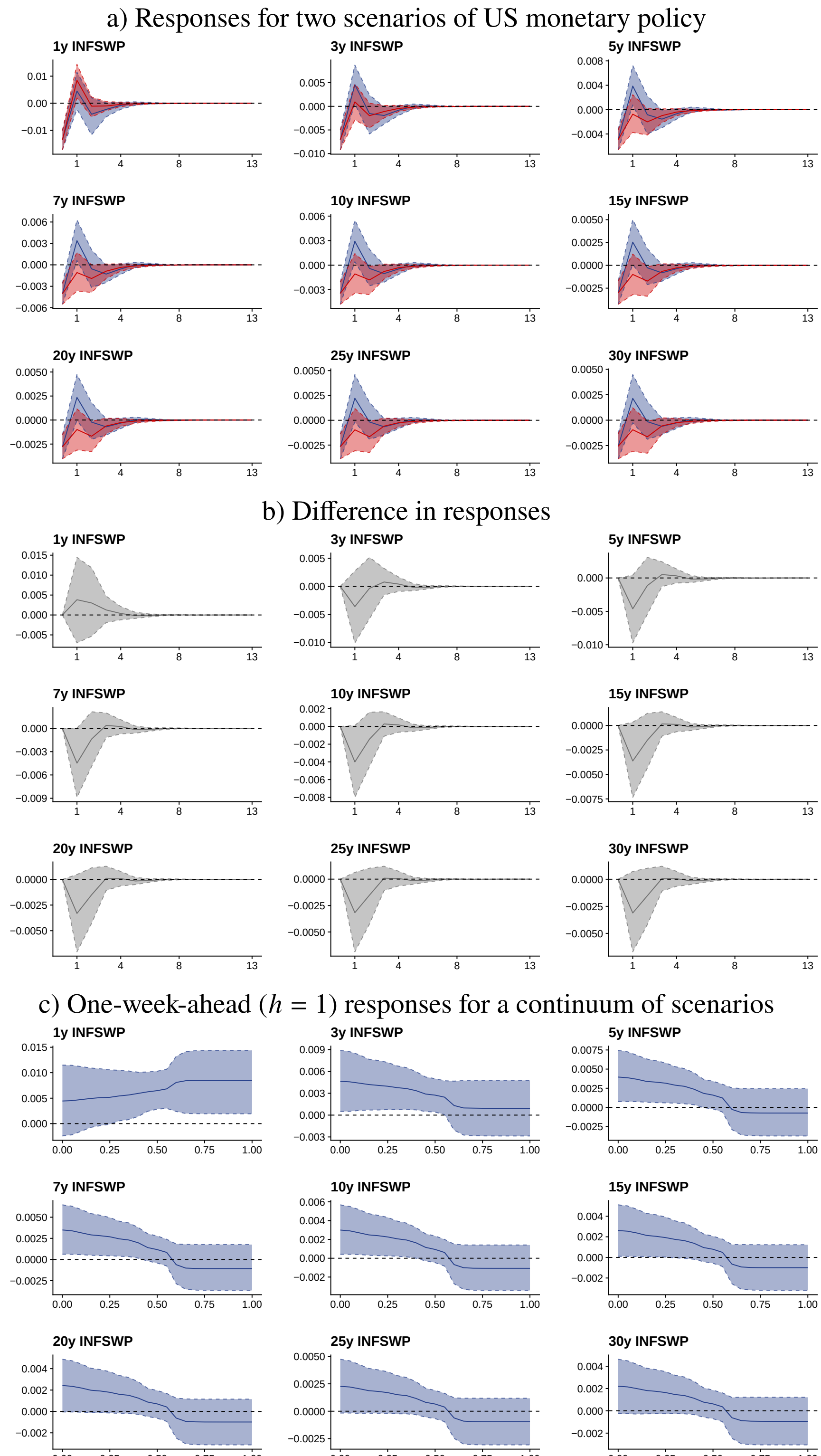
strong and somewhat more pronounced at higher maturities. However, for a US policy stance above the 60% percentile, the one-week-ahead responses are rather muted and even disappear for maturities between 7 and 10 years. For longer maturities, we observe a significant negative reaction when the 60% percentile is exceeded. Thus, at higher quantiles (i.e., the EA policy shock occurs when the Fed is hawkish), the yield curve responses are more muted or "wrong" signed. That is, following a restrictive ECB policy shock, yields fall slightly. The rather puzzling results may again point to the presence of rapid investor reactions due to carry trade strategies. Misaligned monetary policies not only lead to a significant interest rate differential, but may also reflect substantially different macroeconomic fundamentals to which the respective central banks may have reacted. This is supported by the observed peak dynamics within the more realistic interval $[0.25, 0.60]$, as exceptionally large divergences are rather rare in our sample. In this region, we observe that yield reactions are much more sensitive to changes in monetary policy divergence.

We continue to analyze how the inflation swap curve, and thus (more informally) the inflation expectations, shifts to restrictive EA monetary shocks conditional on a given US monetary policy stance. Figure 6, similar to Figure 5, plots the response of inflation expectations at different horizons. Panel (a) of Figure 6 shows that during a restrictive US stance, inflation expectations fall for all maturities considered (with somewhat stronger effects at the short end of the curve) and for the first four weeks. This is a common empirical finding and is consistent with the predictions of standard models (without taking into account potential spillovers). However, recent research by Coibion *et al.* (2021) highlights that it depends on the subjects who form their expectations about future inflation. For example, households and firms may react differently to interest rate hikes than professionals and financial market participants. This is of minor importance for our setting, as we analyze *market-based* expectations proxied by inflation-linked swaps.

In a situation where the shock occurs during an expansionary stance, we observe slightly different dynamic reactions. The immediate responses are very similar to the other case, but between one and four weeks after the shock we find that the full swap curve shifts upwards. This result could be indicative of some dampening effects stemming from the US monetary policy stance towards the euro area. In the case where the ECB policy shock occurs during a tight US stance, the reactions remain negative for all maturities.

However, panel (b) of Figure 6 shows that these differences appear to be rather insignificant over the impulse response horizon. Nevertheless, the main shape indicates a negative difference along maturities greater than one year. This negative difference is most pronounced in the intermediate segment of the curve.

Figure 6: Impulse responses of a restrictive EA monetary policy shock for selected inflation swaps.



Notes: Panel (a) plots the responses to a contractionary monetary policy shock in the EA conditional on two scenarios: A restrictive US monetary policy stance (95% percentile of the threshold, colored in red) and an expansionary US monetary policy stance (5% percentile of the threshold, colored in blue). Panel (b) plots the differences between the two US regimes (restrictive minus expansionary stance). In both panels, the horizontal axis indicates the weeks following the shock. Panel (c) reports responses for a continuum of scenarios one week ahead ($h = 1$), where the horizontal axis denotes the respective percentile of the US MP stance. The shaded areas denote the 68% posterior credible set and the solid line denotes the posterior median.

Looking at the one-week-ahead effects across a continuum of values of z_t^* in panel (c) of Figure 6, the one-week-ahead responses are generally positive when the US monetary policy stance is less than the median of z_t^*. Thus, the more expansionary the Fed's stance, the larger the *increase* in inflation swaps across all horizons, despite the ECB's hawkish surprise. As the inflection point of the continuum of reactions is located slightly to the right of 0.5, approaching the median of the US monetary policy stance, the reaction moves towards the intended one. That is, EA inflation swaps decline in response to a hawkish policy surprise. For a more hawkish Fed, this effect is reversed and the responses of inflation expectations turn negative for all maturities considered except for the short one-year maturity. However, the large uncertainty around the estimates renders these results mostly statistically insignificant. In addition, given the odd responses of short-term swaps, which increase along the entire continuum of US policy stances, we highlight two aspects. First, some factors outside the model may be driving the results. Second, since inflation-linked swaps are likely to incorporate risk premia, the responses may also reflect an increase in these premia following a policy surprise.

4.2 Assessment of the real rate effects

Our model also allows us to examine how real interest rates evolve after a restrictive EA monetary policy shock and how these reactions interact with a given US monetary policy stance. The corresponding real interest rate responses are shown in Figure 7 (a) - (c). Again, starting from a restrictive US stance, we see that when the euro area experiences a contractionary monetary policy shock, real interest rates rise for all maturities considered over the entire impulse horizon. However, the effect is most pronounced in the short and medium term and less pronounced for real rates at ten years and longer, eventually becoming insignificant in the impact. In contrast, when the shock occurs during an expansionary Fed stance, we observe heterogeneous dynamics across the term structure. While shorter maturities up to five years start to decline before mean reversion sets in after one to two weeks, medium and longer maturities show an increase.

This suggests that if monetary policy is judged by its effect on real interest rates, a policy surprise during a restrictive U.S. stance leads to positive real interest rate reactions in the first four weeks after the shock. Under an expansionary Fed stance, the immediate increase in real rates turns into a sharp negative reaction after about two weeks for short-term maturities. The differences between the IRFs do not include zero and are thus statistically significant, except for some medium-term maturities, as shown in panel (b). According to the construction of real interest rates, their strong reactions are a consequence of the more pronounced reactions of government bonds under an expansionary US stance. Interestingly, the different reaction patterns seem to be less pronounced in the middle segment of the real yield curve.

The shape of the response across quantiles of the Fed's stance is increasing for real yields up to five years, flattening for intermediate maturities, and decreasing for longer maturities. If nonlinear effects are present, we find that changes at the one-week-ahead horizon are strongest in the interval of the $[0.25, 0.40]$ quantile of z_t^*, comparable to the responses of the yield curve. If the shock occurs during a restrictive US policy stance, the effects on real rates become substantial and statistically significant, but again with diminishing strength at longer maturities. If the Fed remains expansionary (defined by stances between the minimum and the 25% percentile) while the ECB surprises with a contractionary stance, we find no significant reactions for shorter-dated bonds at the one-week-ahead horizon, as the negative effect of inflation swaps offsets the change in nominal yields. Again, the responses of real yields show an opposite pattern for maturities above 10 years, with a small increase in the real rate for a divergent policy stance. In addition, we observe a muted response for an aligned policy stance, but with a significant degree of uncertainty.

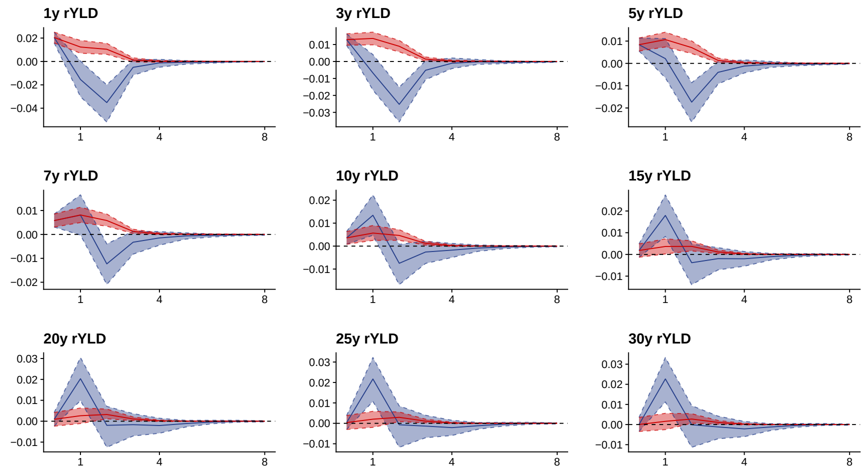
4.3 Reactions of stocks and exchange rates

Recall that our model includes not only bond yields and inflation swaps, but also the exchange rate, stock market returns, and the CISS. In panel (a) of Figure 8 we report the impulse responses for both scenarios of these three quantities along with the differences (panel (b)) and finally the responses for a continuum of scenarios in panel (c).

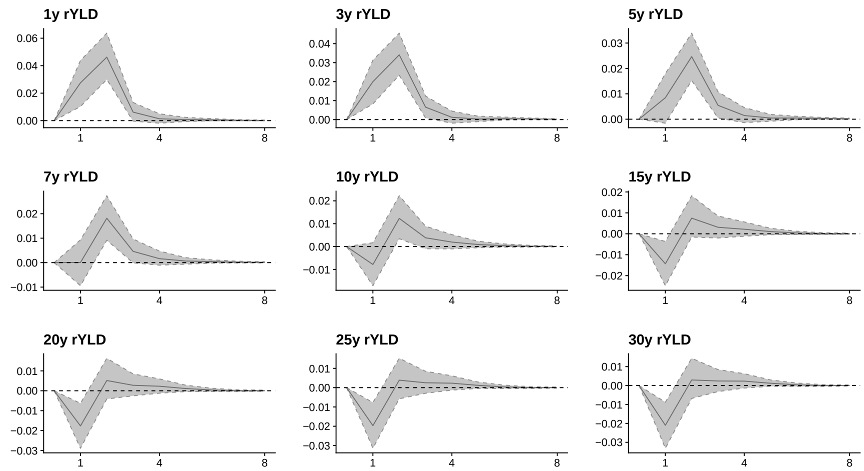
Starting with the Euro exchange rate vis-á-vis the US dollar, a contractionary monetary policy shock in the euro area should, ceteris paribus, lead to an appreciation of the Euro (see, among others, citeeichen-baum1995FX for empirical evidence on the effects of monetary policy on exchange rate dynamics). Our results confirm these expected dynamics, as we observe an appreciation of the Euro against the US dollar for both expansionary and restrictive US policy stances and along the continuum of stances. However, the magnitude of the Euro appreciation depends on the respective stance. In the case of an expansionary US stance (i.e., yield differentials widen as we showed in Figure 5), we see a somewhat more pronounced appreciation of the Euro vis-á-US dollar, with about 0.2 points below the steady state after two weeks. However, this difference is not statistically significant compared to the strongly restrictive stance. It is surprising to see that the prevailing US stance does not matter for the exchange rate reaction after a contractionary ECB shock. This may reflect the special and dominant role of the US dollar in the global financial system. Note, however,

Figure 7: Impulse responses of a restrictive EA monetary policy shock for selected real rates.

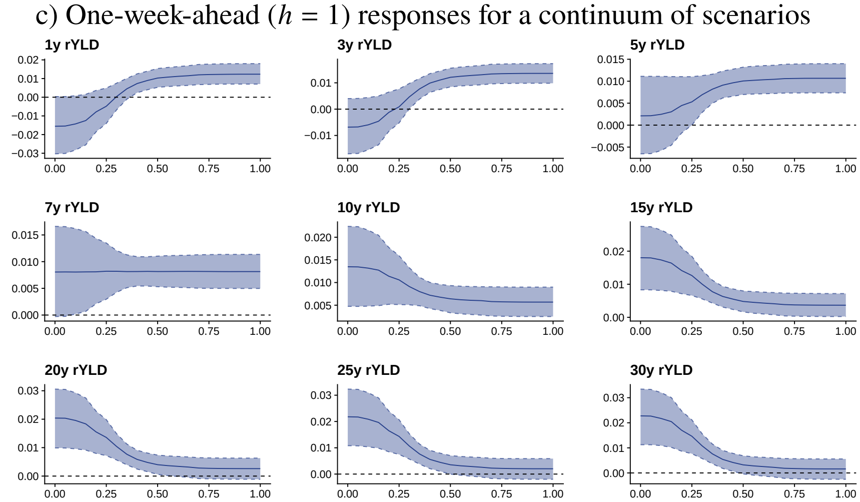
a) Responses for two scenarios of US monetary policy



b) Difference in responses



c) One-week-ahead (h = 1) responses for a continuum of scenarios



Notes: Panel (a) plots the responses to a contractionary monetary policy shock in the EA conditional on two scenarios: A restrictive US monetary policy stance (95% percentile of the threshold, colored in red) and an expansionary US monetary policy stance (5% percentile of the threshold, colored in blue). Panel (b) plots the differences between the two US regimes (restrictive minus expansionary stance). In both panels, the horizontal axis indicates the weeks following the shock. Panel (c) reports responses for a continuum of scenarios one week ahead (h = 1), where the horizontal axis denotes the respective percentile of the US MP stance. The shaded areas denote the 68% posterior credible set and the solid line denotes the posterior median.

that since we use only triple-A rated sovereign bonds, safe-haven motives from a euro area perspective are of minor importance here.

Turning to stock market reactions, we observe an immediate decline after a contractionary monetary policy shock. This is by construction to ensure the separation of monetary policy and information shocks. The reaction remains negative for a shock that occurs during a restrictive Fed stance. However, if the shock occurs during an expansionary Fed, we observe a strong rebound after about a week before mean reversion sets in. This result may be explained by the aggregate Euro Stoxx 50 index, which masks cross-sectional heterogeneity. For example, in a more granular exercise, Jarociński (2022) shows that US-exposed and foreign-exposed stocks, financial and non-financial stocks, and small and large-cap US stocks react differently to EA shocks. As the Euro Stoxx 50 is a broad EA-wide aggregate, similar effects can be expected based on exposure to rising foreign demand or foreign earnings. Here, potential further research could explore the responses of different granular subcomponents of European equity markets.

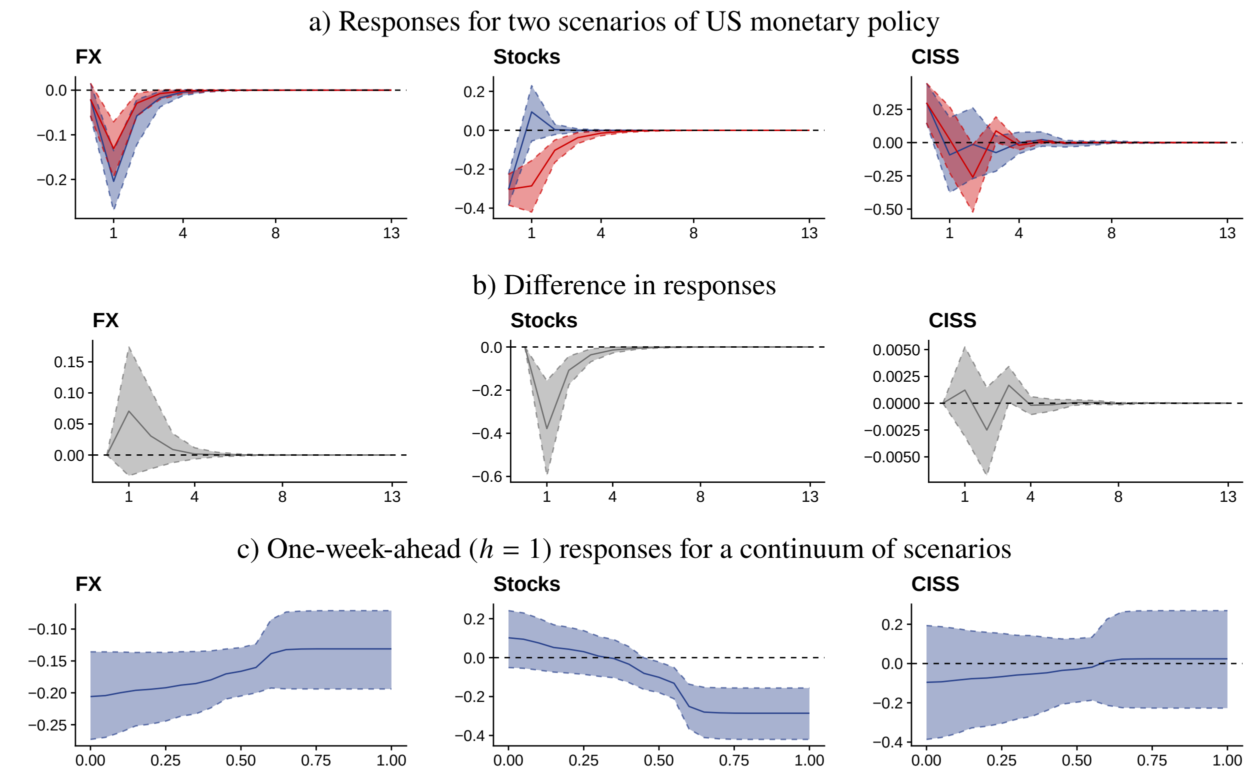
Finally, we observe a statistically significant increase in the composite indicator of systemic stress (CISS), but with negligible differences conditional on the US stance. Neither the one-week responses nor the differences between the two prevailing US stances are statistically significant for this variable. We suggest that our set of variables is sufficient to capture financial market reactions, including the implicit level of financial stress.

4.4 Robustness

In this section, we present some of the results of our robustness checks. To save space, we refrain from reporting detailed results and thus only briefly discuss our findings.[5] While financial variables are conveniently available at a daily/weekly frequency, most macro variables are not. Thus, the general availability of data on real quantities such as output, realized inflation or survey-based expectations represents a limit to this exercise in terms of additional variables. However, the monetary policy stance is a crucial aspect, as are the proxies for the shock. As monetary policy enters our setup in two ways, it is a natural candidate for a robustness check. On the one hand, the euro area monetary policy surprise is treated as observable in our ST-VAR, while the Fed's stance governs the transition between regimes. Thus, in this section we discuss (1) an alternative measure of the EBC monetary policy surprise and (2) the Fed's stance.

[5] All additional results are available from the authors upon request.

Figure 8: Impulse responses of a restrictive EA monetary policy shock for other quantities.



Notes: Panel (a) plots the responses to a contractionary monetary policy shock in the EA conditional on two scenarios: A restrictive US monetary policy stance (95% percentile of the threshold, colored in red) and an expansionary US monetary policy stance (5% percentile of the threshold, colored in blue). Panel (b) plots the differences between the two US regimes (restrictive minus expansionary stance). In both panels, the horizontal axis indicates the weeks following the shock. Panel (c) reports responses for a continuum of scenarios one week ahead ($h = 1$), where the horizontal axis denotes the respective percentile of the US MP stance. The shaded areas denote the 68% posterior credible set and the solid line denotes the posterior median. FX denotes the Euro per one US-dollar (i.e., a decrease denotes an appreciation of the Euro), Stocks the Euro Stoxx 50 index, and CISS the Composite Index of Systemic Stress.

4.4.1 Alternative Monetary Policy Surprises

In our main specification, we used the monetary policy surprise instruments proposed in Altavilla *et al.* (2019) for the EA (henceforth AB-surprises). We repeat our analysis using the database constructed by Jarociński and Karadi (2020) (hereafter JK-surprises) as a proxy for the shock, while keeping the difference between the effective federal funds rate and the (one-sided) natural rate estimate for the US stance. The surprise measure provides another convenient way to obtain policy surprises that are purged of information shocks. It is constructed by tracking the co-movement of interest rates and stock prices in a window about 10 minutes before and 20 minutes after a monetary policy announcement. While "pure" monetary policy shocks are identified by a negative co-movement between interest rate and stock price changes, "information" shocks move both variables in the same direction. Thus, the approach allows the decomposition of surprises into "pure" monetary policy and central bank information effects, which is necessary for our research question.[6] While we re-estimate our specification outlined in Section 2 with the three differently "purged" monetary instruments, we only report the results based on the *rotated* monetary policy surprises. Since we do not change the threshold variable in this exercise, the government allocation remains unchanged. The impulse response functions to a restrictive monetary policy shock in the euro area can be found in Appendix C.

The results show that our findings are generally robust to changes in the measure of monetary surprises used. However, some differences are worth mentioning. For the reactions of government bond yields, we observe the same shapes as in our main specification. However, the magnitude of the peak reaction under JK surprises is somewhat smaller compared to our main specification. Interestingly, the estimation uncertainty is lower for the estimates with AB surprises, especially for more hawkish stances in the US. A similar picture emerges for the inflation swap responses. While the general shape of the responses and the size of the peak response are not different, the uncertainty is larger. Only for one-year swaps do we find clear evidence of nonlinearities when using the JK surprises. The reaction of macrofinancial variables is also not significantly different from our main specification. Only the stock market reaction remains negative in the first month and does not show a positive reaction after one week compared to the main specification when the Fed's stance is expansionary. Both the CISS and the exchange rate show the same responses across the employed surprises. However, the one-week responses of the exchange rate are slightly stronger in our main specification and across Fed stances. Note, however, that the JK surprises neglect to some extent the unconventional nature

[6] We use the constructed shock series on a daily basis, available from https://marekjarocinski.github.io/, and aggregate them to weekly frequency.

of monetary policy in the aftermath of the global financial crisis. This may be one reason for the increased uncertainty around these robustness results.

4.4.2 Alternative US Monetary Policy Stance Indicator

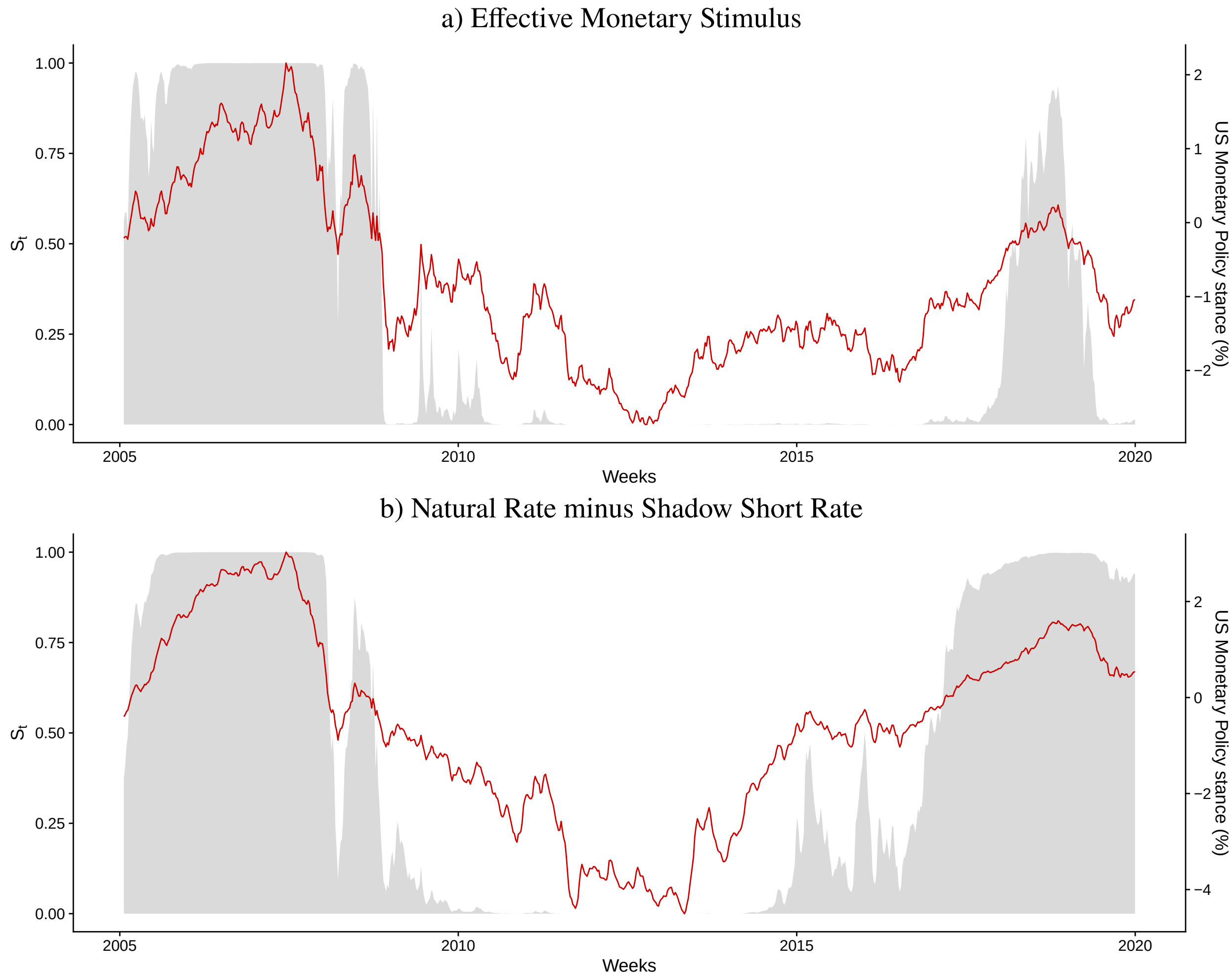
In our main specification, we use the difference between the effective federal funds rate and the (one-sided) natural rate estimate (Laubach and Williams, 2003) to measure the corresponding U.S. monetary policy stance. Here we present results based on two alternative candidates. First, we use the Effective Monetary Stimulus (EMS) compiled by Halberstadt and Krippner (2016). This is a monetary policy measure based on yield curve data that reflects both conventional and unconventional monetary policy environments. As a second measure of the Fed's stance, we again subtract the (one-sided) natural rate estimate (Laubach and Williams, 2003) from a shadow short rate also developed by Halberstadt and Krippner (2016).[7]

Starting with the regime allocation shown in Figure 9, we observe a relatively similar picture until the onset of the financial crisis. Only some probabilities are slightly less restrictive compared to our main specification. While our main specification shows strong support for a restrictive regime from late 2016 until the end of the sample, the EMS measure assigns a lower probability to it. The results with both alternative measures do not support a restrictive stance in the period from 2010 to 2015. A possible explanation lies in the nature of the monetary policy stance series, which explicitly includes unconventional policy measures such as asset purchase programs. As a result, the allocation to the expansionary regime between 2009 and 2015/2016 is much more pronounced in the specification with the EMS metric. Thus, the different measures of the US monetary policy stance used as a threshold variable suggest slightly different regime allocations and thus potentially different responses to an EA monetary policy shock.[8]

Interestingly, the reactions across the yield curve to an EA policy shock remain very close to our benchmark results, despite a somewhat different regime allocation. The shape of the responses for the full continuum of US stances also looks qualitatively very similar. However, for the EMS metric, which shows less support for restrictive stances in the post-financial crisis period through 2016, the shifting reactions between expansionary and restrictive stances are somewhat less smooth. For inflation swaps, the model with the EMS as the threshold variable, the differences disappear and roughly resemble the shape of the expansionary scenario in our benchmark model. The results for the second metric are very similar to our

[7] Both measures are available on a daily basis and have been downloaded from https://www.ljkmfa.com/.
[8] To conserve space, the full set of results displayed like in Section 4 can be obtained by the authors upon request.

Figure 9: State allocation for the EA with two US monetary policy metrics as threshold variable.



Notes: Gray areas denote the posterior mean of S_t, the transition function, where 0 (1) indicates an expansive (restrictive) regime. The red line depicts the US monetary policy indictator in percent. In panel (a) this is the Effective Monetary Stimulus (EMS) compiled by Halberstadt and Krippner (2016), while in panel (b) a Shadow Short Rate developed by Halberstadt and Krippner (2016) was substracted by an estimate of the natural interest rate Laubach and Williams (2003).

main specification, but with less sensitivity to changes in stance up to the turning point. Finally, there is an exchange rate difference between the scenarios for both metrics, but it is small. Stock market reactions show an almost identical shape to the benchmark only for the second metric, with an insignificant positive reaction after one week. The proxy for financial stability shows no differences and is not significant.

To conclude the robustness section, a common feature of all results is the strong evidence of nonlinearities shown by the significant transition variable proxying the Fed's monetary policy stance. Thus, we again conclude that it does matter for euro area monetary policy whether the US stance is expansionary or contractionary and, more importantly, how strong this stance is.

5 Concluding remarks

Monetary policy in a globalized world depends not only on prevailing macroeconomic fundamentals, but also strongly on how central banks interact. This is because key transmission channels have an explicit international dimension, and effective monetary policy depends not only on domestic first-round effects but also on the extent of international spillovers in financial and real markets. In this paper, we develop an empirical weekly time series model that explicitly models the interactions between domestic macrofinancial dynamics in the euro area (EA) and the US monetary policy stance in a nonlinear fashion. The identification of shocks is based on readily available high-frequency instruments designed to measure unexpected movements in monetary policy. The measurement of the US monetary policy stance is based on an estimate of the neutral interest rate.

In sum, our results suggest interaction effects between US and EA monetary policy. For nominal bond yields, we observe that contractionary EA monetary policy has stronger effects when the Fed is following an expansionary stance. These effects reverse after about two weeks. At this horizon, bond yields tend to fall in response to capital inflows into the EA economy, which pushes up bond prices. In contrast, when the EA shock occurs under a restrictive Fed stance, we observe somewhat weaker reactions with heterogeneity across maturities. Looking at the reactions of the inflation swap curve, we observe muted differences when controlling for the US stance. In general, the results point to negative reactions at impact, with different dynamics thereafter depending on the prevailing US monetary policy stance. However, even though they are marginally significant, there is some evidence of nonlinearities. Finally, our model allows for the analysis of real interest rate reactions by computing the differences in IRFs between the nominal bond yield and the inflation swap reactions. Consideration of these reactions suggests that real rates rise in all scenarios, but tend to be stronger when EA shocks occur during a restrictive Fed stance.

The model we use in this paper uses a simple and straightforward way of interacting foreign monetary policy with domestic quantities. As a further avenue of research, it would be possible to allow for a more flexible law of motion (i.e., using nonparametric techniques to remain agnostic about the precise functional form of the interaction effect). In addition, we have only included EA-based quantities and capture international spillovers mainly by including the US monetary policy stance. Estimating larger bi-country models would in principle be feasible, which we leave for further research.

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Appendix A Data

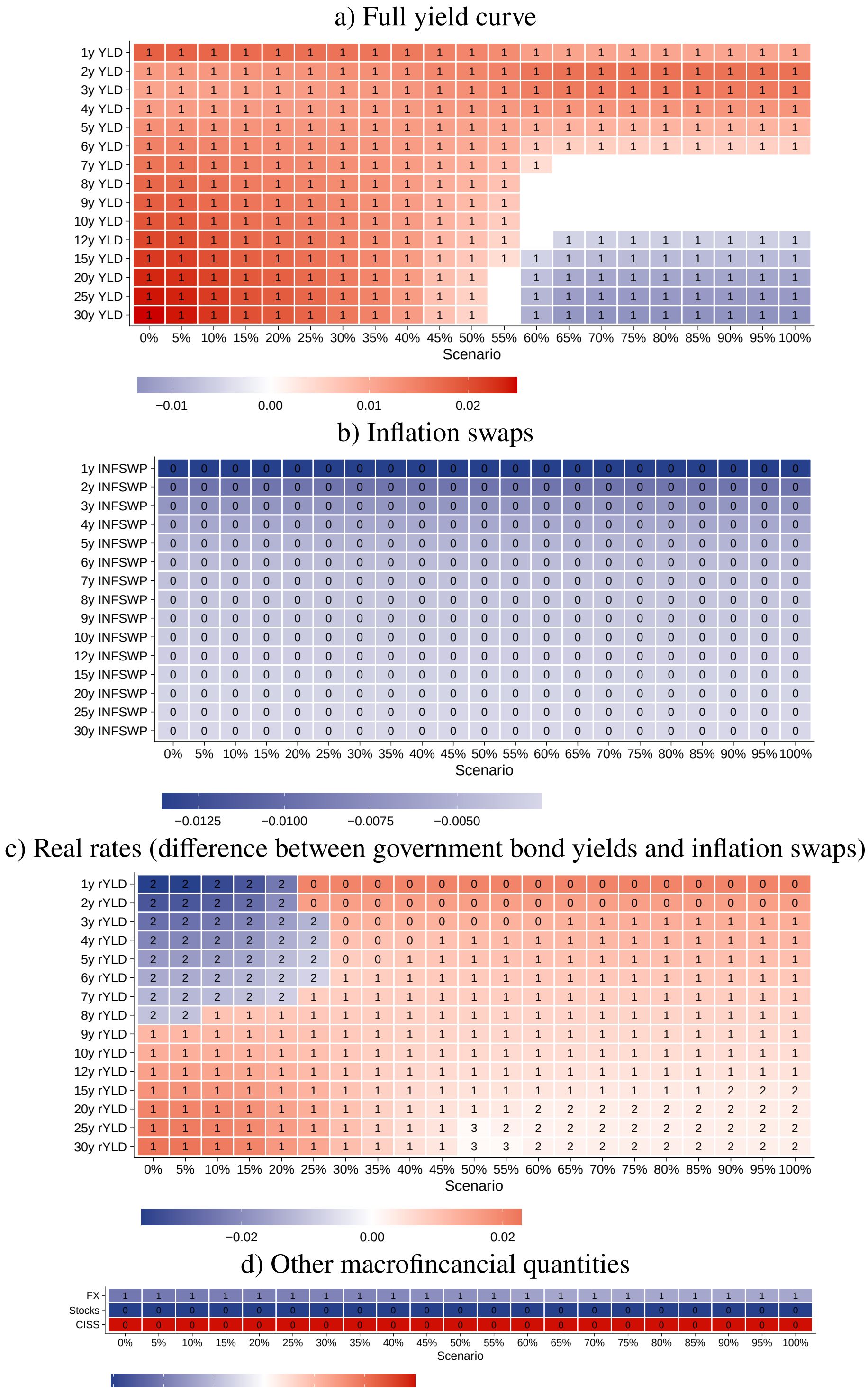
Table A1: Data description

Label	Mnemonic	Description	Frequency	Transformation
		(a) Euro Area		
Yields	$\texttt{YLD}^{\tau Y}$	EA government bond yields (AAA-rated) with maturity τ years ahead	daily, aggregated to weekly	first differences
ILS	$\texttt{INFSWP}^{\tau Y}$	Inflation-linked swaps τ- year ahead	daily, aggregated to weekly	first differences
EuroStoxx50	`Stocks`	Euro area stock market index	daily, aggregated to weekly	percentage changes
CISS	`CISS`	Euro area composite indicator of systemic stress (Kremer *et al.*, 2012)	daily, aggregated to weekly	first differences
EUR-USD Exchange Rate	`FX`	Exchange rate in EUR per USD	daily, aggregated to weekly	percentage changes
		(b) United States		
Neutral rate	r^*	estimate of the natural rate of interest (Laubach and Williams, 2003)	quarterly, interpolated to weekly	level (One-sided)
Policy Rate	`DFF`	Federal Funds Effective Rate	daily, aggregated to weekly	level
Effective Monetary Stimulus	`EMS`	Effective Monetary Stimulus (Halberstadt and Krippner, 2016)	daily, aggregated to weekly	level
Shadow Short Rate	`SSR`	Shadow Short Rate (Halberstadt and Krippner, 2016)	daily, aggregated to weekly	level

Notes: The inflation-linked swaps are gathered from Macrobond, while the remaining EA variables are obtained from the ECB's Statistical Data Warehouse (https://sdw.ecb.europa.eu/). For computing the US monetary policy stance, we obtained the `DFF` from the FRED database of the Federal Reserve Bank of St. Louis (fred.stlouisfed.org) and the real rate estimate of Laubach and Williams (2003) from the database of the Federal Reserve Bank of New York (www.newyorkfed.org/research/policy/rstar). The latter is available on a quarterly basis and was interpolated to match weekly frequency. All remaining variables were obtained on a daily basis and then aggregated to weekly frequency by taking the average. Both alternative US monetary policy measures are available on a daily basis and were downloaded from https://www.ljkmfa.com/.

Appendix B Further empirical results

Figure B.1: Peak responses of a restrictive EA monetary policy shock for a set of US monetary policy scenarios.



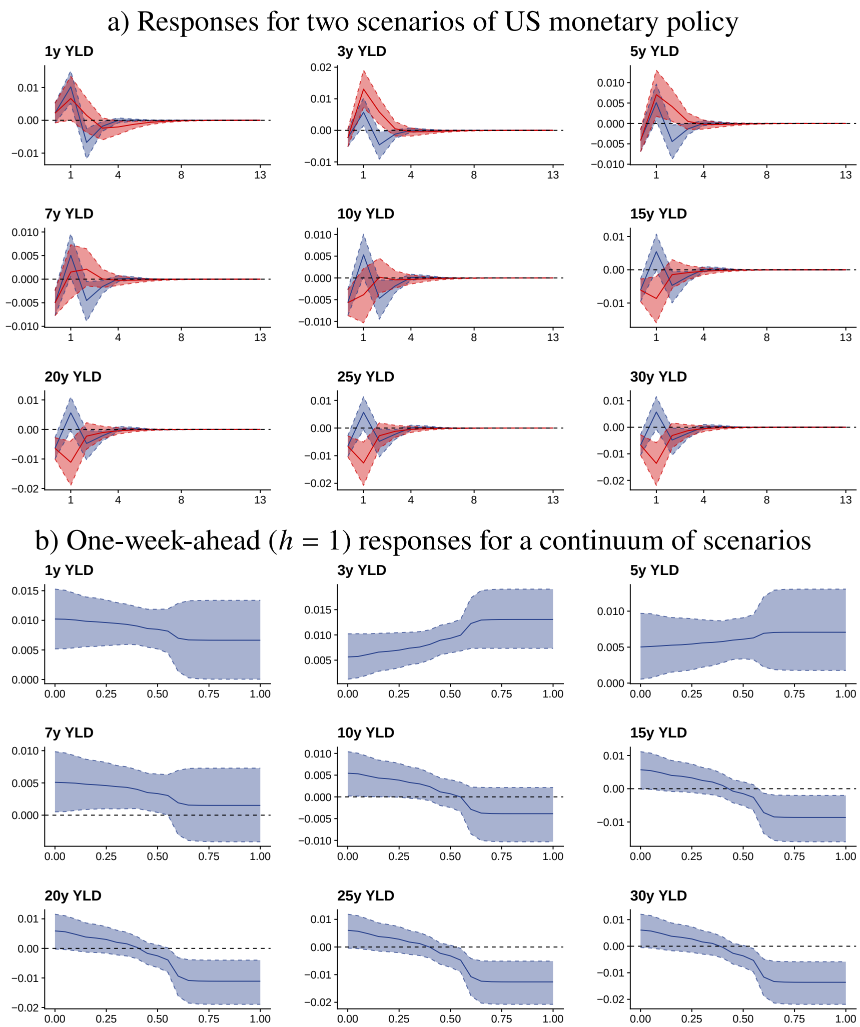
Notes: While red (blue) shaded cells indicate significant negative (positive) peak responses, the cell number refers to the week after the shock at which the a significant response peaks. Significance is defined with respect to the 68% credible set.

Appendix C Robustness Checks

C.1 Robustness results based on a different monetary policy instrument

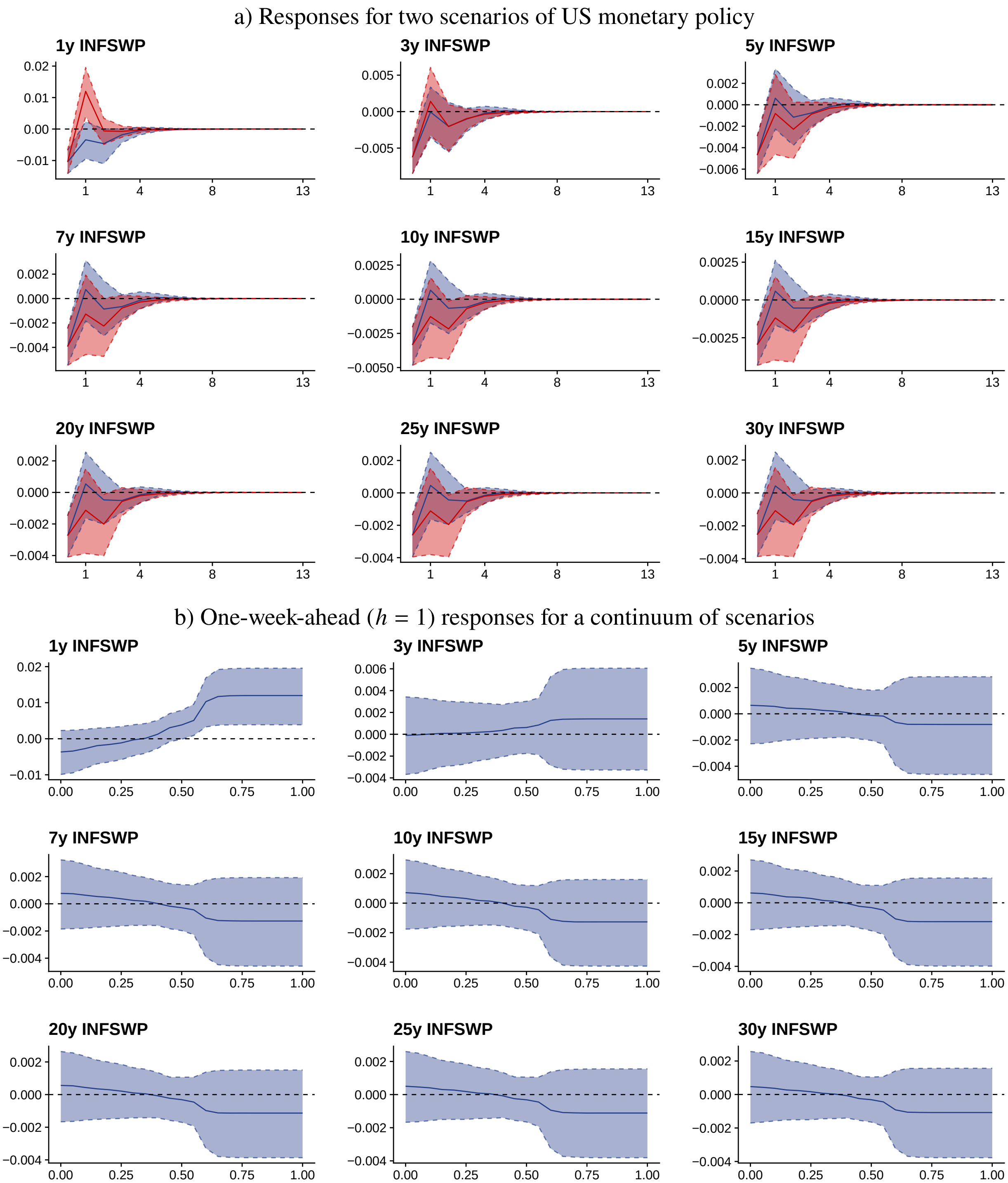
In this subsection, we provide additional results to the robustness checks discussed in Section 4.4. We follow the specification from Section 2 but use the monetary policy instruments compiled by Jarociński and Karadi (2020) instead. The estimation specification and the prior setup remain unchanged.

Figure C.1: Impulse responses of a restrictive EA monetary policy shock for selected government bond yields.



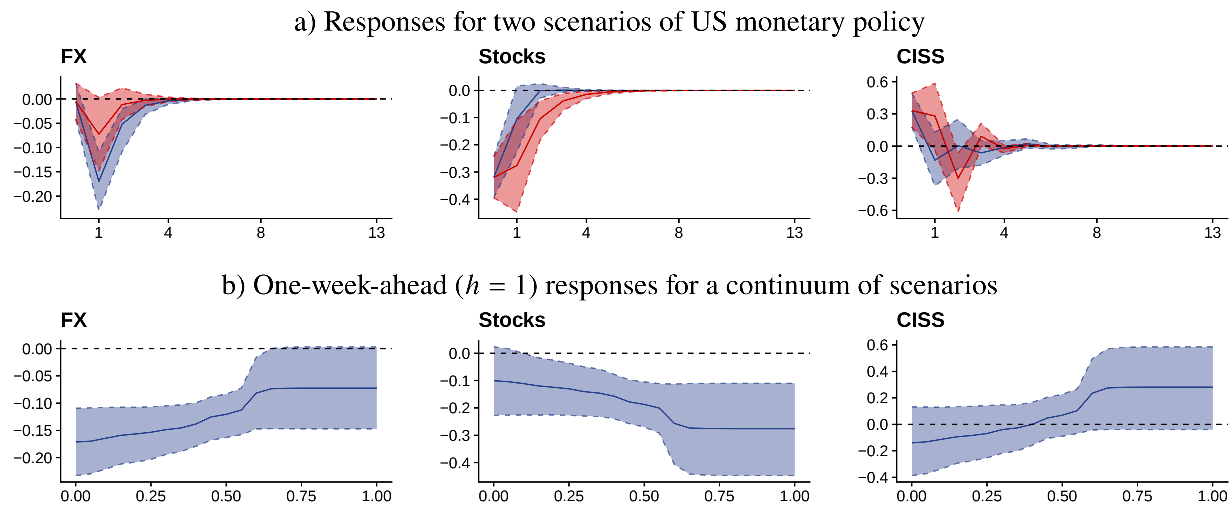
Notes: Panel (a) plots the responses to a contractionary monetary policy shock in the EA conditional on two scenarios: A restrictive US monetary policy stance (95% percentile of the threshold, colored in red) and an expansionary US monetary policy stance (5% percentile of the threshold, colored in blue). Panel (b) reports responses for a continuum of scenarios one week ahead ($h = 1$), where the horizontal axis denotes the respective percentile of the US MP stance. The shaded areas denote the 68% posterior credible set and the solid line denotes the posterior median. Estimates are based on monetary policy surprises obtained from Jarociński and Karadi (2020).

Figure C.2: Impulse responses of a restrictive EA monetary policy shock for selected inflation swaps.



Notes: Panel (a) plots the responses to a contractionary monetary policy shock in the EA conditional on two scenarios: A restrictive US monetary policy stance (95% percentile of the threshold, colored in red) and an expansionary US monetary policy stance (5% percentile of the threshold, colored in blue). Panel (b) reports responses for a continuum of scenarios one week ahead (h = 1), where the horizontal axis denotes the respective percentile of the US MP stance. The shaded areas denote the 68% posterior credible set and the solid line denotes the posterior median. Estimates are based on monetary policy surprises obtained from Jarociński and Karadi (2020).

Figure C.3: Impulse responses of a restrictive EA monetary policy shock for other quantities.



Notes: Panel (a) plots the responses to a contractionary monetary policy shock in the EA conditional on two scenarios: A restrictive US monetary policy stance (95% percentile of the threshold, colored in red) and an expansionary US monetary policy stance (5% percentile of the threshold, colored in blue). Panel (b) reports responses for a continuum of scenarios one week ahead ($h = 1$), where the horizontal axis denotes the respective percentile of the US MP stance. The shaded areas denote the 68% posterior credible set and the solid line denotes the posterior median. Estimates are based on monetary policy surprises obtained from Jarociński and Karadi (2020). FX denotes the EURUSD exchange rate (i.e., a decrease denotes an appreciation of the Euro), Stocks the Euro Stoxx 50 index, and CISS the Composite Index of Systemic Stress.

Appendix D Technical appendix

D.1 Prior set-up

For the prior setup, we closely follow Hauzenberger *et al.* (2021). Our priors are designed to provide sufficient regularization of the VAR coefficients to reduce estimation uncertainty. On the VAR coefficients, we therefore apply a Horseshoe (HS) prior (Carvalho *et al.*, 2010), which belongs to the class of global-local shrinkage techniques. To outline the mechanism of this prior, it proves useful to collect the state-specific coefficients in an $M \times N$-dimensional matrix $\boldsymbol{A}_k = (\boldsymbol{c}_k, \boldsymbol{\gamma}_k, \boldsymbol{A}_k p, \ldots, \boldsymbol{A}_k P)$, for $k = \{0, 1\}$ and $p = 1, ..., P$. In the following, let $\boldsymbol{a}_k = \text{vec}(\boldsymbol{A}_k)$ be a $K(= NM)$-dimensional vector, then, for $k = \{0, 1\}$, then the hierarchical HS prior on the jth $(j = 1, \ldots, K)$ state-specific coefficients is as follows:

$$a_{kj} \sim \mathcal{N}(0, \tau_k \lambda_{kj}), \quad \sqrt{\tau_k} \sim C^+(0,1) \quad \text{and} \quad \sqrt{\lambda_{kj}} \sim C^+(0,1).$$

Here, a global component τ_k pushes all coefficients in $\boldsymbol{a}_k$ toward zero (i.e., the prior mean), while local scaling parameters λ_{kj} are regime- and covariate-specific, providing sufficient flexibility. Both hyperparameters $\sqrt{\tau_k}$ and $\sqrt{\lambda_{kj}}$ follow a half-Cauchy distribution denoted by C^+. This prior is able to introduce sufficient sparsity via the global shrinkage component in a highly parameterized model that would otherwise be prone to overfitting. The local scales still allow for the detection of important individual covariates by mitigating the potential (over)shrinkage of the global parameter.

The prior specification on the variance-covariance $\boldsymbol{\Sigma}$ is standard. Here, $\boldsymbol{\Sigma}$ follows an inverted Wishart distribution:

$$\boldsymbol{\Sigma} \sim \mathcal{IW}(d_0, \boldsymbol{D}_0),$$

with the scalar d_0 denoting the prior degrees of freedom and the $M \times M$-dimensional matrix $\boldsymbol{D}_0$ referring to the prior scaling. In the empirical application, we specify $d_0 = M + 2$ and $\boldsymbol{D}_0 = 0.01\boldsymbol{I}_M$. This choice ensures that the prior is proper, but still weakly informative.

To complete our proposed setup, we also need to specify priors on the hyperparameters μ and ρ, which define the transition between the two regimes. The threshold parameter μ follows a weakly informative Uniform distribution:

$$\mu \sim \mathcal{U}\left(\min(z_t^*), \max(z_t^*)\right),$$

with the prior support being bounded between the minimum and maximum value of the signal variable z_t^*. On the speed of adjustment parameter ρ we impose a Gamma prior:

$$\rho \sim \mathcal{G}(s_0, s_1),$$

with the hyperparameters $s_0 = 7/0.001$ and $s_1 = 1/0.001$ chosen such that the transition between the two states occurs relatively fast. This informative choice ensures a sensible state allocation and a clear separation between the two regimes, one characterized by an expansionary US monetary policy and the other characterized by a restrictive stance (see also Figure 4 for the estimated state allocation).

D.2 Posterior quantities and sampling algorithm

Next, we outline the main steps of our Markov Chain Monte Carlo (MCMC) algorithm, which is used to simulate from the full posterior distribution. To obtain the conditional posteriors, we combine the priors with the likelihood of the model, as described in Eq. (2.2) and Eq. (2.3). For most parameters, this leads to well-known conditional distributions, and simulating from them is straightforward. Let the symbol $\bullet$ indicate that we condition on the remaining parameters of the system, then these sampling steps read as follows:

(i) For simplicity we collect all VAR coefficients in a $2K$-dimensional vector $\boldsymbol{a} = (\boldsymbol{a}_1', \boldsymbol{a}_0')'$ and define the corresponding $2K$-dimensional vector of predictors $\boldsymbol{x}_t = \left((1, z_t, \boldsymbol{y}_{t-1}', \dots, \boldsymbol{y}_{t-p}') \times S_t, (1, z_t, \boldsymbol{y}_{t-1}', \dots, \boldsymbol{y}_{t-p}') \times (1 - S_t)\right)'$. In what follows, we sample $\boldsymbol{a}$ from a multivariate Gaussian distribution:

$$\boldsymbol{a}|\bullet \sim \mathcal{N}\left(\bar{\boldsymbol{a}}, \bar{\boldsymbol{V}}\right),$$

with $\bar{\boldsymbol{a}}$ denoting the posterior mean and $\bar{\boldsymbol{V}}$ the posterior variance-covariance matrix:

$$\begin{aligned}\bar{\boldsymbol{V}} &= \left(\boldsymbol{\Sigma}^{-1} \otimes \boldsymbol{X}'\boldsymbol{X} + \underline{\boldsymbol{V}}^{-1}\right),\\ \bar{\boldsymbol{a}} &= \bar{\boldsymbol{V}}^{-1}\left(\boldsymbol{\Sigma}^{-1} \otimes \boldsymbol{X}'\text{vec}(\boldsymbol{Y})\right).\end{aligned}$$

Here, $\boldsymbol{X} = (\boldsymbol{x}_1, \dots, \boldsymbol{x}_T)'$ and $\boldsymbol{Y} = (\boldsymbol{y}_1, \dots, \boldsymbol{y}_T)'$ denote the full data matrices and $\underline{\boldsymbol{V}} = \text{diag}\left(\{v_{1j}\}_{j=1}^K, \{v_{0j}\}_{j=1}^K\right)$, where $v_{kj} = \tau_k \lambda_{kj}$, collects the prior variances of all coefficients on the main diagonal.

(ii) To draw the HS scaling parameters we rely on the following steps, outlined in Makalic and Schmidt (2015):

$$\begin{aligned}\tau_k|\bullet &\sim \mathcal{IG}\left(\frac{K+1}{2}, \frac{1}{\phi_k} + \sum_{j=1}^{K} \frac{a_{kj}}{2\lambda_{kj}}\right),\\ \lambda_{kj}|\bullet &\sim \mathcal{IG}\left(1, \frac{1}{\zeta_{kj}} + \frac{a_{kj}}{2\tau_k}\right),\end{aligned}$$

where the two auxiliary variables ϕ_k and ζ_{kj} are also drawn from inverted Gamma distributions:

$$\begin{aligned}\phi_k|\bullet &\sim \mathcal{IG}\left(1, 1 + \frac{1}{\tau_k}\right),\\ \zeta_{kj}|\bullet &\sim \mathcal{IG}\left(1, 1 + \frac{1}{\lambda_{kj}}\right).\end{aligned}$$

(iii) We sample $\boldsymbol{\Sigma}$ from an inverted Wishart distribution:

$$\boldsymbol{\Sigma}|\bullet \sim \mathcal{IW}(d_1, \boldsymbol{D}_1),$$

with posterior degrees of freedom $d_1 = d_0 + T/2$ and $M \times M$-dimensional posterior scaling matrix $\boldsymbol{D}_1 = \boldsymbol{D}_0 + (\boldsymbol{Y} - \boldsymbol{X}\boldsymbol{a})'(\boldsymbol{Y} - \boldsymbol{X}\boldsymbol{a})/2$.

(iv) The parameters of the state transition function are sampled jointly with a Metropolis-within-Gibbs step. The candidate values are sampled from independent Gaussian proposal densities: $\mu^{(*)} \sim \mathcal{N}(\mu^{(s)}, \kappa_\mu)$ and $\rho^{(*)} \sim \mathcal{N}(\rho^{(s)}, \kappa_\rho)$, with the candidate value $\mu^{(*)}$ $(\rho^{(*)})$ being centered on the last accepted draw $\mu^{(s)}$ $(\rho^{(s)})$ and featuring a variance κ_μ (κ_ρ). The acceptance probability of the proposed values is then given by:

$$\min\left(1, \frac{\mathcal{L}(\mu^{(*)}, \rho^{(*)}|\bullet)p(\mu^{(*)}, \rho^{(*)})}{\mathcal{L}(\mu^{(s)}, \rho^{(s)}|\bullet)p(\mu^{(s)}, \rho^{(s)})}\right),$$

with $\mathcal{L}$ denoting the data likelihood conditional on the proposed values $(\mu^{(*)}, \rho^{(*)})$ and last accepted draw $(\mu^{(s)}, \rho^{(s)})$, while $p(\mu^{(*)}, \rho^{(*)})$ and $p(\mu^{(s)}, \rho^{(s)})$ refer to joint prior distribution evaluated at the proposed values and the last accepted draw.

We repeat these steps $15,000$ times and discard the initial $5,000$ draws as burn-in.

Index of Working Papers:

July 1, 2022	Marcel Barmeier	242	The new normal: bank lending and negative interest rates in Austria
July 14, 2022	Pavel Ciaian, Andrej Cupak, Pirmin Fessler, d’Artis Kancs	243	Environmental-Social-Governance Preferences and Investments in Crypto-Assets
October 18, 2022	Burkhard Raunig, Michael Sigmund	244	The ECB Single Supervisory Mechanism: Effects on Bank Performance and Capital Requirements
April 5, 2023	Norbert Ernst, Michael Sigmund	245	Are zombie firms really contagious?
May 8, 2023	Richard Sellner, Nico Pintar, Norbert Ernst	246	Resource Misallocation and TFP Gap Development in Austria
September 5, 2023	Katharina Allinger, Fabio Rumler	247	Inflation Expectations in CESEE: The Role of Sentiment and Experiences
October 16, 2023	Pietro Saggese, Esther Segalla, Michael Sigmund, Burkhard Raunig, Felix Zangerl, Bernhard Haslhofer	248	Assessing the Solvency of Virtual Asset Service Providers: Are Current Standards Sufficient?
October 20, 2023	Pirmin Fessler, Severin Rapp	249	The subjective wealth distribution: How it arises and why it matters to inform policy?
October 27, 2023	Burkhard Raunig, Michael Sigmund	250	Watching over 21,000 Billion Euros: Does the ECB Single Supervisory Mechanism Affect Bank Competition in the Euro Area?
December 5, 2023	Markus Knell	251	Housing and the secular decline in real interest rates
December 14, 2023	Niko Hauzenberger, Florian Huber, Thomas O. Zörner	252	Hawks vs. Doves: ECB’s Monetary Policy in Light of the Fed’s Policy Stance