International Effects of Euro Area Forward Guidance

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In this paper, we investigate the international effects of euro rate forward guidance (FG) and compare them to spillovers from a conventional monetary policy shock (MP). We identify the forward guidance shock via a combination of zero and sign-restrictions that use the relationship between expectations and observed data and additionally draw on insights from recent event studies using high frequency data. To address potential time-variation, we use a fully flexible approach that allows to handle both drifts in residual variances and the structural coefficients. Our results show that both shocks lead to considerable international effects on output growth, inflation and equity returns. Moreover, we find that effects are stronger during the period of the global financial crisis, which is particularly true for the FG shock. This implies that monetary policy is generally not hindered in affecting the real economy by the zero lower bound. Also, shocks to expectations can have real domestic effects with international consequences.

JEL CODES: E5, F3, C11, C30

KEYWORDS: Spillovers, interest rate expectations, forward guidance, GVAR

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1. INTRODUCTION: MOTIVATION AND BACKGROUND

Best practice in the conduct of monetary policy requires a central bank to be forward-looking. Ordinarily, this means providing markets and the public with information about the economic outlook. Even if, as Michael Woodford points out in his seminal book, *Interest and Prices* (Woodford, 2003), the outlook contains an element of history dependence, transparency puts pressure on the central bank to guide markets about the likely future path of interest rates, so-called rate forward guidance (FG). Consequently, FG operates through a compression of the expectations component of the term structure. Successful rate FG might also reduce uncertainty about future interest rates thereby depressing the term premium. Thus, the transmission generally works in a manner that is reminiscent of conventional monetary policy. The difference is that FG appears to impact the middle segment of the yield curve (Brand et al., 2010; Rogers et al., 2014; Altavilla et al., 2019; Kim et al., 2020) and potentially leads to a larger compression of the term premium. Through both channels, FG affects asset prices and exchange rates (Altavilla et al., 2019; Bank for International Settlements, 2019) which spillover abroad.

While the literature dealing with domestic effects of rate FG is fast growing (see, e.g., Andrade and Ferroni, 2020; Lakdawala, 2019; Bundick and Smith, 2020b; Kim et al., 2020; Kortela and Nelimarkka, 2020), potential effects on neighboring countries have been overlooked so far. To do so requires a global model to investigate the macroeconomic effects of rate forward guidance in the euro area. As a benchmark, and to better isolate the effects of rate FG, we will compare them with the effects from the implementation of a conventional monetary policy. Recent studies suggest that the effectiveness of monetary policy is different during periods of heightened financial stress (Jannsen et al., 2019), when interest rates reached the zero lower bound (ZLB, Eggertsson and Woodford, 2003; Liu et al., 2019; Ikeda et al., 2015) or the yield curve has flattened. To address potential time-variation of the effects, we estimate the fully flexible global time-varying parameter vector autoregressive model of Crespo Cuaresma et al. (2019). We also expand considerably the number of countries and regions of the world investigated compared to Jannsen et al. (2019) so that we can empirically determine whether spillover effects amplify or not domestic responses to a FG shock.

A precise quantitative assessment of FG, though, is impeded by several challenges which go beyond those associated with the identification of the impact of conventional monetary policy.[1] Foremost, an empirical investigation is plagued by an inherent endogeneity that arises because current policy decisions are based on a mix of history dependent factors, as well as forecasts of future economic conditions. Hence, identification is necessary to disentangle in observed data promises to act in future from the central bank's communication of the outlook to set the current stance of monetary policy. This is the inevitable consequence of a monetary policy that is forward-looking, as best practice requires. As a result, a plethora of identifying restrictions have been proposed to enable researchers to make inferences about the respective roles of past performance versus policy makers' outlook and how they intend to respond to them. Another challenge is that there are different types of FG, such as calendar-based or state-dependent. Ehrmann et al. (2019) show that long-horizon calendar-based FG and state-dependent

[1]One of series of complications, but one outside the scope of this paper, is that FG has often accompanied QE interventions (e.g., see Lombardi et al., 2018, and references therein). There is a reasonable consensus that QE has been successful in reducing long-term yields. (inter alia, see Swanson and Williams, 2014; Altavilla et al., 2019; Lombardi et al., 2019).

FG are very effective forms of FG. Calendar-based FG with a short horizon can have side effects, while purely qualitative FG is largely ineffective but does not do any harm.

Another and more fundamental issue is to separate Odyssean from Delphic forward guidance. Delphic – or news-driven – forward guidance is determined by the central bank's own interest rate projections. By contrast, Odyssean guidance represents an explicit commitment to a particular future interest rate path.[2] The implications of these two types for the effectiveness of monetary policy of FG differ considerably: while output and inflation under Odyssean FG behave much like the response to a conventional monetary policy shock, Delphic FG predicts the opposite. The reason is that accommodative FG could be perceived by financial markets as the central bank signalling bad economic news in future as it attempts to forestall this state by promising, if necessary, to maintain an easing of monetary policy. Note that the reasons why markets perceive FG as Odyssean or Delphic are unclear and do not depend on the type of FG. However, there is a tendency for news-based FG to be interpreted as Delphic as indicated in a recent study by Lunsford (2020). In the euro area, since the launch of explicit use of rate forward guidance in 2013, the majority of FG shocks have been interpreted as being Odyssean in nature (Altavilla et al., 2019; Andrade and Ferroni, 2020), which is why our study focuses on this type of FG interpretation. Moreover, our empirical analysis coincides with the time period when overnight rates were hitting the ZLB limiting the possibilities of monetary policy on the short end of the yield curve.[3] A number of prominent studies have emerged that investigate, both theoretically and empirically, the effectiveness of monetary policy at the ZLB (Eggertsson and Woodford, 2003; Liu et al., 2019; Ikeda et al., 2015; Debortoli et al., 2019). The global VAR time-varying estimation approach adopted in this paper permits us to determine with greater precision whether the relative importance of the Odyssean-Delphic distinction versus identifying the response to FG as between calmer times and crisis periods is warranted. Indeed, time-varying estimation also permits us to conclude that the ZLB need not limit the effectiveness of monetary policy. The broader economic environment at the time also matters. Others have reached the same conclusion (e.g., Debortoli et al., 2019).

Against this backdrop, it is not surprising that the literature on the macroeconomic effects of FG has arrived at mixed conclusions. One strand of the literature exploits high frequency financial data (i.e., yields at different maturities, exchange rates, etc.) shortly before and after a FG announcement (Campbell et al., 2012; Gürkaynak et al., 2015; Altavilla et al., 2019; Bundick and Smith, 2020a; Jarociński and Karadi, 2019; Nakamura and Steinsson, 2018). Specifying a narrow time window is one strategy to prevent other news events from contaminating the source of any changes in, say, market interest rates and the wider public response to potential changes in the policy stance. A drawback of high frequency identification is that it can reveal effects that are not consistent with standard economic theory. For example, Miranda-Agrippino and Rey (2020) show that monetary policy shocks identified with high frequency data often lead to output and / or price puzzles. Campbell et al. (2012) find that a FG tightening lowers the unemployment rate and increases inflation for the US economy. Bundick and Smith (2020a) find the opposite, namely a gradual decline in economic activity accompanied by a decrease in investment and capital utilization

[2] As an aside, it is often forgotten that the commitment is either explicitly, and almost certainly, conditional on how future economic data evolve or what has been termed data dependence.

[3] At the outset, the ZLB was intended to communicate the inability or unwillingness of central banks to allow interest rates to be negative, as in the UK and the US. Since then several economies have breached the ZLB, including the ECB, giving rise to the effective lower bound. For the sake of consistency we continue to use ZLB to refer to both conditions.

and modest disinflation. A second strand of the literature combines high frequency identification with data on expectations. Surveys of professional forecasters are often used and relate to expectations on output, inflation and interest rates.[4] Andrade and Ferroni (2020) disentangle Delphic from Odyssean FG shocks by assuming that the former is accompanied by an upward movement in inflation expectations whereas the opposite is true for FG shocks of the Odyssean type. This assumption is based on a small scale theoretical model and empirically corroborated by de la Barrera et al. (2017) using an event study approach. Aksit (2020b) shows that FG has elements of both the Delphic and Odyseean types while Aksit (2020a) sees FG as a succession of policy shocks into the future. This can explain away the so-called FG puzzle wherein the impact of FG shocks is larger than would be expected from a standard New Keynesian macro model (Del Negro et al., 2015). This paper cannot address all of the questions just raised, but we retain the use of a time series approach in preference to the event study methodology as well as the important role played by market expectations. Moreover, we explicitly control for the communication of the ECB. To our knowledge, this has not been done so far in the present context. More importantly, as noted above, the literature on international effects on forward guidance is rather limited. Theoretically, Callum et al. (2018) set up a two-country DSGE model to capture the transmission of FG shocks from the USA to Canada. Their work emphasizes the importance of time variation in FG spillovers: while, in general, the effects of FG on Canadian output are twice as large as those from conventional monetary policy, they decrease significantly if aggregate demand in the US economy is weak. Inoue and Rossi (2019) focus on the effects of unconventional monetary policy including FG on international exchange rates. They do not find significant differences between conventional and unconventional monetary policy shocks on bilateral exchange rates. D'Amico and King (2015) identify interest rate expectation shocks for the US economy using a novel combination of zero and sign restrictions (Arias et al., 2018). They find that a decrease in expected interest rates triggers large, immediate, and persistent upward effects on inflation and real activity. These effects tend to rise with the expectations' horizon. We retain the essence of their identification strategy but make a substantive modification to it by relaxing how inflation and output expectations adjust to observed values of these same variables.

Specifically, we investigate the macroeconomic effects of rate forward guidance in the euro area. We do this by augmenting our empirical model with expectations data and using a combination of zero and sign restrictions advocated in D'Amico and King (2015). These "rationality" conditions ensure that responses of observed and expectations data do not diverge by the end of the forecast horizon. Therefore, since restrictions are not enforced in the interim periods, rationality is not imposed at all times. Forecasts can deviate from observed values in the short-run. Our main results show that both shocks generate a broad based decline in economic activity and inflation as well as global equity returns. Neighboring countries' currencies tend to weaken, but here results are more country-specific. Taken at face value, our results show that FG can have significant spillover effects beyond ones that impact domestic economies.

[4]Since all of the relevant studies focus on financial market reactions to various forms of central bank intervention it is natural to resort to the forecasts of professionals. The other advantage is that the data are more readily available and usually at a higher sampling frequency than household forecasts. As a referee points out, it is well known that households' expectations behave differently from those of other groups. See, for example, Coibion and Gorodnichenko (2012). D'Acunto et al. (2020), using household level data from Europe find that "... only experts with extremely high levels of sophistication, such as experts, react to forward guidance." (op.cit, p. 6-7). The precise role played by different kinds of expectations in the present context is a topic left for future research.

Last, we find evidence of considerable time variation in the size of spillovers. Both shocks trigger larger effects on international output growth during the period of the global financial crisis compared to the rest of the sample. On top of that, FG seems also able to trigger large effects on inflation during that period rendering it a particularly attractive tool for the policymaker in steering inflation. By contrast, we do not find significantly different effects of both shocks when the ZLB is binding.

The paper is structured as follows. The next section introduces the data and Sec. 3 the Bayesian global VAR model with time-varing parameters and stochastic volatility (TTVP-SV-GVAR). Section 4 describes the empirical results and Sec. 5 concludes.

2. DATA AND SOME STYLIZED FACTS

We use monthly data spanning the period from 2001m01 to 2018m06. For most economies, we have collected data on short-term interest rates (3-months money market rates, i^s_{it}), consumer price inflation (Dp_{it}, year-on-year growth rate of index 2010=100), long-term interest rates (i^l_{it}, 10-year government bond yields), monthly stock market returns (eq_{it}), and the nominal exchange rate vis-à-vis the euro (er_{it}). For the latter an increase implies a depreciation of the local currency against the euro. As a measure of economic activity, we use real GDP growth (y_{it}, in year-on-year terms), disaggregated to the monthly frequency by using the Chow-Lin method (Chow and Lin, 1971) and industrial production alone as a high frequency indicator. This yields somewhat less volatile growth rates than using industrial production as an economic activity indicator. Data on industrial production and consumer prices are de-seasonalized.

Data are collected for a broad set of countries, namely the euro area (EA), countries from Central, Eastern and Southeastern Europe (*CESEE*: Bulgaria, Croatia, Czechia, Hungary, Poland, Romania, Russia and Turkey) and other advanced European countries (*Europe*: Denmark, Great Britain, Norway, Sweden and Switzerland). We include Russia and Turkey due to their close economic ties with the euro area as well as the USA, China, Canada and Japan (Rest of the world, *RoW*) to control for global factors. That leaves us with a sample of good coverage of European and G-8 industrialized advanced economies. The countries in our sample account for about 70% of global nominal output (averaged over the years 2010 to 2016) and reflect the most important trading partners of the euro area.

For a typical non-euro area country i, this yields a $k_i \times 1$ vector of endogenous variables, $\boldsymbol{x}_{it}$:

$$\boldsymbol{x}_{it} = (y_{it}, Dp_{it}, i^s_{it}, i^l_{it}, eq_{it}, er_{it})'. \tag{1}$$

The euro area model differs from the remaining countries in terms of variable coverage. To identify the forward guidance shock, we require expectations data. More specifically, we include one-year ahead forecasts of real GDP growth (y^{t+12}_{it}), inflation (Dp^{t+12}_{it}) and short-term interest rates ($i^{s,t+12}_{it}$). The data are from Consensus economics and are transformed to a fixed-horizon forecast as in Siklos (2013).

As a control variable, we also include the Swiss Economic Institute's (KOF) Monetary Policy Communicator index (mpc^{comm}_{it}). This variable quantifies the ECB president's statements concerning risks to price stability as made during the monthly press conference into an index. The methodology is implemented by analysts from Media Tenor, a media research institute. Media analysts read the text of the introductory statement of the monthly press conference sentence by sentence and code them.[5] The coding is aggregated by the KOF Swiss Economic Institute by evaluating the balance of statements where the ECB sees upside risks to future price stability and statements versus ones that communicate downside risks to future price stability, relative to all statements about future price stability (including neutral ones). By construction, the values of the KOF MPC are restricted to be in the range of minus one to plus one with larger and positive values indicating a stronger communication of the ECB of upside risks for future price stability. It is also shown that the indicator leads the policy rate by two to three months and has a strong positive relationship with inflation expectations (Neuenkirch, 2013). We include the indicator to control for short-term interest rate expectations, and because it helps to identify a conventional monetary policy shock (Neuenkirch, 2013). Most importantly though, we want to control for Delphic FG as well as news shocks. Andrade and Ferroni (2020), de la Barrera et al. (2017) show that Delphic FG can be characterized by a positive correlation of interest rate expectations and inflation expectations and Lunsford (2020) demonstrate that news shocks produce a similar pattern. Since the KOF indicator is a proxy for interest rate expectations (Neuenkirch, 2013) and by construction positively correlated with inflation expectations it is a suitable control variable. Finally, we also include three Nelson-Siegel factors that reflect, level (β^0_{it}), slope (β^1_{it}) and curvature (β^2_{it}) of the joint euro area yield curve (Nelson and Siegel, 1987; Diebold and Li, 2006; Diebold et al., 2006). The factors have been estimated using a dynamic, arbitrage free Nelson-Siegel model and daily data on overnight indexed swaps (OIS).[6]

Note that in accordance with the literature, the slope factor is inversely defined, i.e., β^2_{it} approximates the difference between short- and long-term rates (Diebold and Li, 2006). We include the yield curve factors to further distinguish the conventional monetary policy from the FG shocks – the latter should mainly work on the middle segment of the yield curve (Brand et al., 2010; Altavilla et al., 2019) and hence affect the curvature factor.[7]

The euro area country model $i = 0$ then features the following variables:

$$\boldsymbol{x}_{0t} = (y_{0t}, Dp_{0t}, i^s_{0t}, y^{t+12}_{0t}, Dp^{t+12}_{0t}, i^{s,t+12}_{0t}, eq_{0t}, mpc^{comm}_{0t}, \beta^0_{0t}, \beta^1_{0t}, \beta^2_{0t})'. \tag{2}$$

Fig. 1 shows the variables related to monetary policy.

We see that inflation pressure, measured by the KOF indicator, evolves in cycles, i.e., there are consecutive press conferences that either suggest positive or negative pressure on inflation. This is not surprising, since the monetary policy stance and economic conditions do not change abruptly. Over the latest sub-period of the sample, values for the KOF indicator are positive. This can be interpreted as a sign that the ECB expects the unconventional monetary policy measures undertaken over this sub-period to be successful in driving up inflation. On the right-hand side of the figure, we plot expectations of real

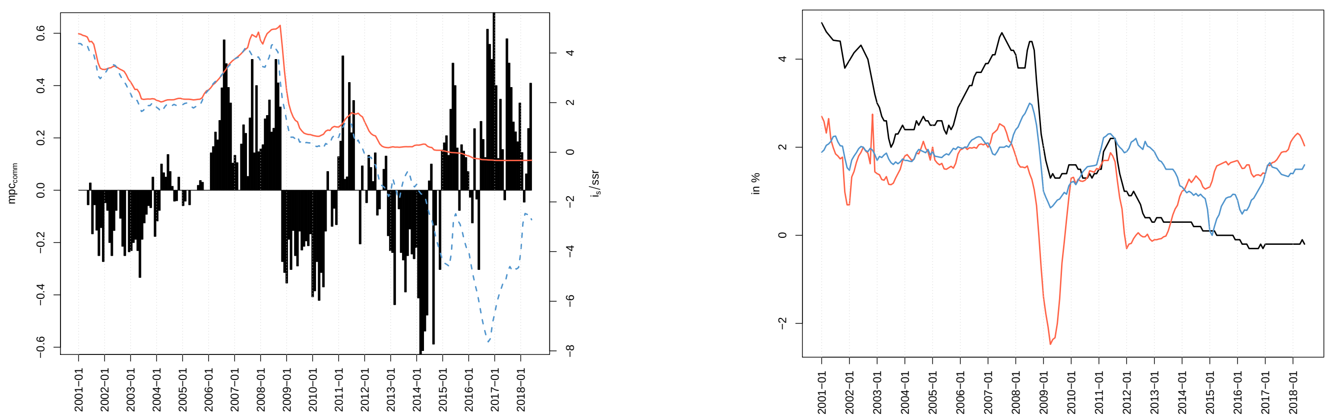
[5] Media analysts read the text of the introductory statement of the monthly press conference sentence by sentence and code them. The coding is available at `https://www.kof.ethz.ch/en/forecasts-and-indicators/indicators/kof-monetary-policy-communicator.html`.

[6] We thank Tomi Kortela for sharing his estimates with us. Prior to 2006 and due to lack of availability, OIS rates are approximated by the mean of German and French government bond yields. For more details, see Kortela and Nelimarkka (2020) and Altavilla et al. (2019) for a similar approximation of euro area yields.

[7] Conversely, Hubert and Labondance (2018), using an event study approach, find that forward guidance has a significant effect on the whole term structure of interest rates – also on longer term yields.

Fig. 1: Euro area monetary policy variables - data overview

(a) mpc^{comm} / i^s (b) Expectations



Notes: The figure on the left-hand side shows the KOF MPC communicator (mpc^{comm}, black solid line) index on the left-hand scale, and the euro area short term interest rates (i^s, red solid line) and the shadow rate of Krippner (ssr, blue dashed line Krippner, 2013) on the right-hand scale. The right-hand side figure shows one-year head survey expectations of real GDP growth (red, y^{t+12}), inflation (Dp^{t+12}, blue) and short-term interest rates ($i^{s,t+12}$, black).

GDP growth, inflation and short-term interest rates, transformed to fixed-event, one-year ahead forecasts. We observe positive expectations of economic activity and inflation over the most recent sample period.

Tab. 1 displays simple correlations of the monetary policy and expectations variables over the pre- and post-crisis periods.

The table shows that interest rates, actual and expectations and the shadow rate, all were closely related to each other and positively correlated in the pre-crisis period. The KOF communicator index is also positively correlated with the remaining variables (most strongly so with expected inflation). If we now look at the post-crisis period we see that correlations change, from positive to negative. More specifically, both the KOF communicator index and real GDP growth expectations are now negatively related with interest rates and interest rate expectations. Also, the association between output growth expectations and the monetary policy variables differs across the two sub-periods. Frequent changes of the macroeconomic effects of FG have also been reported in Andrade and Ferroni (2020). This motivates resort to a time-varying parameter framework to properly assess the effects of FG in the euro area.

3. ECONOMETRIC FRAMEWORK

To assess spillovers of the euro area shocks, and to account for potential changes in these effects over time, we estimate a global vector autoregressive model with time-varying parameters and stochastic volatility (TVP-SV-GVAR model), recently proposed in Crespo Cuaresma et al. (2019)

In general, the structure of a GVAR model implies two distinct stages in the estimation process. In the first stage, $N + 1$ country-specific multivariate time series models are specified, each of them including weakly exogenous regressors that aim to capture cross-country linkages. It is in this stage that we introduce time variation by using the mixture innovation specification of Huber et al. (2019). In the second stage,

Tab. 1: Correlations of monetary policy variables and expectations

2001m1-2008m12	y^{t+12}	Dp^{t+12}	$i^{s,t+12}$	i^s	ssr	mpc^{comm}
y^{t+12}	1					
Dp^{t+12}	0.11	1				
$i^{s,t+12}$	0.40	0.50	1			
i^s	0.17	0.69	0.88	1		
ssr	0.42	0.59	0.97	0.92	1	
mpc^{comm}	0.35	0.62	0.30	0.39	0.42	1
2009m1-2018m6	y^{t+12}	Dp^{t+12}	$i^{s,t+12}$	i^s	ssr	mpc^{comm}
y^{t+12}	1					
Dp^{t+12}	0.02	1				
$i^{s,t+12}$	-0.38	0.38	1			
i^s	-0.53	0.33	0.93	1		
ssr	-0.47	0.43	0.88	0.85	1	
mpc^{comm}	0.42	0.13	-0.30	-0.32	-0.47	1

these models are combined using country weights to form a global model which then forms the basis for the impulse response analysis.

3.1. *A Dynamic Global Macroeconomic Model*

Let the endogenous variables for country $i = 0, \dots, N$ be contained in a $k_i \times 1$ vector $\boldsymbol{x}_{it} = (x_{i1,t}, \dots, x_{ik_i,t})'$. In addition, all country-specific models feature a set of k_i^* weakly exogenous regressors $\boldsymbol{x}_{it}^* = (x_{i1,t}^*, \dots, x_{ik_i,t}^*)'$ constructed as weighted averages of the endogenous variables in other economies,

$$x_{ij,t}^* = \sum_{c=0}^{N} w_{ic} x_{cj,t} \text{ for } j = 1, \dots, k_i^*, \tag{3}$$

where w_{ic} is the weight corresponding to the jth variable of country c in country i's specification. These weights are assumed to be related to bilateral trade exposure[8], sum up to unity and only off-diagonal elements are non-zero ($\sum_{c=0}^{N} w_{ic} = 1$ and $w_{ii} = 0$). In line with the bulk of the literature on GVAR modeling, we assume that all variables and countries are linked by the same set of weights which is fixed over time (Dees et al., 2007). We follow Crespo Cuaresma et al. (2019) and specify country-specific structural VARX*(1,1) models with time-varying parameters and stochastic volatility:

$$\boldsymbol{Q}_{i0,t}\boldsymbol{x}_{it} = \boldsymbol{A}_{i1,t}\boldsymbol{x}_{it-1} + \boldsymbol{B}_{i0,t}\boldsymbol{x}_{it}^* + \boldsymbol{B}_{i1,t}\boldsymbol{x}_{it-1}^* + \boldsymbol{\varepsilon}_{it} \quad \boldsymbol{\varepsilon}_{it} \sim \mathcal{N}(0, \boldsymbol{D}_{it}) \tag{4}$$

[8]More precisely, we use annual data from the World Input Output Database (WIOD), averaged over the period from from 2000 to 2014. The data were retrieved from `http://www.wiod.org/home` and only available up until 2014. For a detailed description see Timmer et al. (2015).

with $\boldsymbol{Q}_{i0,t}$ denoting a $k_i \times k_i$ matrix of structural coefficients $\boldsymbol{A}_{i1,t}$ is a $k_i \times k_i$ matrix of coefficients associated with the lagged endogenous variables and $\boldsymbol{B}_{iq,t}$ $(q = 0, \ldots, 1)$ denote $k_i \times k_i^*$ dimensional coefficient matrices corresponding to the k_i^* weakly exogenous variables in $\boldsymbol{x}_{it}^*$. The matrix of structural coefficients, $\boldsymbol{Q}_{i0,t}$, is a lower triangular matrix with a diagonal of ones and captures the contemporaneous relationships between the variables in $\boldsymbol{x}_{it}$. Estimating the structural form of the model (i.e., with the errors being orthogonal) allows us to use equation-by-equation estimation which considerably speeds up computational time (for a recent similar approach, see Crespo Cuaresma et al., 2019; Huber et al., 2019; Jannsen et al., 2019; Koop et al., 2019).[9] For the error term, we assume that $\boldsymbol{D}_{it} = \text{diag}(\lambda_{i0,t}, \ldots, \lambda_{ik_i,t})$. That is, while the off-diagonal elements within the country model are zero (the relationships already captured by $\boldsymbol{Q}_{i0,t}$), the variances of the residuals of the model are allowed to vary over time. Stochastic volatility is now considered as important building block in time series models to capture possibly large shocks leading to better predictive properties (Clark, 2011). These are assumed to follow a stationary autoregressive process,

$$\log(\lambda_{il,t}) = \mu_{il} + \rho_{il}(\log(\lambda_{il,t-1}) - \mu_{il}) + \upsilon_{il,t},\ \upsilon_{il,t} \sim \mathcal{N}(0, \varsigma_{il}^2), \tag{5}$$

where μ_{il} denotes the unconditional expectation of the log-volatility, ρ_{il} the corresponding persistence parameter and ς_{il}^2 is the innovation variance of the process. Allowing for stochastic volatility ensures that changes in the parameters reflect changes in the underlying macroeconomic relationships and are not confounded by a wrongly assumed constant error variance. Also, with reference to the following empirical application that features expectations data for the euro area, having a specification that allows for heteroskedasticity, seems essential (Galati et al., 2011; de la Barrera et al., 2017).

In the second stage of the GVAR framework, the coefficients of the single country models can be stacked to yield a global vector autoregressive model. For that purpose, we rewrite the $N + 1$ country specific models in terms of a global vector $\boldsymbol{x}_t = (\boldsymbol{x}_{0t}', \ldots, \boldsymbol{x}_{Nt}')'$ of dimension $k = \sum_{i=0}^{N} k_i$ (Pesaran et al., 2004). More specifically, we define country-specific link matrices $\boldsymbol{W}_i$ $(i = 1, \ldots, N)$ and selection matrices $\boldsymbol{S}_i$, both of dimension $(k_i + k_i^*) \times k$. The selection matrix is composed of zero elements, with unit values only on the elements that singles out the variables for country i from the global vector $\boldsymbol{x}_t$.

$$\begin{aligned}
\boldsymbol{Q}_{i0,t}\boldsymbol{S}_i\boldsymbol{x}_t &= \boldsymbol{A}_{i1,t}\boldsymbol{S}_i\boldsymbol{x}_{t-1} + \boldsymbol{B}_{i0,t}\boldsymbol{W}_i\boldsymbol{x}_t + \boldsymbol{B}_{i1,t}\boldsymbol{W}_i\boldsymbol{x}_{t-1} + \boldsymbol{\varepsilon}_{it}\\
(\boldsymbol{Q}_{i0,t}\boldsymbol{S}_i - \boldsymbol{B}_{i0,t}\boldsymbol{W}_i)\boldsymbol{x}_t &= (\boldsymbol{A}_{i1,t}\boldsymbol{S}_i + \boldsymbol{B}_{i1,t}\boldsymbol{W}_i)\boldsymbol{x}_{t-1} + \boldsymbol{\varepsilon}_{it}\\
\boldsymbol{G}_{i,t}\boldsymbol{x}_t &= \boldsymbol{H}_{it}\boldsymbol{x}_{t-1} + \boldsymbol{\varepsilon}_{it}
\end{aligned}$$

with $\boldsymbol{G}_{i,t} = (\boldsymbol{Q}_{i0,t}\boldsymbol{S}_i - \boldsymbol{B}_{i0,t}\boldsymbol{W}_i)$ and $\boldsymbol{H}_{it} = \boldsymbol{A}_{i1,t}\boldsymbol{S}_i + \boldsymbol{B}_{i1,t}\boldsymbol{W}_i$. Stacking the equations $N + 1$ times yields the global representation of the model

$$\boldsymbol{G}_t\boldsymbol{x}_t = \boldsymbol{H}_t\boldsymbol{x}_{\boldsymbol{t-1}} + \boldsymbol{\varepsilon}_t, \quad \boldsymbol{\varepsilon}_t \sim N(0, \boldsymbol{\Sigma}_t) \tag{6}$$

[9] Introducing the lower triangular matrix renders our estimation as potentially dependent on the ordering of the variables. However, since we are ultimately interested in impulse responses identified using an order-invariant identification scheme – zero and sign restrictions – our results turn out not to be sensitive to the arrangement of the variables in the model.

Eq. (6) resembles a large VAR model with drifting coefficients and stochastic volatility, which we are going to estimate using Bayesian techniques. Following Huber (2016), we assume a block-diagonal $\mathbf{\Sigma}_t$. This is a stricter version of the implicit assumption in the GVAR framework that cross-country reduced form residual correlation is weak, which can be indirectly tested by examining the pairwise cross-country correlation of the country models' residuals (see the appendix, Fig. B.1). Note, however, that block-diagonality of $\mathbf{\Sigma}_t$ does not restrict immediate cross-country spillovers to be zero, since pre-multiplying Eq. (6) with $\boldsymbol{G}_t^{-1}$ establishes contemporaneous links as a function of the underlying weights.

3.2. *Modeling Time Variation in the Regression Coefficients*

To model and estimate time variation in the coefficients, we use the model of Huber et al. (2019) that has been recently applied in the GVAR context by Crespo Cuaresma et al. (2019). In a nutshell, this model estimates a latent threshold for each coefficient of the model. Small movements of the coefficients that do not surpass the threshold (in absolute values) are deemed negligible and the corresponding coefficient held constant. This specification hence allows for a parsimonious way of modeling time-variation, which is necessary in order to obtain precise estimates and avoid overfitting.

Stacking the lagged endogenous and weakly exogenous variables in an m_i-dimensional vector, with $m_i = k_i + k_i^*(1+1)$,

$$z_{it} = (\boldsymbol{x}'_{it-1}, \boldsymbol{x}^{*'}_{it}, \boldsymbol{x}^{*'}_{it-1})' \tag{7}$$

and collecting all coefficients in a $m_i \times k_i$ matrix $\mathbf{\Psi}_{it} = (\boldsymbol{A}_{i1,t}, \boldsymbol{B}_{i0,t}, \boldsymbol{B}_{i1,t})'$ allows us to rewrite Eq. (4) as

$$\boldsymbol{Q}_{i0,t}\boldsymbol{x}_{it} = (\boldsymbol{I}_{k_i} \otimes z'_{it})\, \mathbf{\Psi}_{it} + \varepsilon_{it}. \tag{8}$$

We let $\boldsymbol{\psi}_{it} = \text{vec}(\mathbf{\Psi}_{it})$ and collect the free elements elements of $\boldsymbol{Q}_{i0,t}$ in a $l_i = k_i(k_i-1)/2$-dimensional vector $\boldsymbol{a}_{i0,t}$.

For each individual coefficient in $\boldsymbol{\xi}_i = (\boldsymbol{q}'_{i0,t}, \boldsymbol{\psi}'_{it})'$, we assume a random walk law of motion,

$$\xi_{ij,t} = \xi_{ij,t-1} + \eta_{ij,t}, \text{ for } j = 1, \dots, s_i, \tag{9}$$

with $s_i = l_i + k_i(m_i k_i)$ and $\eta_{ij,t}$ being a white noise shock with time-varying variance $\vartheta_{ij,t}$. Here we follow Huber et al. (2019) and assume that $\vartheta_{ij,t}$ evolves according to

$$\vartheta_{ij,t} = (1 - d_{ij,t})\vartheta_{ij,0} + d_{ij,t}\vartheta_{ij,1}, \tag{10}$$

whereby $\vartheta_{ij,1} \gg \vartheta_{ij,0}$ and $\vartheta_{ij,0}$ is set close to zero. Moreover, let $d_{ij,t}$ denote a binary random variable that follows an independent Bernoulli distribution with,

$$d_{ij,t} = \begin{cases} 1 & \text{with probability} \quad p_{ij} \\ 0 & \text{with probability} \quad 1 - p_{ij}. \end{cases} \tag{11}$$

This specification captures the notion that if the period-on-period change in the respective parameter is large, the unconditional probability (i.e., with $\xi_{ij,t}$ integrated out) of $d_{ij,t} = 1$ is also large.[10] It also nests

[10]Due to the high computational burden, the latent indicators ($d_{ij,t}$) are approximated rather than exactly estimated during the Markov chain Monte Carlo sampling (Huber et al., 2019).

a wide variety of competing models. For instance, if $d_{ij,t} = 1$ for all t we obtain a standard time-varying parameter specification whereas in the case of $d_{ij,t} = 0$ for all t, we end up having a nearly constant parameter specification (as the variance of $\eta_{ij,t}$ will be relatively small). Cases in between are also possible, implying that our framework accommodates situations where parameters might be time-varying during certain points in time and effectively constant during other periods. The degree of time variation is determined by the threshold for each coefficient separately, which is a strong feature of the model and allows for a great deal of flexibility.

We estimate the above model using Bayesian techniques using 30,000 draws as burn-ins and 10,000 posterior draws for inference. While the latent threshold specification ensures a parsimonious handling of time-variation in the coefficients, we augment it with Bayesian shrinkage priors on the initial state of each coefficient. This facilitates variable selection and should mitigate the risk of overparameterization that is inherent vector autoregressive models. The exact prior specifications are detailed in A in the appendix section.

3.3. *Structural Identification*

In this section, we outline the empirical strategy to identify euro area monetary policy and FG shocks. For that purpose, we use a combination of zero and sign restrictions and locally identify the shocks in the euro area country model. In the GVAR context, local identification has been proposed in Dees et al. (2007) and applied – among others – in Eickmeier and Ng (2015) and Feldkircher and Huber (2016).

Recall that the euro area country model is indexed by $i = 0$ and that we estimate already a structural (Cholesky-type) form of the model. Sign restrictions are then introduced by drawing a $k_0 \times k_0$ rotation matrix $\boldsymbol{R}_{0t}$ (with $\boldsymbol{R}_{0t}\boldsymbol{R}'_{0t} = \boldsymbol{I}_{k_0}$) using the algorithm of Arias et al. (2018).

We pre-multiply Eq. (4) to get

$$\tilde{\boldsymbol{Q}}_{00,t}\boldsymbol{y}_{it} = \tilde{\boldsymbol{A}}_{0p,t}\boldsymbol{x}_{0,t-1} + \tilde{\boldsymbol{B}}_{00,t}\boldsymbol{x}^*_{0,t} + \tilde{\boldsymbol{B}}_{01,t}\boldsymbol{x}^*_{0,t-1} + \tilde{\boldsymbol{R}}_{i0,t}\boldsymbol{\varepsilon}_{it}, \tag{12}$$

with $\tilde{\boldsymbol{Q}}_{00,t} = \boldsymbol{R}_{0t}\boldsymbol{Q}_{00,t}$, $\tilde{\boldsymbol{A}}_{01,t} = \boldsymbol{R}_{0t}\boldsymbol{A}_{01,t}$ and $\tilde{\boldsymbol{B}}_{0q,t} = \boldsymbol{R}_{0t}\boldsymbol{B}_{0q,t}$. Notice that Eq. (12) is observationally equivalent to Eq. (4) and the introduction of the rotation matrix R_{0t} leaves the likelihood function untouched. In addition, notice that the rotation matrices are time-varying and we thus need to simulate T rotation matrices in order to identify the model for each time point. The main implication of local identification in the GVAR context is that the obtained responses are only orthogonal in the country where the shock is applied (i.e., the euro area). For all other countries, the responses are of the generalized form, as in Dees et al. (2007). This implies that the international responses capture the direct spillover effect, effects via third-countries and the domestic responses to these effects.

We identify the two shocks based on the restrictions outlined in Tab. 2 below:

First, an unexpected increase in euro area interest rates (i^s) is supposed to decrease output growth (y), consumer price inflation (Dp) and equity returns (eq). These restrictions imply that we rule out counterintuitive reactions of prices and output by construction.[11] Note that in general, the inclusion

[11] The early literature on monetary policy shocks often focused on the US economy and recursive identification (e.g., Christiano et al., 2005). More recently, a number of authors propose the use of external instruments, high frequency information (Gertler and Karadi, 2015; Altavilla et al., 2019). Miranda-Agrippino and Ricco (2015), however show that using these measures often lead to output and / or price puzzles.

Tab. 2: Sign restrictions.

Shock	$i^s(MP)$	$i^{s,t+12}(FG)$
i_s	$\pmb{\uparrow_{1,2}}$	0_1, $IRF(i^s, 13) = IRF(i^{s,t+12}, 1)$
y	$\downarrow_{1,2}$	$IRF(y, 13) = IRF(y^{t+12}, 1)$
Dp	$\downarrow_{1,2}$	$IRF(Dp, 13) = IRF(Dp^{t+12}, 1)$
eq	$\downarrow_{1,2}$	$\downarrow_{1,2}$
$i^{s,t+12}$	–	$\pmb{\uparrow_{1,2}}$
y^{t+12}	–	$\downarrow_{1,2}$
Dp^{t+12}	–	$\downarrow_{1,2}$

Notes: The sign restrictions are imposed as ≥ / ≤ and on impact and the following period. The zero restriction is imposed on impact only. Rotation matrices drawn following Arias et al. (2018).

of expectations data should mitigate the potential of price puzzles (Castelnuovo and Surico, 2010) and introduces additional information to the analysis otherwise not contained in standard macroeconomic data (D'Amico and King, 2015).[12] The assumption on equity returns is based on empirical evidence for the reaction of stock markets to monetary policy-induced interest rate changes (Thorbecke, 1997; Rigobon and Sack, 2004; Bernanke and Kuttner, 2005; Bohl et al., 2017; Li et al., 2010). Also, in a recent paper, Jarociński and Karadi (2019) use the movements of equity prices to disentangle "pure" monetary policy shocks from "information" shocks that arise from the central bank's statements about the policy decision. All assumptions are imposed on impact and the following period. The short-horizon of the restrictions and the fact that we use differenced as opposed to level data, ensures that we use a minimum of assumptions and our results are largely data driven. We do not impose restrictions on expectations, nor the Nelson-Siegel yield curve factors. We expect that the conventional MP shock mainly affects the slope factor (β^1) of the yield curve.

Second, a FG tightening is modeled as an increase in *expected* interest rates, while actual rates are unchanged. D'Amico and King (2015) lay out a New Keynesian model augmented with rational expectations and demonstrate that a shock to interest rate expectations leads to an opposite movement of expected inflation and expected output for Odyssean FG. They state that "Intuitively, expectations of future inflation, output, and interest rates are affected by expectations of a future monetary-policy change in the same way that current inflation, output, and interest rates are affected by a current unanticipated policy shock". Recent empirical evidence support the predictions of the model of D'Amico and King (2015). More specifically, Bundick and Smith (2020a) use high frequency identification of FG shocks for the USA. They show that expansionary FG shocks that lead to lower future policy rates, indeed trigger (moderate) increases in output and inflation. D'Amico and King (2015) also show how to link responses of expectations and observed data in a VAR framework by proposing to use "rationality conditions". The rationality conditions restrict responses of *observed* output growth, inflation and interest rates to match

[12]Recently, Bu et al. (2020) made the case that to capture forward looking monetary policy it is best to include central bankers' forecasts as opposed to market based expectations for the USA.

the *impact response of their expectations counterparts* at the one-year ahead forecast horizon.[13] The same applies to inflation and output growth. Note that these rationality conditions can be considered as weak assumptions. They do not rule out price / output puzzles in the short-run as output and inflation reactions in response to the FG shock can fluctuate freely up to 12 observations of the forecast horizon. Only at the forecast horizon (one year) they have to coincide. One might argue that the rationality conditions impose some "implicit" restrictions before the constraint binds. While we cannot rule out that responses at impulse response horizon 0 to 12 are somehow affected, examining the empirical results in the following section will allow us at least qualitatively to gauge how much of this is in fact a relevant issue given the data and econometric framework we use in this paper. To be consistent with the monetary policy shock, we again pursue a weak identification strategy and abstain from placing further restrictions on, say the yield curve (see, e.g., Andrade and Ferroni, 2020). Leaving the yield curve parameters free also allows us to investigate differences that arise due to the two shocks. We expect that the forward guidance shock mainly affects the curvature factor (β^2) of the yield curve.

4. EMPIRICAL RESULTS

The model presented in Sec. 3 is estimated using mean-standardized data. The results below consider the responses of a standard deviation shock in interest rates and interest rate expectations in terms of standard deviations of the variable under scrutiny.

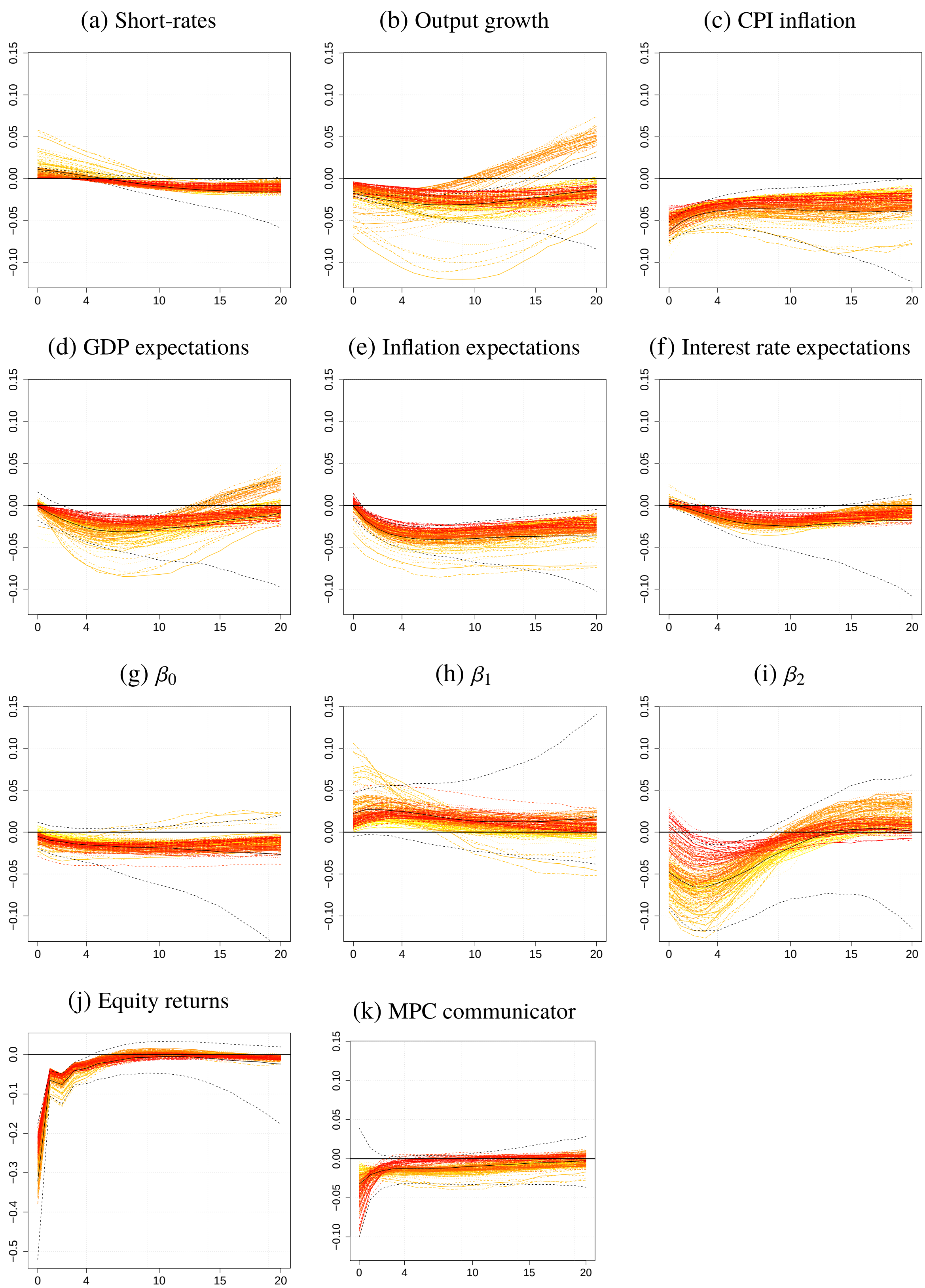
4.1. *Domestic Effects*

We begin by focusing on the domestic, euro area effects. These are depicted in Fig. 2 and Fig. 3. In each plot we show the posterior median response over the sample considered. Lighter, yellow responses refer to the beginning of the sample, darker, red responses to the end of the sample. A continuous increase in the shade of a response thus reflects a gradually increasing effect on a particular variable. We also show the posterior median of the time-averaged responses (solid, black line) along with 68% credible intervals. The time-averaged responses give a good, first impression of the size and significance of an estimated effect. In general, by averaging over time, we reduce one dimension of uncertainty. Hence, a significant time-averaged response does not necessarily imply that all the underlying, time-varying responses are significant at the same level.

Looking at the euro area conventional tightening monetary policy shock, we find persistent negative effects on both inflation and output growth. Regarding the latter, the color shading of the impulses suggests that the middle segment of the sample span shows a more pronounced rebound than the rest of the sample. This period corresponds roughly to the global financial crisis and the euro area sovereign debt crisis (2009 to 2012). In line with responses of actual output growth and inflation, expectations are also persistently negative. These results are purely data driven since we did not put restrictions on expectations for the monetary policy shock. Interest rates turn negative after about five months, while negative effects

[13]D'Amico and King (2015) impose an average restriction on interest rate, while end-point restrictions on output and inflation are used in this study. This represents an important distinction since they use expectations of interest rates defined as the average over the following our data are forecasts of output and inflation with a one year ahead horizon. The same is true for interest rate forecasts.

Fig. 2: Euro area responses to a euro area monetary policy tightening



Notes: The plot shows posterior median responses over the sample period. Light yellow responses correspond to the beginning of the sample (i.e., 2001m1), dark red responses to the end (i.e., 2018m6). The black line corresponds to the posterior median of the time averaged response along with 68% credible bounds.

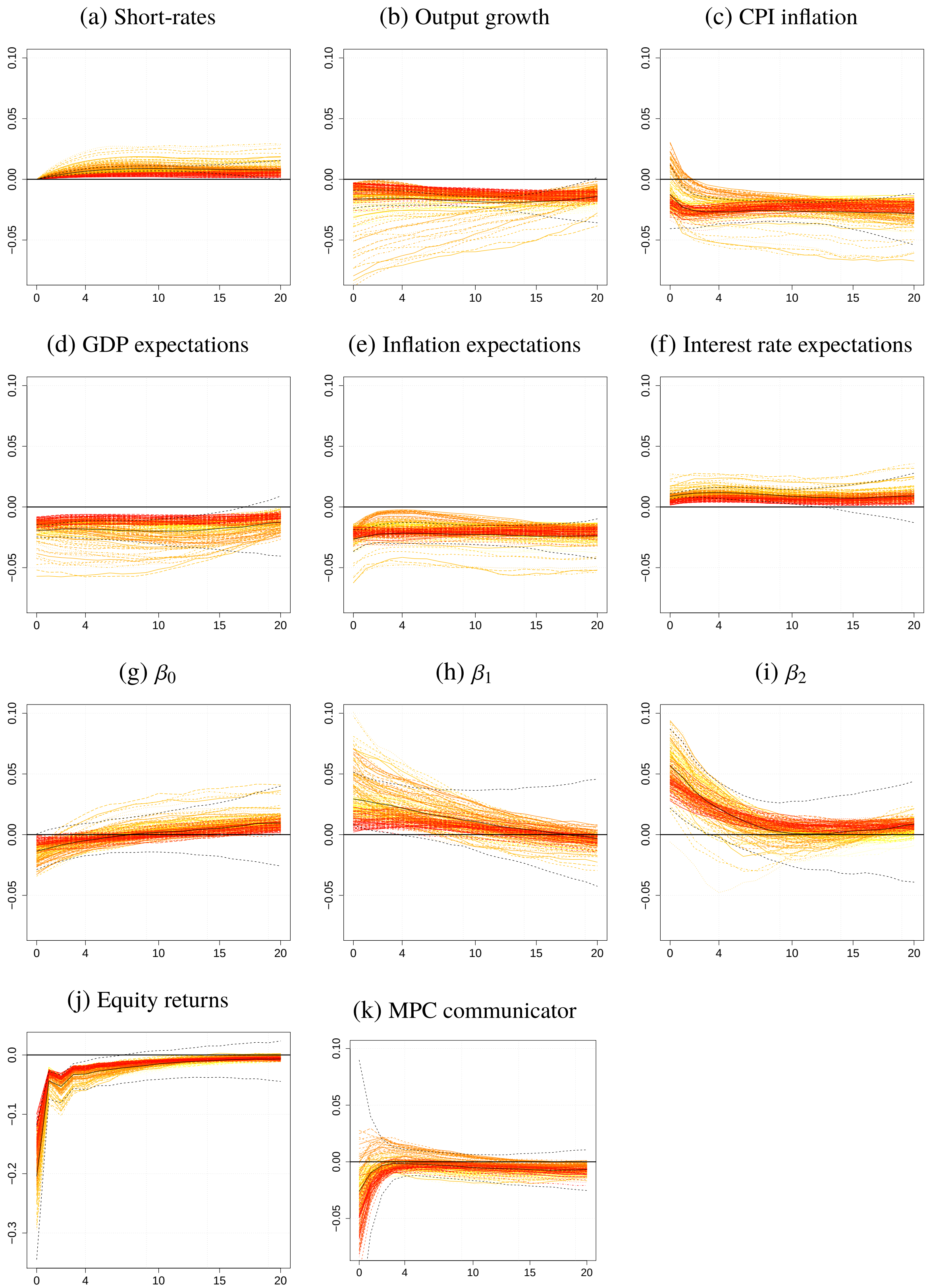
on interest rate expectations are more front-loaded. Equity prices show the most pronounced, negative reaction, which is precisely estimated up to four periods. Looking at movements of the yield curve, we find a positive effect on the inverse slope coefficient (β^1) suggesting a flattening of the yield curve which is typical after a monetary policy tightening. The reason is that the rate increase at the short-end is not fully passed on to the long end of the curve (Benati and Goodhart, 2008). Here, additional downward pressure on the long-end of the curve stems from the negative reaction of inflation expectations. We also find a significant and negative response of the curvature factor (β^2), which implies less mass of the yield curve in the medium-term segment and more in the short-end. There is no significant response of the MPC communicator control variable or the level of the yield curve (β_0), the latter which typically depends on macroeconomic fundamentals such as long-run inflation expectations (Diebold and Li, 2006).

Next, we investigate the responses to an increase in interest rate expectations. These are displayed in Fig. 3, which indicates some similarities of responses to both shocks. For example, responses of both output and inflation expectations and their observed counterparts decline. Moreover, these effects are rather persistent. The response of observed inflation also allows us to gauge how restrictive the rationality conditions are for IRF horizons up until one year. The results show a lot of variation with immediate responses for some time periods being even positive. We also see that responses of output growth to both shocks tend to be rather persistent. This indicates that the rationality conditions indeed can be considered weak over the period up to one year. Furthermore, and also consistent with responses to the monetary policy shock, the most pronounced responses are for equity returns. There are, however, also some differences between MP and FG responses. More precisely, while responses of interest rate expectations and actual interest rates are both persistently positive, the shape of interest rate responses differs from the one for the conventional monetary policy shock: In response to the FG shock, interest rates stay elevated for a considerable period of time. This is in contrast to responses from the MP shock, which show a rebound of interest rates over the medium- to longer-term horizon. This behavior might be triggered by the very nature of FG. As expectations are steered towards a period of higher interest rates, expectations and responses of observed interest rates behave accordingly. Another difference arises through effects on the yield curve. Responses of the yield curve factors imply a flattening of the yield curve (positive effect on the inverse slope coefficient, β_1) with more mass on the middle-segment (positive effect on the curvature factor, β_2). The latter fact is in line with findings from a range of recent studies. For example, Rogers et al. (2014), state that FG does typically not affect yields on instruments maturities of five years or more. Brand et al. (2010) reports a hump-shaped maturity response pattern of euro area yields to FG. Also more recently, Altavilla et al. (2019) using high frequency identification show that FG affects yields with a maturity over two years most strongly.[14] Note that in this paper, we did not place any restrictions on the yield curve factors. The fact that the FG shock loads more strongly on the middle segment compared to the MP shock is hence a strong indication that we successfully captured an FG shock by following the approach of D'Amico and King (2015).

To put our results in perspective, we finally calculate the average output growth and inflation effect over the first year. The shocks are normalized to an average increase of 10bp over the sample period in either observed (MP, shown in blue) or expected (FG, shown in orange) interest rates and the results

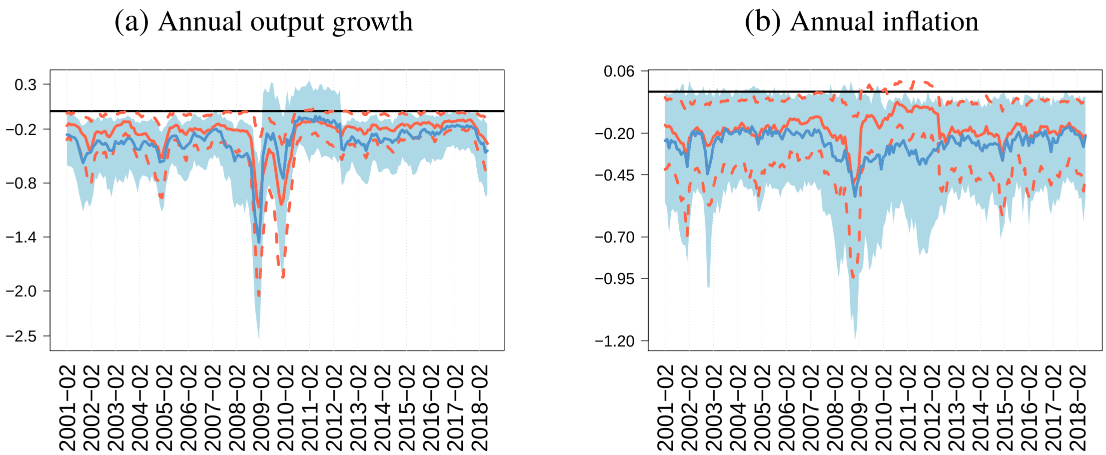
[14] Kortela and Nelimarkka (2020) take this one step further and identify the FG shock via its effect on the yield curve and curvature factor.

Fig. 3: Euro area responses to a euro area forward guidance tightening



Notes: The plot shows posterior median responses over the sample period. Light yellow responses correspond to the beginning of the sample (i.e., 2001m1), dark red responses to the end (i.e., 2018m6). The black line corresponds to the posterior median of the time averaged response along with 68% credible bounds.

Fig. 4: Annualized effects on output growth and inflation over time



Notes: The plot shows effects of an increase of 10bp (averaged over the sample period) in either interest rates (MP) or expected interest rates (FG) on output growth and inflation after one year, over time. In blue, posterior median and 68% credible interval of the MP shock; in orange posterior median and 68% credible interval of the FG shock.

depicted in Fig. 4. The results show that output growth decreases by about 0.3 percentage points and CPI inflation by about 0.25 percentage points in response to a standard 10bp monetary policy hike. These results are well in line with those of recent studies (see e.g., Lakdawala, 2019; Kortela and Nelimarkka, 2020). The numbers are somewhat smaller in case of the FG shock: 0.25 percentage points for output growth and 0.18 percentage points for inflation. Again, these effects fall in the range spanned by recent studies. More precisely, Jarociński and Karadi (2019) and Bundick and Smith (2020b) estimate trough effects in response to a 10bp increase in expected interest rates of about 0.4% of GDP and Andrade and Ferroni (2020), Kortela and Nelimarkka (2020) and Kim et al. (2020) find somewhat smaller trough effects of 0.2% on industrial production. The comparison with the existing literature serves as another plausibility check of our results and the appropriateness of our identification strategy.

Fig. 4 also suggests that effects are not time-invariant over the sample period. For conventional monetary policy, the effect peaks prior to the global financial crisis, when the ECB in parallel with other major central banks, slashed interest rates. The same trough response can be seen by looking at the response to the FG shock. However, in addition to the trough response at the onset of the global financial crisis, there is another negative spike that coincides with the euro area sovereign debt crisis. Compared to output growth responses, reactions of inflation to the monetary policy shock seem less time-varying. This is not the case for the FG shock, which shows a pronounced response of inflation at the onset of the global financial crisis. These results suggest two conclusions: First, monetary policy is more effective during a financial crisis in the sense that it triggers larger output effects. This finding corroborates results of Jannsen et al. (2019) who investigated monetary policy effectiveness during financial crises for a broad set of countries. Jannsen et al. (2019) postulate that monetary policy is especially powerful during these periods since it is able to significantly alleviate financial market distress and reduce uncertainty. Our results also show that FG is relatively more effective in steering inflation, again a finding that is in line with recent studies (see e.g., Bluwstein and Canova, 2016). Second, our evidence can also be seen in support of the ZLB irrelevance hypothesis since we find little evidence for stark differences in the responses before and after the ZLB was binding. Hence, our results compare favorably with the findings of Debortoli et al. (2019) who also find similar responses comparing the ZLB period to the pre-ZLB period.

4.2. *International Effects*

In this section, we examine the effects of the euro area shocks on its neighboring countries. To do this more systematically, we will proceed in two steps. First, we focus on regional impulse response functions that are displayed in a similar fashion than in the previous section. This will allow us to study the average effect over time as well as time-variation. Second, we zoom in on the cross-country dimension by analyzing the total effect of the monetary policy shocks, aggregated over the sample period. The full set of responses for all countries and over time is shown in the appendix (Fig. C.1 to Fig. C.12).

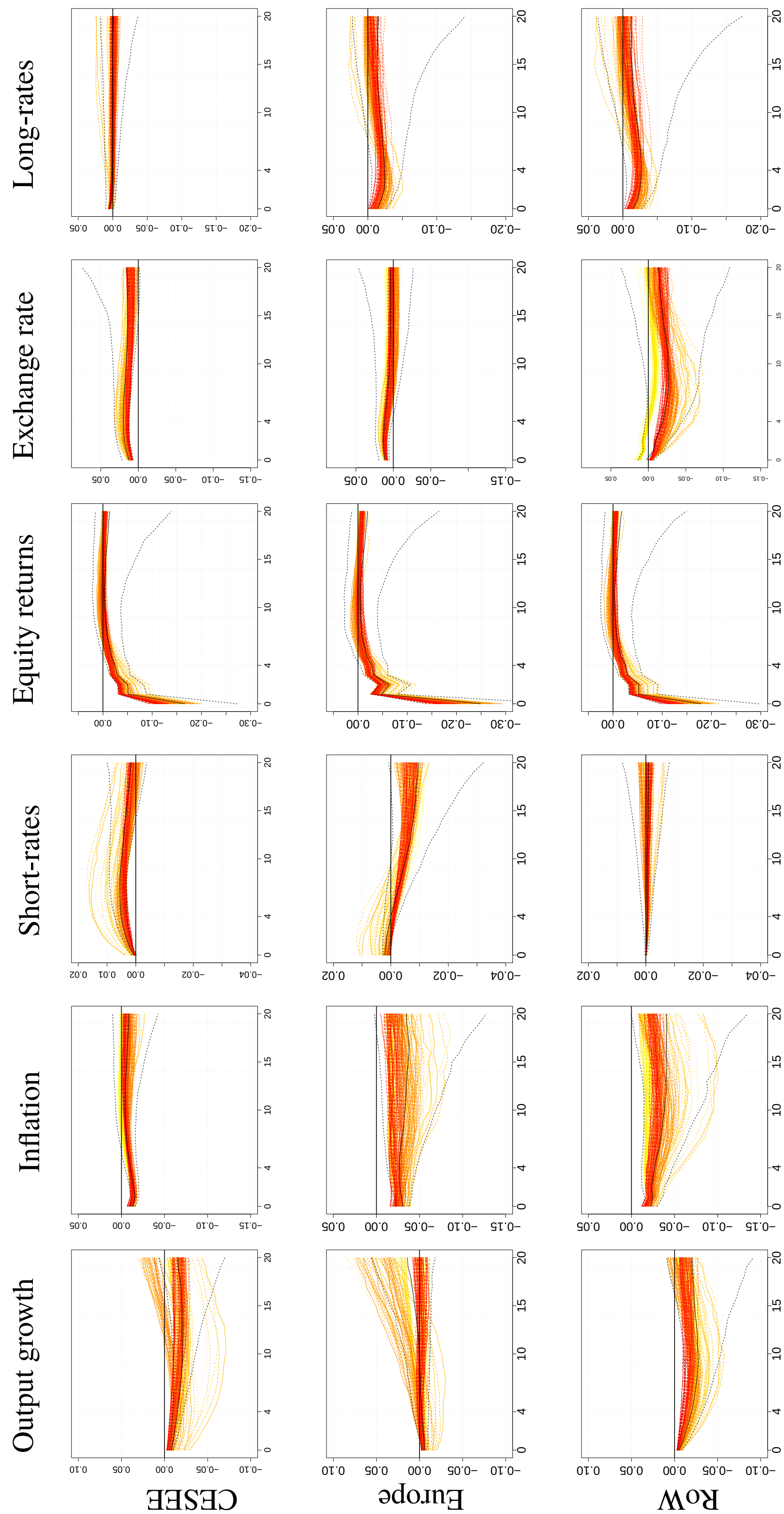
We start with the regional analysis in Figures Fig. 5 and Fig. 6. Regional means are calculated using simple averages for the three regions defined in Sec. 2: the CESEE region, other non-euro area European countries (Europe) and major advanced and emerging economies (RoW).

The effects of a conventional MP tightening shock on inflation are significantly negative in all three regions. They also show less time variation compared to output effects and are more front-loaded in CESEE economies compared to their regional peers. Policy rates rise in the short-term in all three regions. They revert, however, in European countries in the longer-term. Zooming in again reveals that this is mainly driven by responses in Great Britain (see Fig. C.3), but that interest rates also decline in the USA.[15] The latter finding is in line with predictions of standard open-economy DSGE models which suggest that, if the country where the monetary policy shock originates is large relative to the spillover receiving economies, domestic interest rates tend to increase in parallel with foreign rates - facilitated indirectly via a hike in world interest rates (Svensson and van Wijnbergen, 1989; Galí and Monacelli, 2005). This supporting effect via world interest rates is not present in case the receiving country is comparably large, which applies for example to the USA. Next, we examine effects on financial variables, such as asset prices and exchange rates. Not surprisingly, we find strong, front-loaded effects on equity returns, which mirror domestic effects in the euro area. Looking at exchange rates, we would expect a broad-based depreciation of countries with floating exchange rate regimes (Inoue and Rossi, 2019). However, we find evidence for regional heterogeneity across the three country groups: While in CESEE and neighboring European economies, currencies depreciate, other major economies in the rest of the world show a slight appreciation tendency. This applies in particular to China, Japan, Switzerland and the USA. Finally, we find evidence for a front-loaded decline in long-term yields. Together with the increase in policy rates, this implies a compression of the term spread, at least in the short-run.

The responses for the FG shock are depicted in Fig. 6. Compared to the conventional MP shock, the increase in interest rate expectations triggers a more persistent and significant decline in output growth throughout all three regions. Negative spillovers also carry over to inflation, which persistently decline in the longer term in European neighboring countries as well as in the rest of the world. The response of short-term interest rates is positive and significant in CESEE and non-euro area European economies. Compared to effects of a conventional monetary policy shock, interest rate reactions are much more

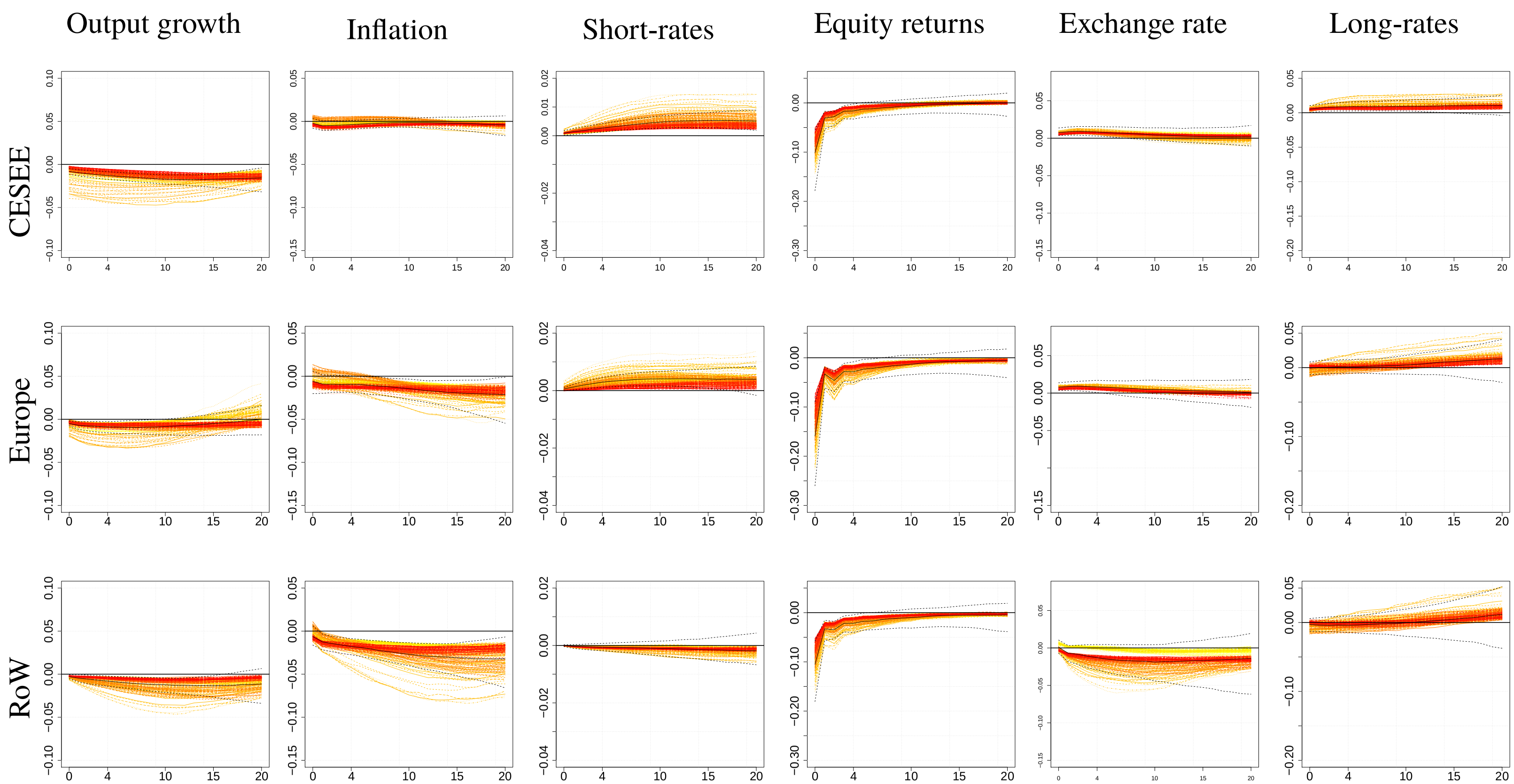
[15] Both countries were early and aggressive adopters of QE and the responses are also consistent with the view that, as QE persisted, its impact on the real economy declined. See, for example, Lombardi et al. (2019), and references therein.

Fig. 5: Regional impulse responses to a conventional monetary policy (MP) tightening



Notes: The plot shows posterior median responses over the sample period. Light yellow responses correspond to the beginning of the sample (i.e., 2001m1), dark red responses to the end (i.e., 2018m6). The black line corresponds to the posterior median of the time averaged response along with 68% credible bounds. Regional responses calculated as simple averages over the countries in a region.

Fig. 6: Regional impulse responses to a forward guidance (FG) tightening



Notes: The plot shows posterior median responses over the sample period. Light yellow responses correspond to the beginning of the sample (i.e., 2001m1), dark red responses to the end (i.e., 2018m6). The black line corresponds to the posterior median of the time averaged response along with 68% credible bounds. Regional responses calculated as simple averages over the countries in a region.

persistent. Similar to responses from the MP shock, we find significant and pronounced effects on equity returns. Also, the responses of exchange rates are rather similar to the ones from the MP shock. This finding is in line with results of Inoue and Rossi (2019). Long-term rates increase significantly in CESEE and tend to increase in the other regions. This distinct reaction in response to the FG shock might be explained by different yield curve reactions in the euro area to the two shocks.

Next, we consider effects over the full sample period which are displayed in Fig. 7. As a summary measure, we calculate the area under the curve, i.e., the area spanned by the posterior median response and the origin over the parts of the impulse response horizon that are significant (according to the 68% credible intervals). We take the posterior median as a reference point as opposed to the 68% credible bounds since the median is the most commonly used point estimate to quantify the effects of a shock. Note that either a large but short-lived effect or a small but persistently significant effect can lead to the same area under the curve measure.

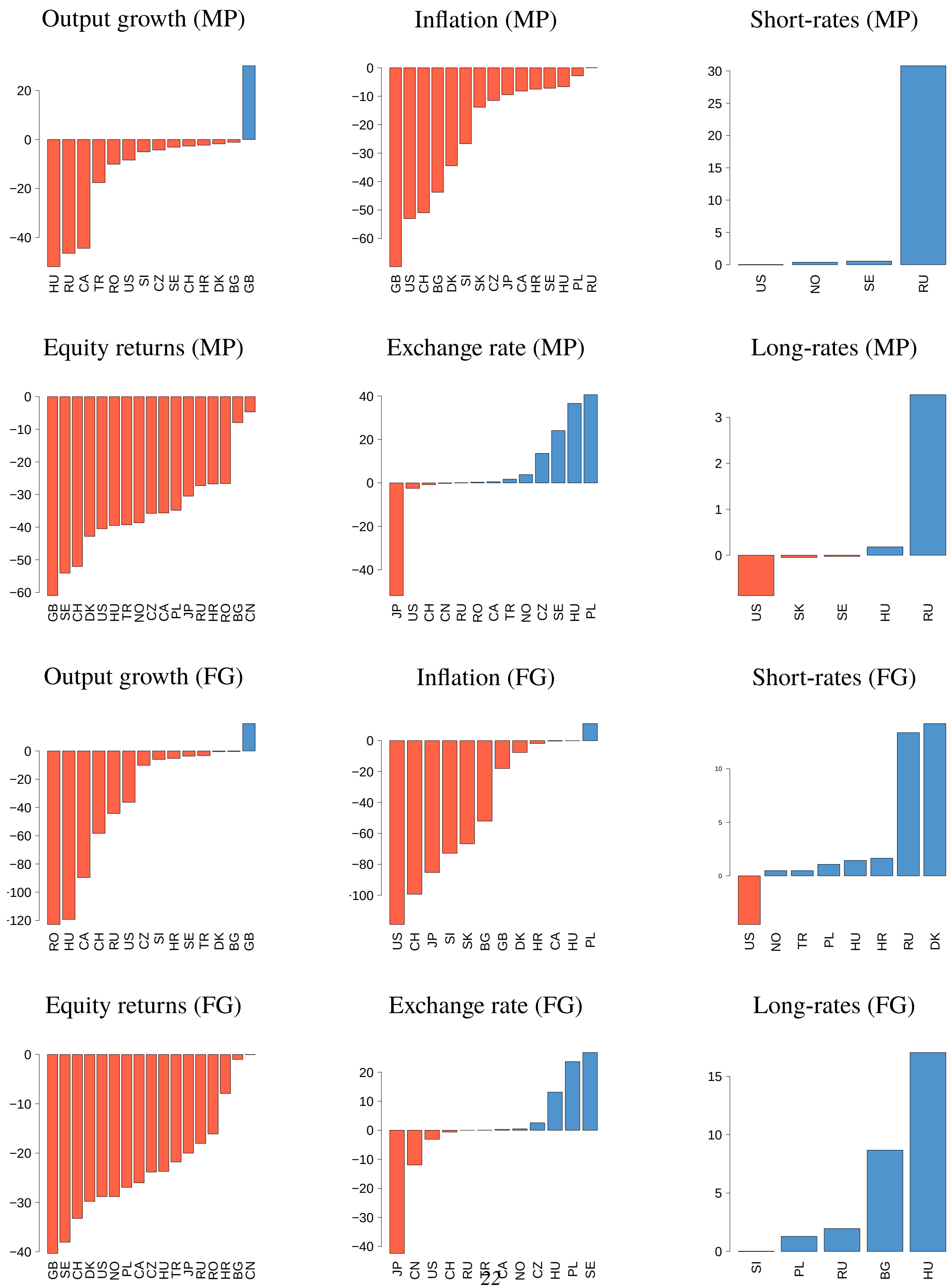
Starting with international responses to the conventional monetary tightening we find broad-based significant and negative effects on output growth, inflation and equity returns. Recall that structural generalized impulse responses in a GVAR framework reflect the direct (spillover) effects, indirect effects via third-countries, as well as the domestic response to the shock of a given country. Nevertheless, the figure shows that for these variables, responses are negative for most economies. In some cases, exchange rates depreciate against the euro and interest rates ease. More specifically, currencies weaken against the euro in European, neighboring countries (e.g., Czechia, Hungary, Poland, Norway, Sweden, Romania), but also in Canada, Russia and Turkey. In general, the countries affected range from neighboring countries of the euro area to major economies such as the USA. This conclusion also carries over to financial variables such as equity returns indicating that euro area shocks can have far reaching consequences. Some countries deviate from the general, international pattern. For example, in Great Britain we find a positive peak effect on output growth and a negative peak effect on inflation. This resembles the reaction of a negative domestic supply shock.

Finally, we examine responses to the increase in expected interest rates. Again, responses are mostly negative, implying that the FG shock depresses international output growth, inflation and equity returns. Compared to international reactions to the conventional monetary policy shock, the area under the curve in response to an increase in interest rate expectations is generally larger. This implies either that responses to $i^{s,t+12}$ exceed ones to i^{s} in terms of magnitude or are estimated with greater precision. Another difference compared to spillovers from the conventional monetary policy shock can be seen by looking at interest rates. There are more positive reactions of both short- and long-term interest rates, which might be related to the more persistently positive domestic reaction of euro area short-term interest rates and the transmission over the yield curve.

4.3. *Do Spillovers Differ Over the Sample Period?*

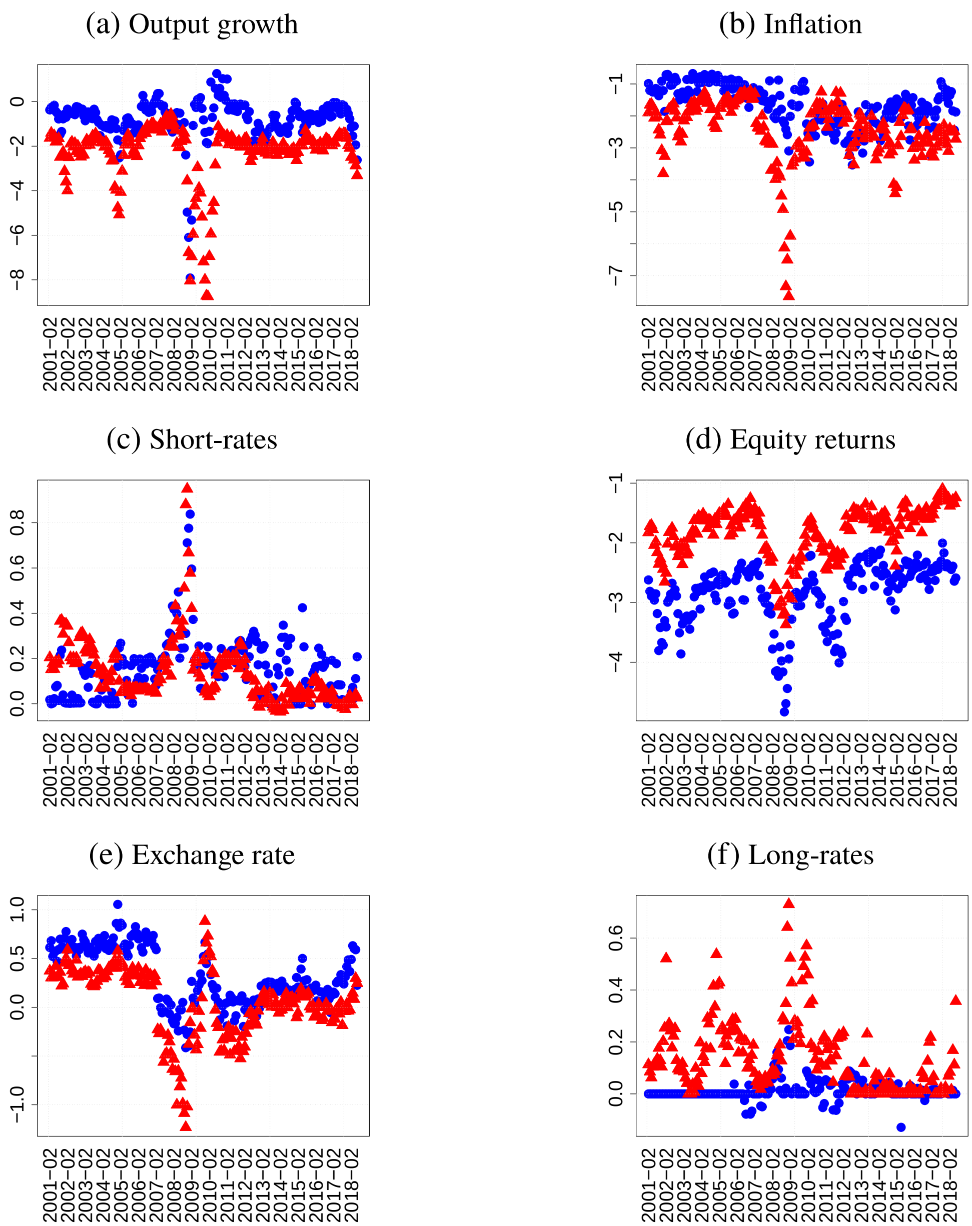
In this section, we investigate in more detail whether spillover effects vary over time. For that purpose, we calculate the area under the curve, aggregated per international (i.e., non-euro area) variable for each point in time. The results are provided in Fig. 8, where blue circles correspond to the conventional monetary policy shock and red triangles to results for the forward guidance shock.

Fig. 7: Area under the curve of significant international responses



Notes: The plot shows the area under the curve, calculated from the posterior median to the origin, over impulse response horizons that are significant according to the 68% credible intervals. The areas under the curve are summed over all time periods in the sample. Red bars indicate negative (total) areas, blue bars positive (total) areas.

Fig. 8: Area under the curve over time



Notes: The plot shows the area under the curve, calculated from the posterior median to the origin, over impulse response horizons that are significant according to the 68% credible intervals per variable and over time. Blue circles (●) refer to to monetary policy, red triangles (▲) to forward guidance.

Focusing on the responses to the conventional monetary policy shock first, we see that effects are strongest and most precisely estimated starting with the period of the onset of the global financial crisis. For most variables, trough responses occur roughly in the early months of 2009. A difference arises if we compare the time profile of spillovers to real and financial variables. More specifically, we see considerably more time variation for equity returns and exchange rates up until 2012, whereas output responses vary considerably only in a tight window from 2009 to 2010. Turning next to the FG shock, we see a similar pattern for output which responds most strongly during the period of the global financial crisis and financial variables, which show considerable time variation up until 2010. Comparing responses from the two shocks reveals two striking differences, namely the time profile of spillovers to inflation and long-term interest rates. International effects on inflation are rather time-invariant in response to the MP shock, but show a considerable trough effect during the global financial crisis for the FG shock (Bluwstein and Canova, 2016). Interestingly, the time profile of spillovers to long-term rates is much more pronounced for the FG relative to the MP shock. This again might be driven by different yield curve dynamics in the euro area triggered by the two shocks.

Our findings broadens the results of Jannsen et al. (2019) to include international effects. They investigate only domestic monetary policy effectiveness to spillovers and forward guidance. The reason why monetary policy might be particularly effective during episodes of financial crises are manifold: through the impact on uncertainty (Bekaert et al., 2013), by providing signals about future economic prospects (Ilut and Schneider, 2014) and through effects on financial markets and hence the credit channel (Jannsen et al., 2019). Our results also relate to the literature on monetary policy effectiveness during the period of the ZLB. More specifically, if we compare responses to the two shocks during the pre-ZLB (i.e., pre-crisis) and the ZLB (i.e., post-crisis) period, we find rather similar effects (Debortoli et al., 2019).

A note on how our model captures time-variation, seems in order. Our empirical framework models both time variation that stems from the variance of the residuals (stochastic volatility) up- / down-scaling the responses as well as from changes in the relationships among the variables (drifts in the coefficients) which impact the shape of the responses. It is not straightforward to disentangle the two sources.[16] A model that wrongly shuts down stochastic volatility or holds the coefficients constant would lead to misleading inference. We also cannot normalize the shocks (i.e., holding the size on impact constant) since both interest rates and interest rate expectations are nearly constant during large parts of the sample period (and hence the volatility close to zero). In a robustness exercise, we have normalized the shocks to yield always the same trough effect on output. This exercise has shown that even when holding the strength of the shocks constant (in terms of output effects), impulse responses show considerable time variation. That said, having volatilities and coefficients potentially drift is the most comprehensive way in dealing with changes in the economic environment, the business climate and more generally the overall level of uncertainty, all factors that can influence the effectiveness of monetary policy (Jannsen et al., 2019).

5. CONCLUSIONS

In this paper, we investigate the domestic and international effects of euro area forward guidance shock modelled by a change in interest rate expectations. To gauge the importance of FG, we compare these effects to those of a standard or conventional monetary policy shock that works through a change in actual

[16]As another source of time variation in the GVAR context, one could consider letting the trade weights vary over time. In our setting, potential changes in the connectivity would be captured by drifts in the coefficients. It is beyond the scope of this paper to investigate whether these changes are driven by changes in the economic relationship between countries or by changes how these countries react to foreign factors.

interest rates. Both shocks are identified using a combination of zero and sign restrictions following D'Amico and King (2015) which puts structure on the joint behavior of impulse responses of observed and expectations data.

Our principal conclusions are as follows. First, a conventional and contractionary monetary policy shock in the euro area triggers a broad-based deceleration in domestic and international output growth, inflation and equity returns. Domestic negative responses are amplified by negative output growth and inflation outlooks for the euro area. Our results also indicate strong, negative effects on equity returns. However, in contrast to the remaining variables, these are rather short-lived. Exchange rate responses are country-specific with currencies of euro area's neighbors tending to weaken against the euro.

Second, we investigate the international effects of an increase in euro area interest rate expectations under conditions of unchanged interest rates. This FG shock yields mostly similar responses as when there is a conventional monetary policy shock, namely: domestic and international output growth and inflation decelerate and equity returns decline. Compared to effects from a conventional monetary policy shock, we find more positive effects on international short- and long-term interest rates. This might be related to the domestic response of euro area interest rates, which is more persistent for the FG shock. Moreover, and in line with the recent literature (e.g., Brand et al., 2010; Altavilla et al., 2019), the FG shock loads more strongly on the middle segment of the euro area yield curve.

Finally, and perhaps most importantly, we find evidence of time variation for both shocks. Spillovers to international output and financial variables are particularly pronounced during the period of the global financial crisis. This finding generalizes results of Jannsen et al. (2019) on the domestic effectiveness of monetary policy to spillovers during financial crises. During that period, we also find pronounced effects of forward guidance – both domestically and internationally – on inflation, which contrasts with MP responses (Bluwstein and Canova, 2016). Our results also relate to the literature discussing monetary policy effectiveness at the zero lower bound. In line with Debortoli et al. (2019) we do not find significant differences in the domestic and international effects of MP and FG during the ZLB period.

Our careful and exhaustive analysis suggests that FG is a powerful tool especially in crisis conditions. That said, there are a number of caveats suggestive of extensions for future research. Of course, not all crises are of the financial variety nor are they all alike. As this is written, with a health crisis underway, how central banks manage interest rate expectations is set to become even more important. We have focused on restrictions that impact the short-end of the term structure. A natural extension is to more explicitly consider the case of yield curve control, introduced in Japan and being considered elsewhere, even for the euro area. While the distinction between Delphic and Odyssean forward guidance is helpful to fix ideas, our results demonstrate that FG can be effective regardless of the manner it is portrayed. It is worth noting that there is no explicit modelling of the feedback relationship between central bank actions and expectations in our framework. An extension might consist in asking what would the impact of FG be if credibility is explicitly considered. This is an especially challenging task since there is no agreement on how to measure credibility. It is usually known only ex post whereas, ideally, we would like to know ex ante what drives this ingredient of monetary policy. We leave these extensions for future research. Another difficulty with forward guidance studies is that theory has yet to provide clearer indications of the macroeconomic effects of forward guidance. Standard New Keynesian models are seen as inadequate. Hence, models are being developed to handle various kinds of financial frictions, non-heterogeneity of agents, and the reaction of the fiscal authorities when interest rates are at the effective lower bound (e.g., Piazzesi et al., 2019; Woodford and Xie, 2020; Kaplan et al., 2003), not to mention the type of expectations mechanism at play. In the meantime, the results of this paper offer indications of the potential of forward guidance as a policy tool. Additional research will be necessary to work out the various channels and mechanism at play in crisis or at other times.

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A. PRIOR SPECIFICATION

To select whether a given regressor should be included or excluded at time $t = 0$, we impose a normal-gamma (NG) prior akin to Griffin and Brown (2010) on $\boldsymbol{\xi}_{i0}$, the initial state of $\boldsymbol{\xi}_{it}$.

$$\boldsymbol{\xi}_{i0} \sim \mathcal{N}(0_{s_i}, \underline{V}_{\xi_i}), \tag{A.1}$$

with $\mathbf{0}_{s_i}$ a s_i-dimensional vector of zeros and $\underline{\boldsymbol{V}}_{\xi_i}$ is a $k_i m_i \times k_i m_i$ diagonal prior variance-covariance matrix. We assume that each diagonal element of $\underline{\boldsymbol{V}}_{\xi_i}$, labeled $\underline{v}_{\xi_i j}$, features a Gamma prior with,

$$\underline{v}_{\xi_i j} \sim \mathcal{G}(\theta_i, \theta_i \kappa_i/2). \tag{A.2}$$

The prior has two elements that control the degree of shrinkage, a local parameter and variable specific parameter $\underline{v}_{\xi_i j}$ and a global parameter, that pushes all coefficients in a country model towards zero, $\kappa_i \sim \mathcal{G}(q_0, q_1)$. Also, θ_i denotes a scalar hyperparameter specific to each country that serves to control the tail behavior of the marginal prior. The characteristic of the normal-gamma prior is that even if the global degree of shrinkage is estimated to be very large, the marginal prior (i.e., with $\underline{v}_{\xi_i j}$ integrated out) allows for non-zero regression coefficients since it features fat tails. We set $\theta_i = 0.1$ and $q_0 = q_1 = 0.01$ (Crespo Cuaresma et al., 2019; Huber et al., 2019).

The degree of time-variation is governed by the latent threshold approach. Here, we use a Gamma prior on $\vartheta_{ij,1}^{-1}$

$$\vartheta_{ij,1}^{-1} \sim \mathcal{G}(n_0, n_1), \tag{A.3}$$

For the thresholds, we introduce a prior that is uniformly distributed and depends on $\vartheta_{ij,1}$,

$$c_{ij}|\vartheta_{ij,1} \sim \mathcal{U}(\pi_{ij,0}\sqrt{\vartheta}_{ij,1}, \pi_{ij,1}\sqrt{\vartheta}_{ij,1}). \tag{A.4}$$

We follow Huber et al. (2019) and set $n_0 = 3$ and $n_1 = 0.03$ and the variance of $\vartheta_{ij,0} = 0.05 \times \hat{\sigma}_{ij}$, with $\hat{\sigma}_{ij}$ denoting the OLS standard deviation of a time-invariant VAR model. We also set $\pi_{ij,0} = 0.1$ and $\pi_{ij,1} = 3$, effectively bounding the thresholds away from zero. This implies that high frequency noise in the latent states is always set equal to zero, effectively reducing uncertainty stemming from this source without seriously distorting inference.

For the log-volatilities, we use the prior setup proposed in Kastner and Frühwirth-Schnatter (2014). This implies that we use a normal prior on μ_{il} $(l = 1, \dots, k_i)$ with mean $\underline{\mu}_i = 0$ and variance $\underline{V}_{\mu_i} = 100$

$$\mu_{il} \sim \mathcal{N}(\underline{\mu}_i, \underline{V}_{\mu_i}). \tag{A.5}$$

and a beta prior for the persistence parameter ρ_{il},

$$\frac{\rho_{il} + 1}{2} \sim \text{Beta}(a_0, b_0), \tag{A.6}$$

with $a_0 = 2$ and $b_0 = 2$.

Finally, we use a non-conjugate gamma prior for ς_{ij}^2, $(j = 1, \dots, k_i)$,

$$\varsigma_{ij}^2 \sim \mathcal{G}\left(\frac{1}{2}, \frac{1}{2B_\varsigma}\right). \tag{A.7}$$

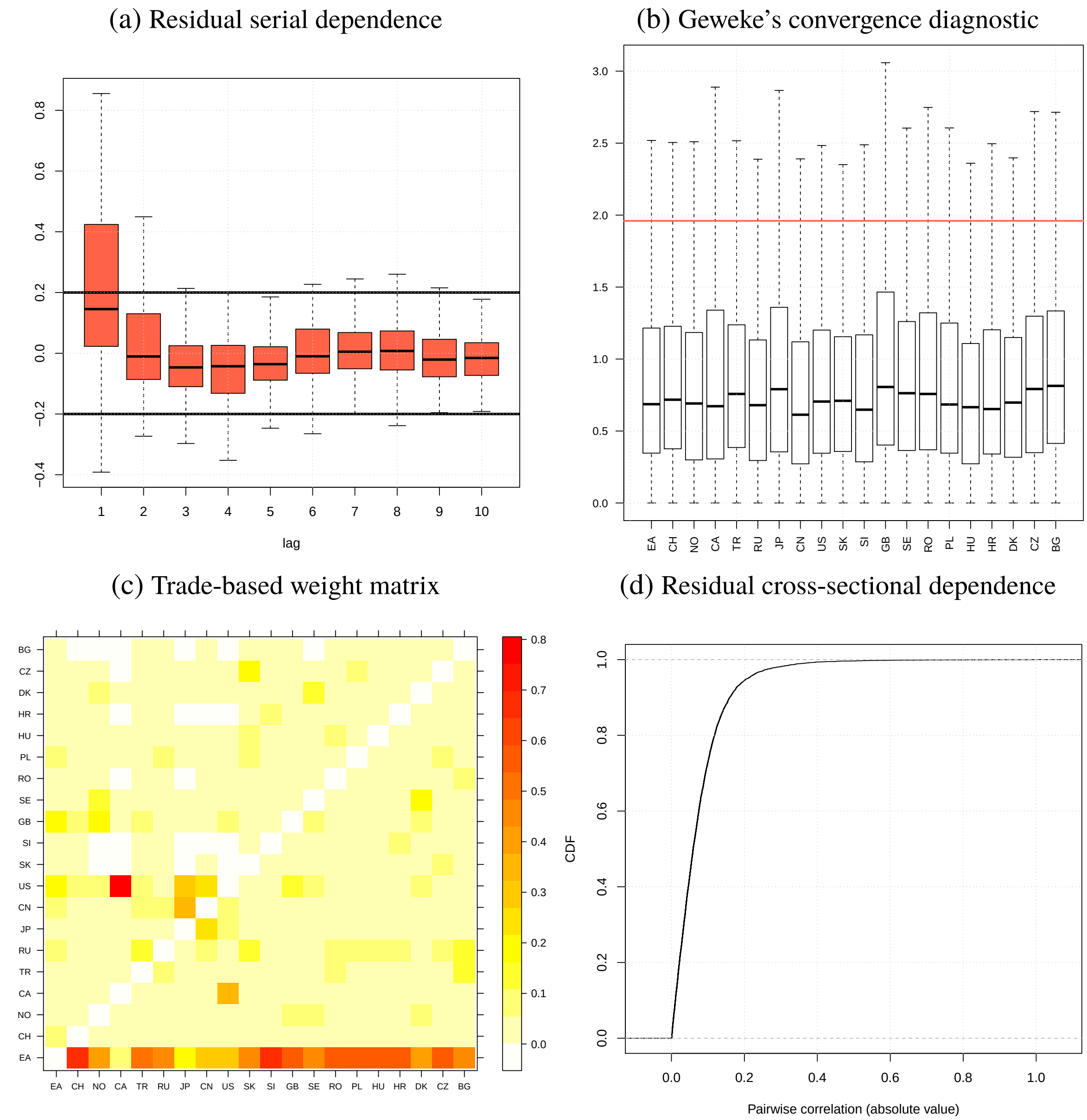
This choice does not bound ς_{il}^2 away from zero, thus providing more shrinkage than standard typical conjugate inverted gamma priors do. Here, we follow the recommendations provided in Kastner and Frühwirth-Schnatter (2014) and set $B_\varsigma = 1$.

B. CONVERGENCE PROPERTIES OF THE MCMC ALGORITHM AND RESIDUAL DIAGNOSTICS

In Fig. B.1 we show several diagnostic checks based on the residuals of the country models, obtained from 10,000 posterior draws after a burn-in phase of 30,000 draws. From the upper left panel we see that the residuals are generally not serially autocorrelated. In the top right panel we show box plots of Z-scores of Geweke's convergence diagnostic (Geweke, 1992) per country. These indicate that the MCMC algorithm has converged to its target distribution since most (absolute) values of the statistic are below the 1.96 threshold. The bottom left panel illustrates the distribution of the trade weights. One assumption underlying the GVAR framework is that the weights are relatively small (see Pesaran et al., 2004). We see that a range of mostly CESEE countries share significant trade links with the euro area (bottom row). For these countries, weights would be more equally distributed if we would analyze not the euro area aggregate but its single countries. However, in the context of monetary policy, this introduces additional challenges since assumptions about how the joint euro area monetary policy is modeled have to be made (see e.g., Georgiadis, 2015; Feldkircher et al., 2020), which is not straightforward in the context of forward guidance. Last, we show in the bottom panel of Fig. B.1, right-hand side that cross-sectional dependence of the country residuals is generally weak. The cumulative density function of the pairwise correlations across the country residuals show that 90% of the mass lies below 30% and indicating weak cross-sectional dependence (Burriel and Galesi, 2018).

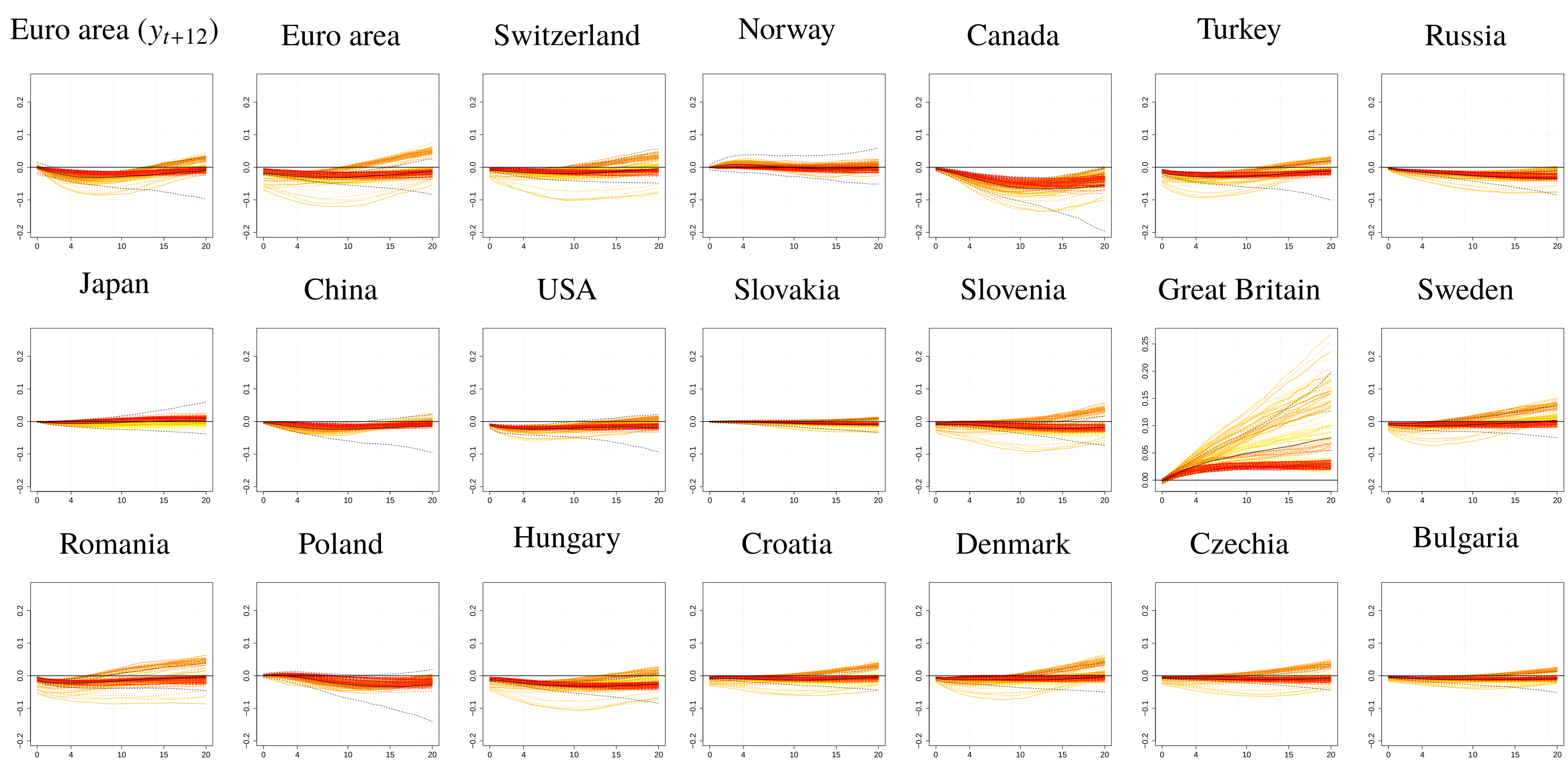
C. ADDITIONAL RESULTS

Fig. B.1: Diagnostics of the estimated GVAR.



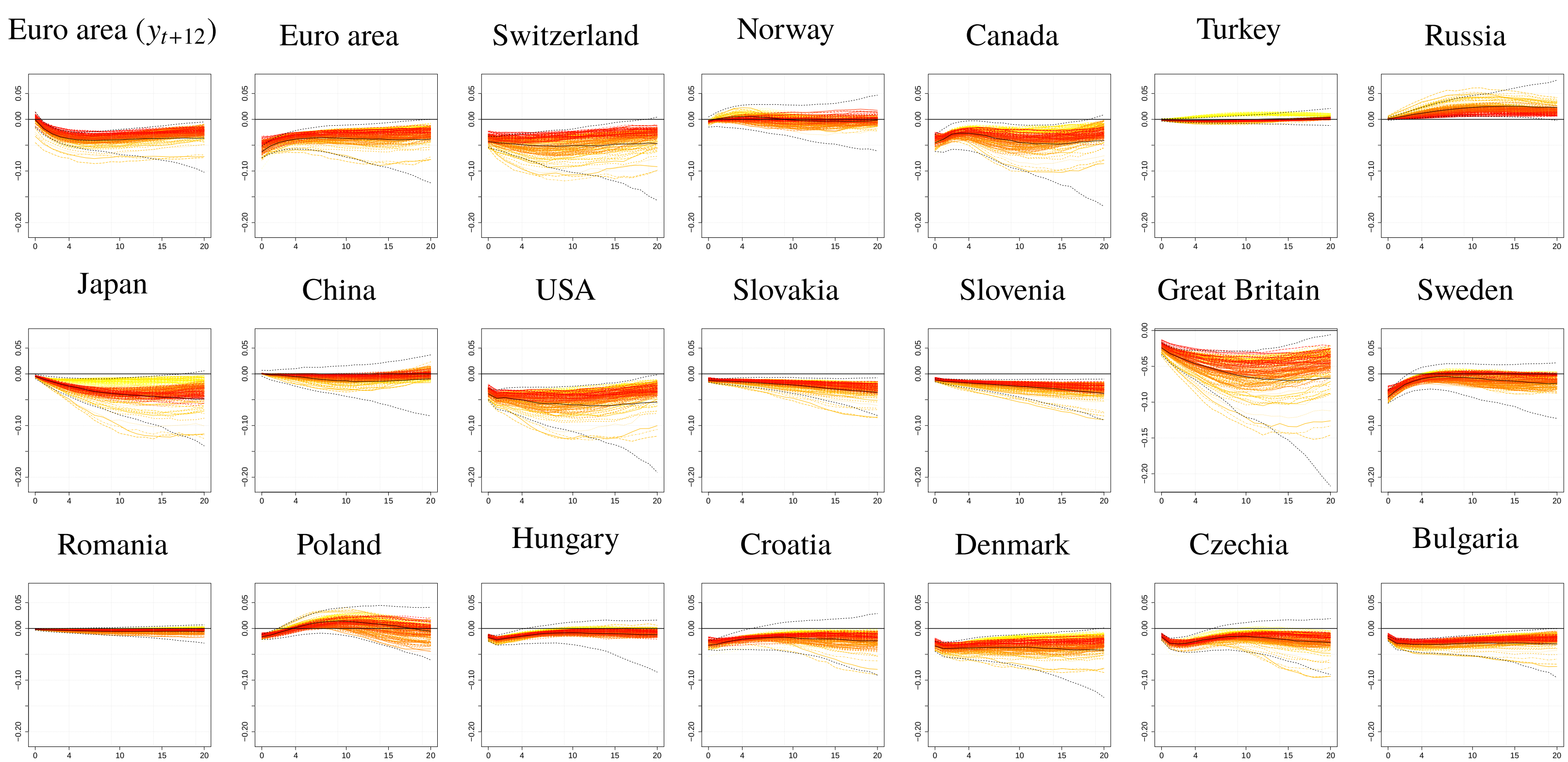
Notes: The top left panel shows the autocorrelation function (ACF) of the cross-country residuals, the right panel boxplots of Z-scores of Geweke's convergence diagnostic (Geweke, 1992). The bottom panel, left hand side shows the distribution of trade weights, the right hand side panel the empirical cumulative density function of average pairwise cross-country residual correlations (in absolute values).

Fig. C.1: Responses of output growth to a euro area monetary tigthening



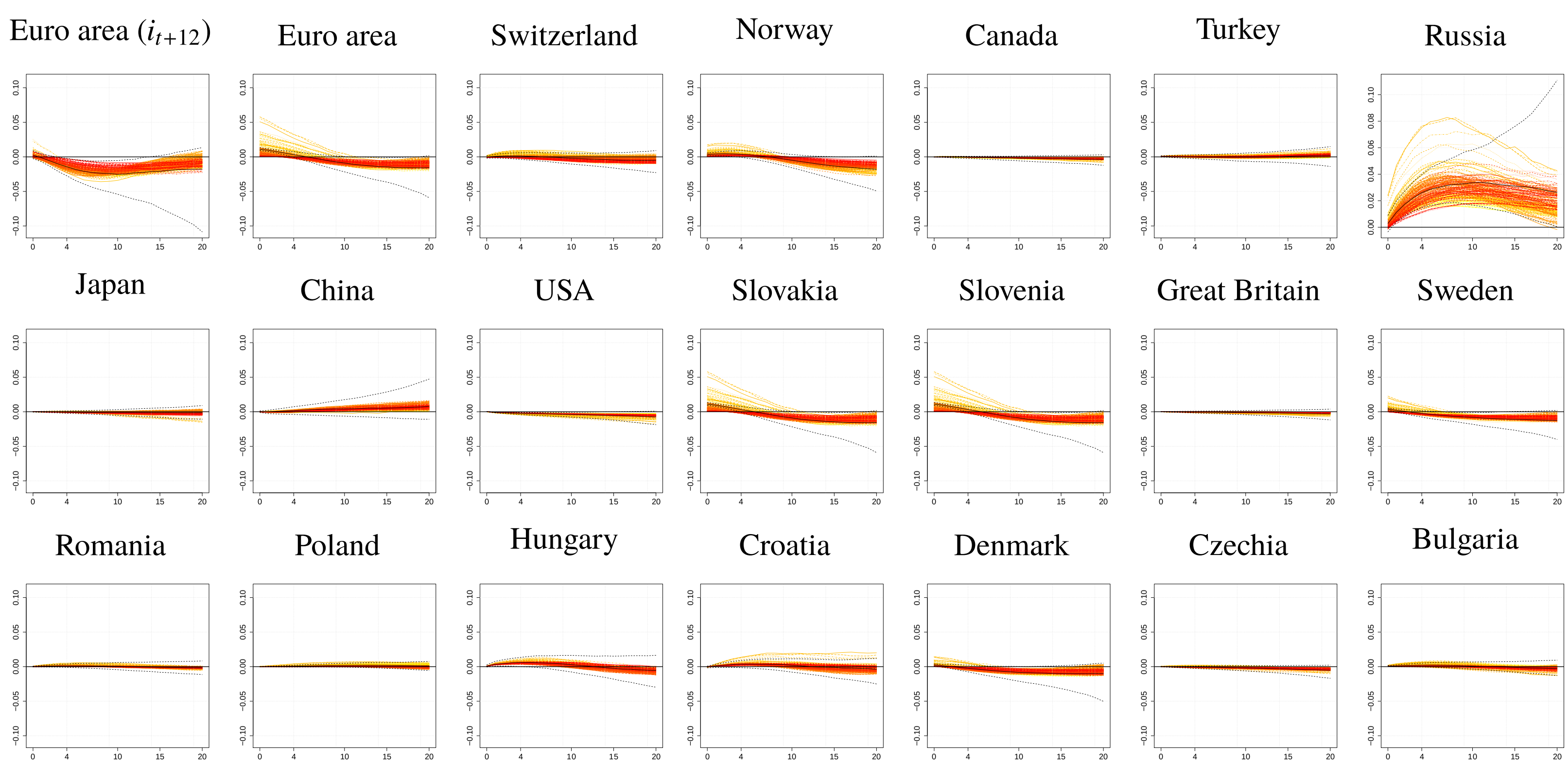
Notes: The plot shows posterior median responses over the sample period. Light yellow responses correspond to the beginning of the sample (i.e., 2001m1), dark red responses to the end (i.e., 2018m6). The black line corresponds to the posterior median of the time averaged response along with 68% credible bounds.

Fig. C.2: Responses of inflation to a euro area monetary tigthening



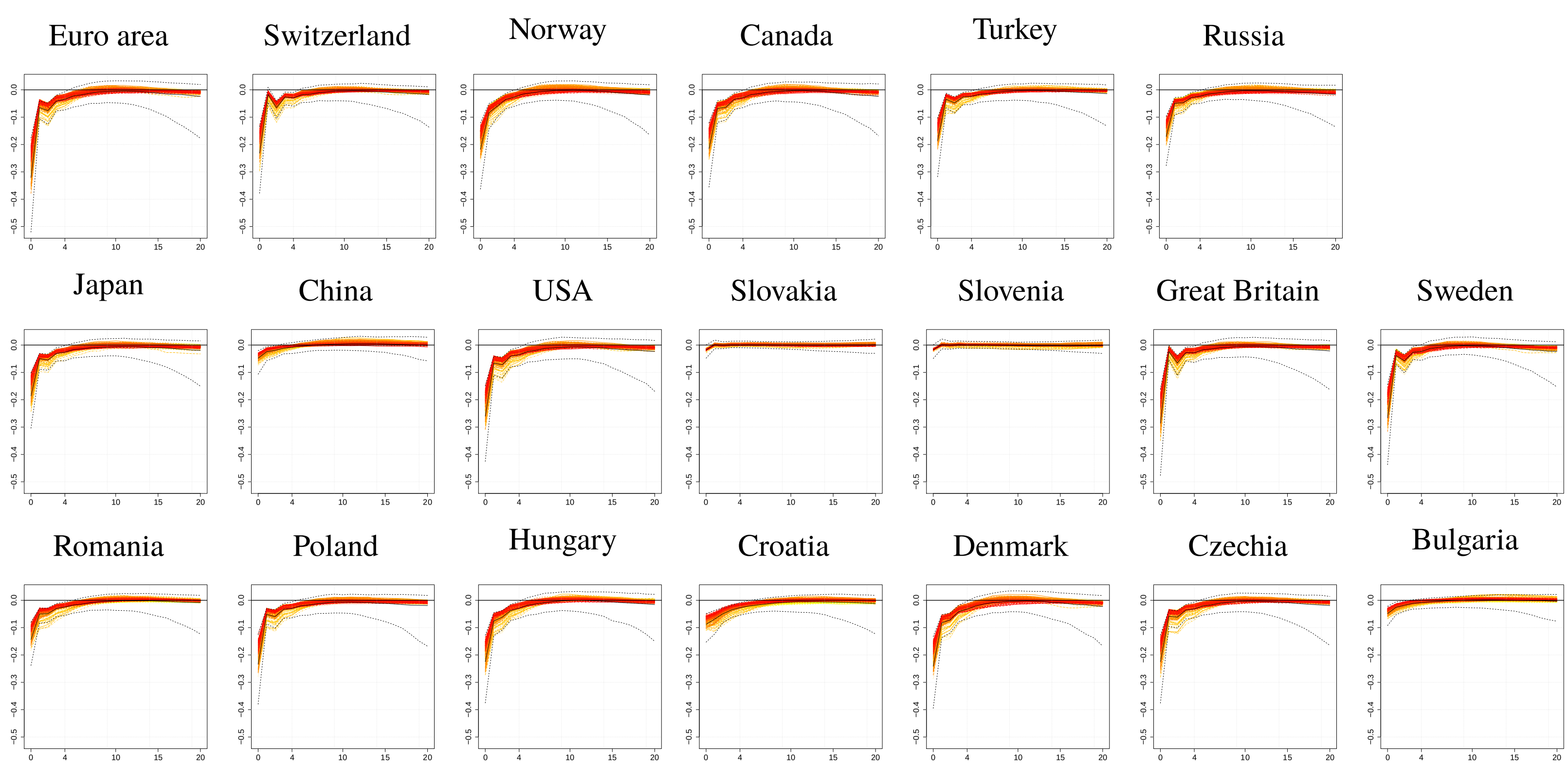
Notes: The plot shows posterior median responses over the sample period. Light yellow responses correspond to the beginning of the sample (i.e., 2001m1), dark red responses to the end (i.e., 2018m6). The black line corresponds to the posterior median of the time averaged response along with 68% credible bounds.

Fig. C.3: Responses of short-term interest rates to a euro area monetary tigthening



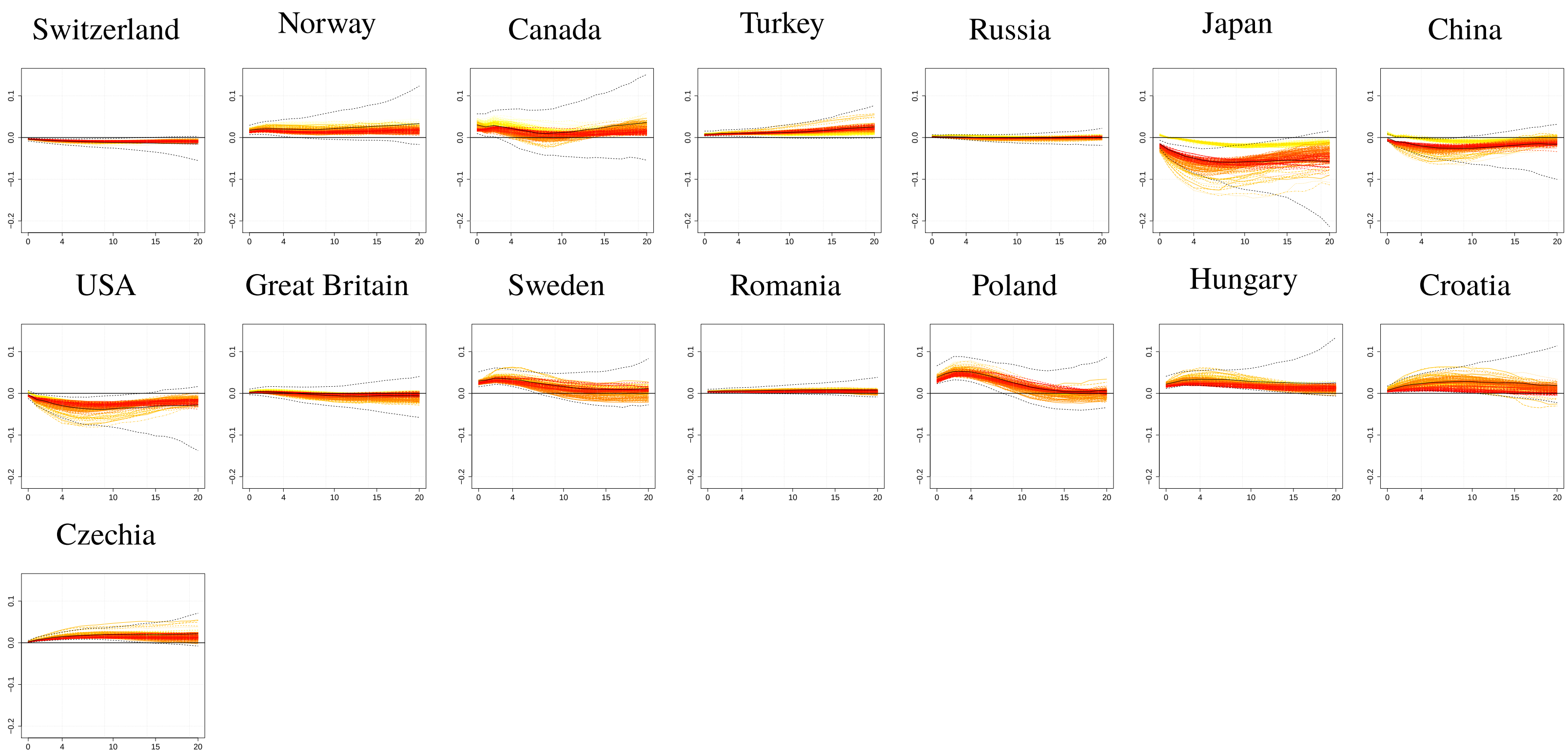
Notes: The plot shows posterior median responses over the sample period. Light yellow responses correspond to the beginning of the sample (i.e., 2001m1), dark red responses to the end (i.e., 2018m6). The black line corresponds to the posterior median of the time averaged response along with 68% credible bounds.

Fig. C.4: Responses of equity returns to a euro area monetary tigthening



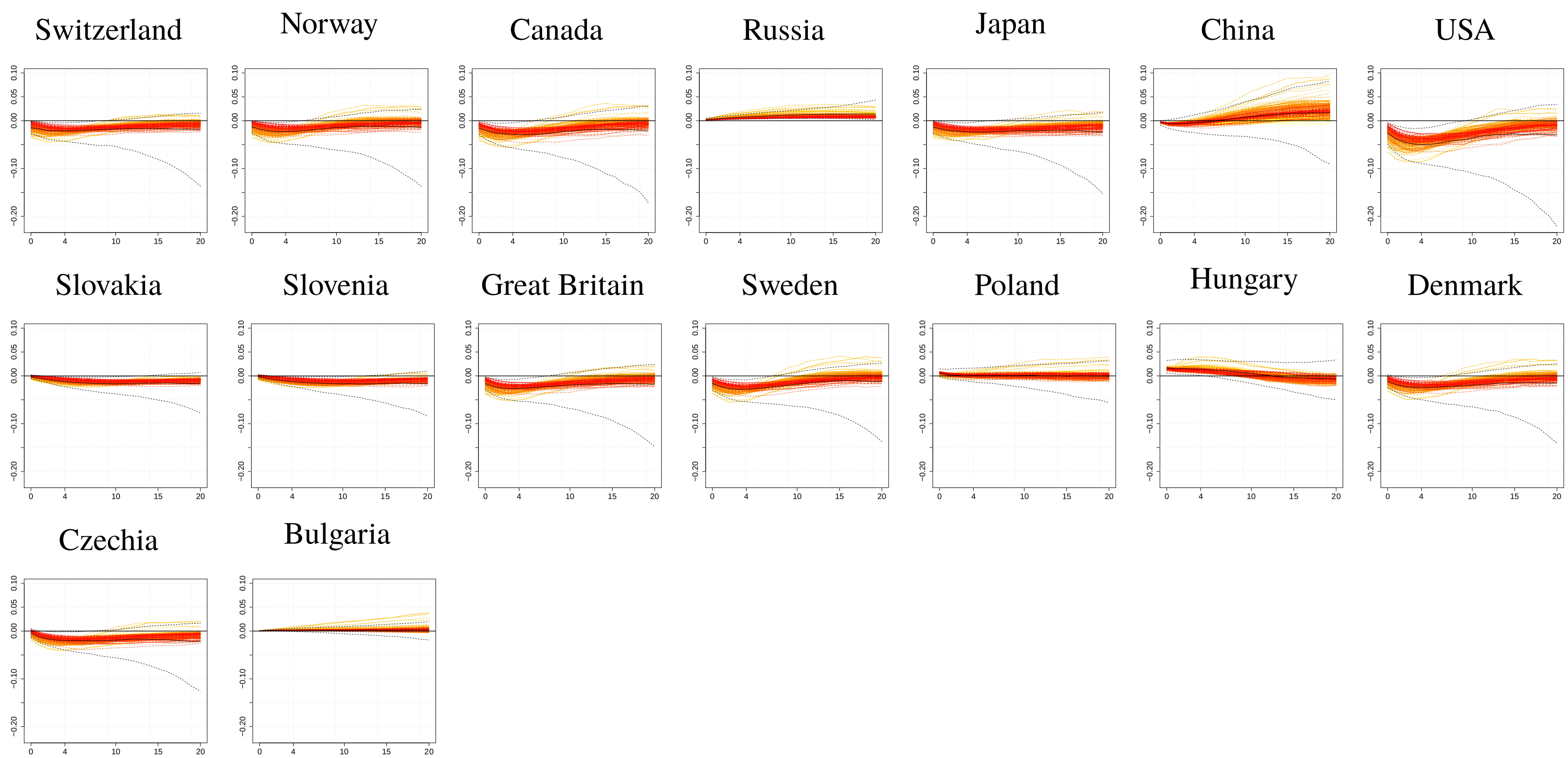
Notes: The plot shows posterior median responses over the sample period. Light yellow responses correspond to the beginning of the sample (i.e., 2001m1), dark red responses to the end (i.e., 2018m6). The black line corresponds to the posterior median of the time averaged response along with 68% credible bounds.

Fig. C.5: Responses of exchange rate changes to a euro area monetary tigthening



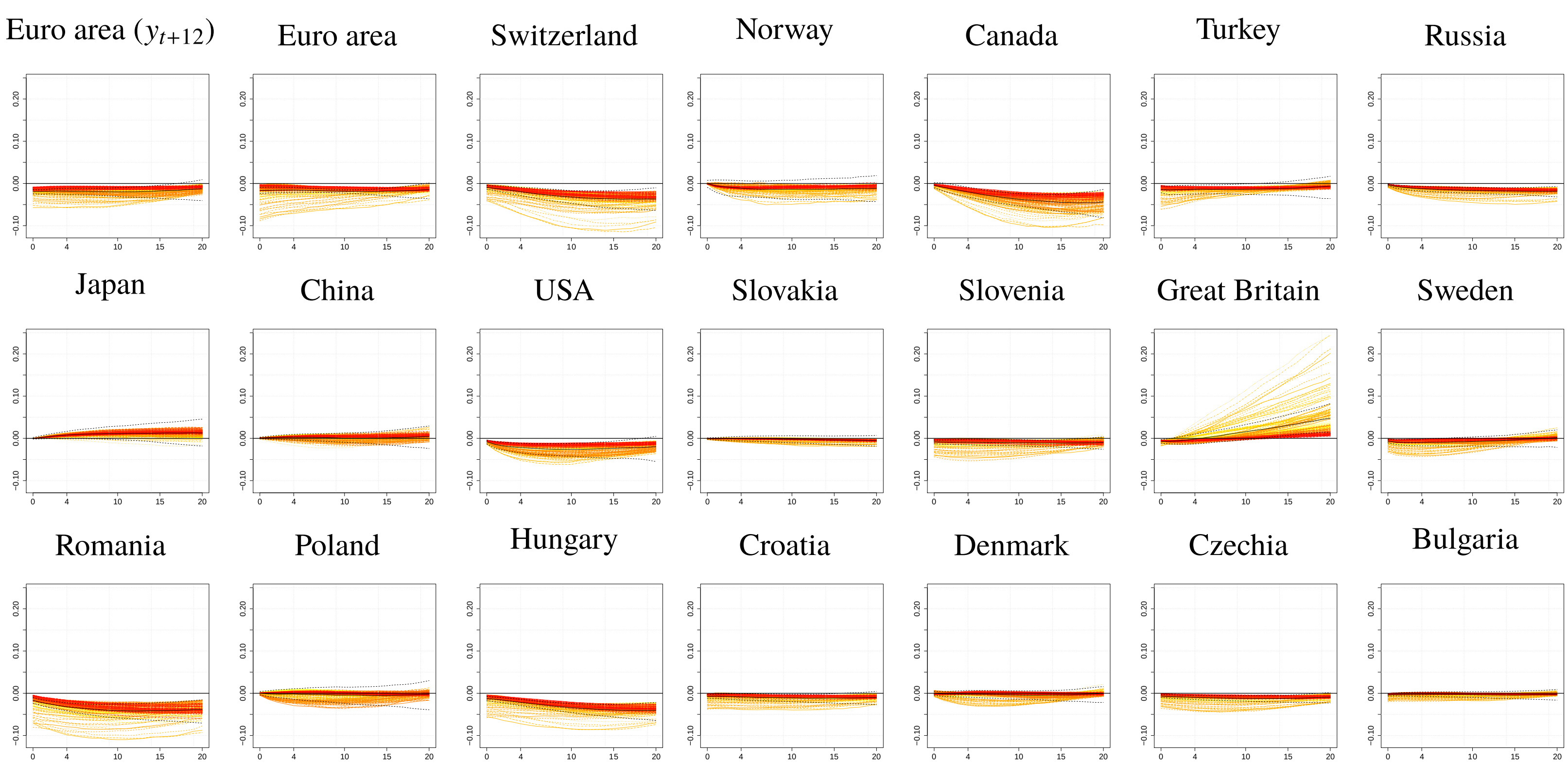
Notes: The plot shows posterior median responses over the sample period. Light yellow responses correspond to the beginning of the sample (i.e., 2001m1), dark red responses to the end (i.e., 2018m6). The black line corresponds to the posterior median of the time averaged response along with 68% credible bounds.

Fig. C.6: Responses of long-term interest rates to a euro area monetary tigthening



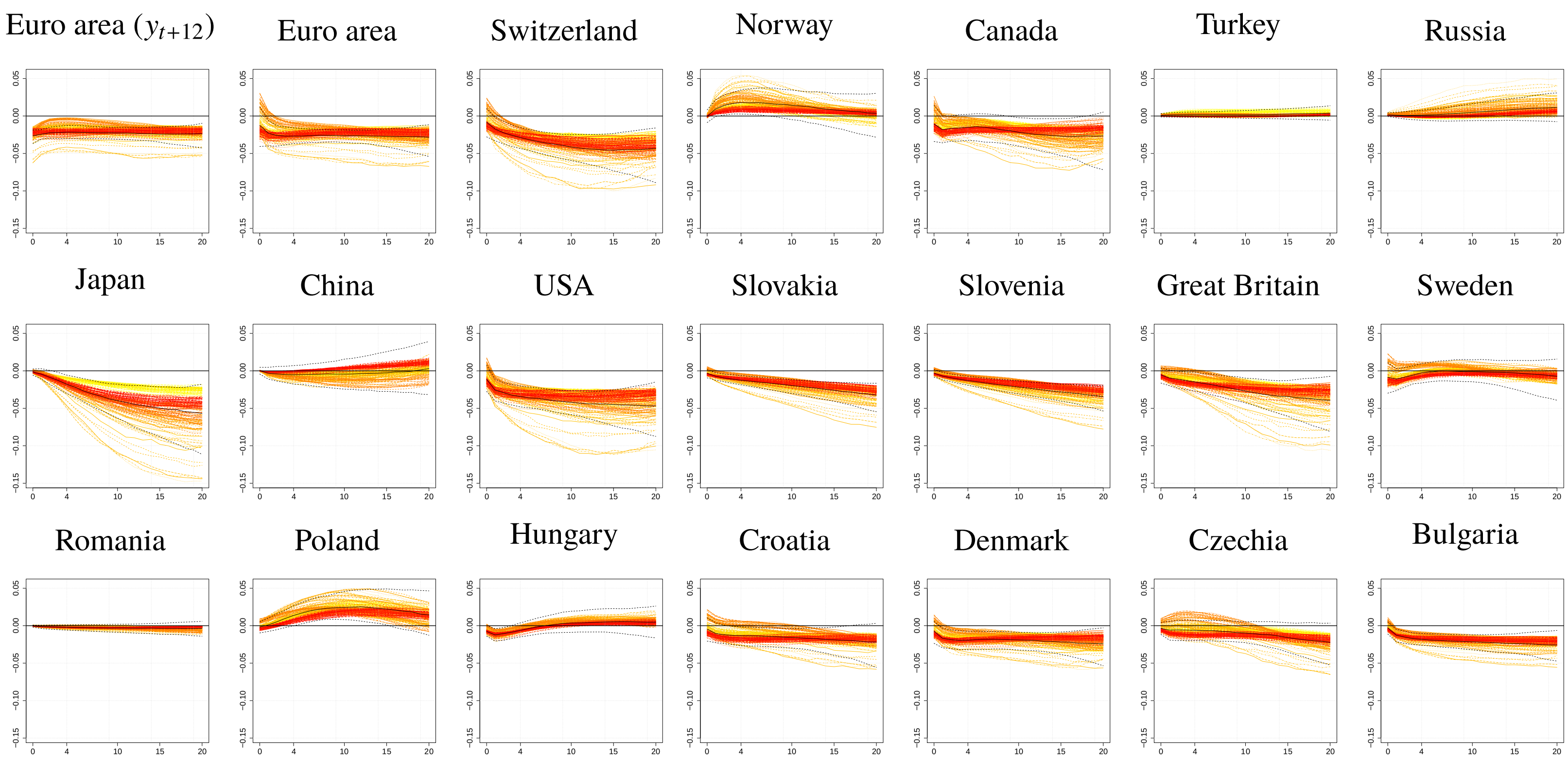
Notes: The plot shows posterior median responses over the sample period. Light yellow responses correspond to the beginning of the sample (i.e., 2001m1), dark red responses to the end (i.e., 2018m6). The black line corresponds to the posterior median of the time averaged response along with 68% credible bounds.

Fig. C.7: Responses of output growth to a euro area forward guidance tigthening



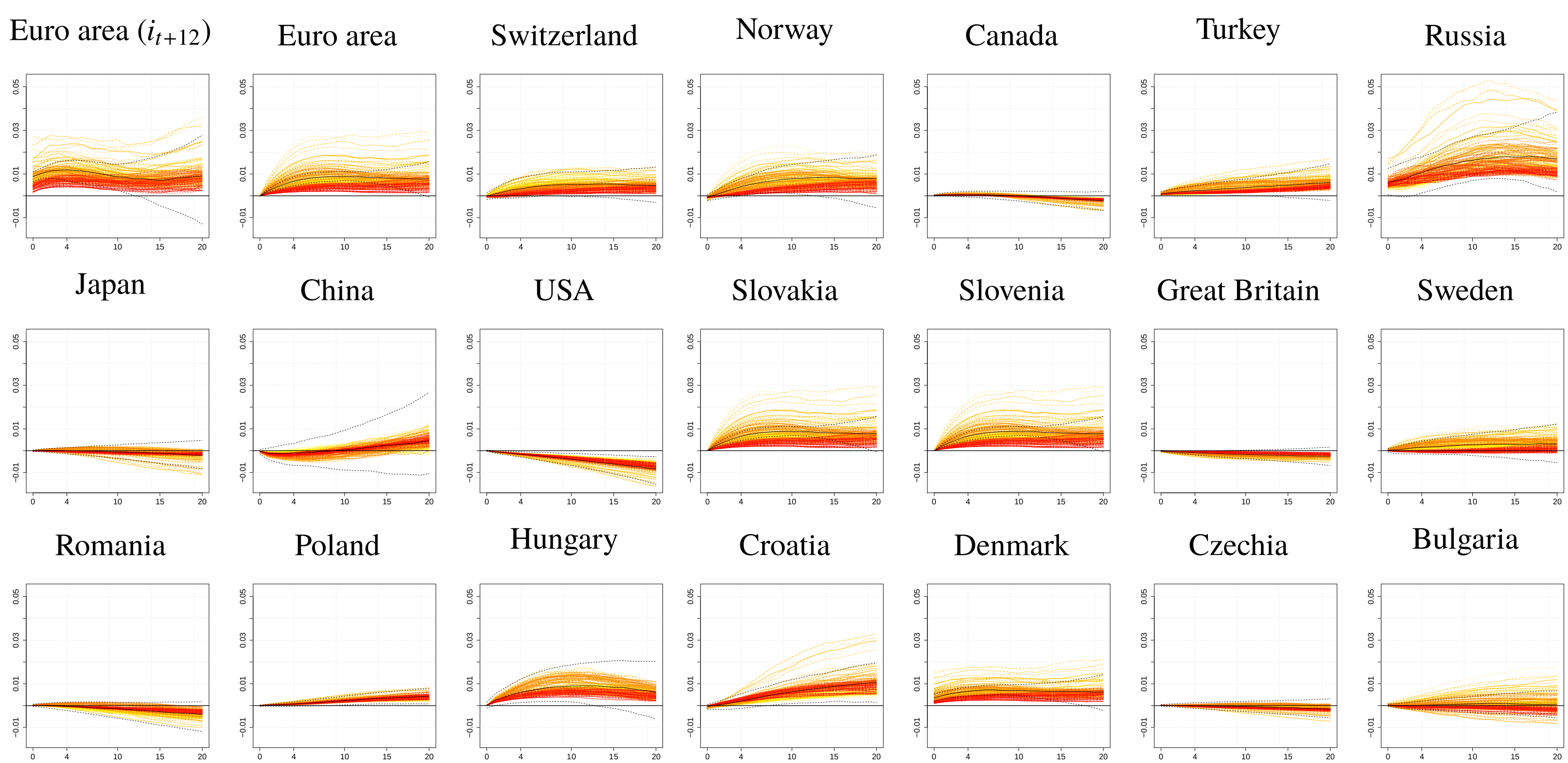
Notes: The plot shows posterior median responses over the sample period. Light yellow responses correspond to the beginning of the sample (i.e., 2001m1), dark red responses to the end (i.e., 2018m6). The black line corresponds to the posterior median of the time averaged response along with 68% credible bounds.

Fig. C.8: Responses of inflation to a euro area forward guidance tigthening



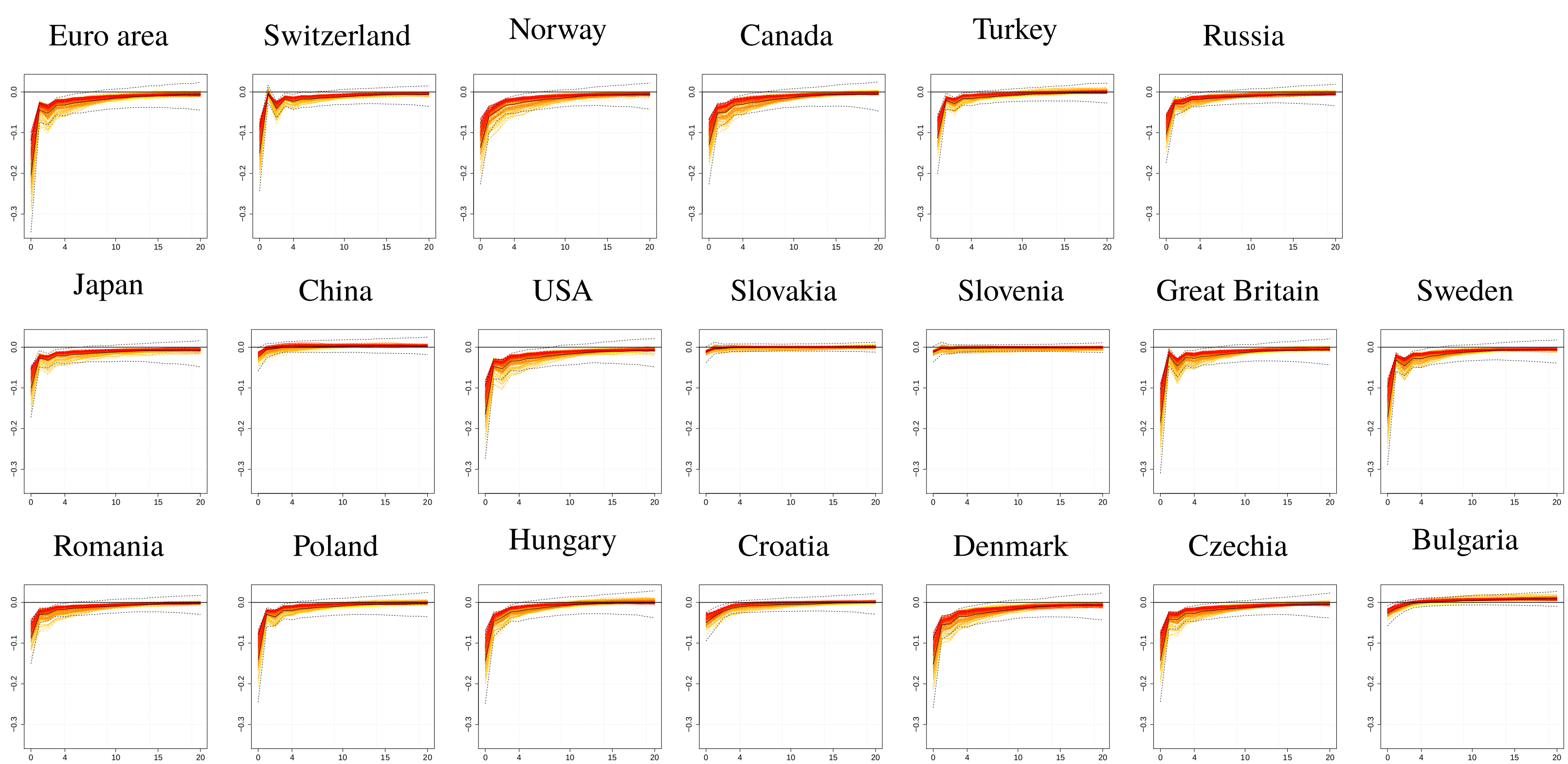
Notes: The plot shows posterior median responses over the sample period. Light yellow responses correspond to the beginning of the sample (i.e., 2001m1), dark red responses to the end (i.e., 2018m6). The black line corresponds to the posterior median of the time averaged response along with 68% credible bounds.

Fig. C.9: Responses of short-term interest rates to a euro area forward guidance tigthening



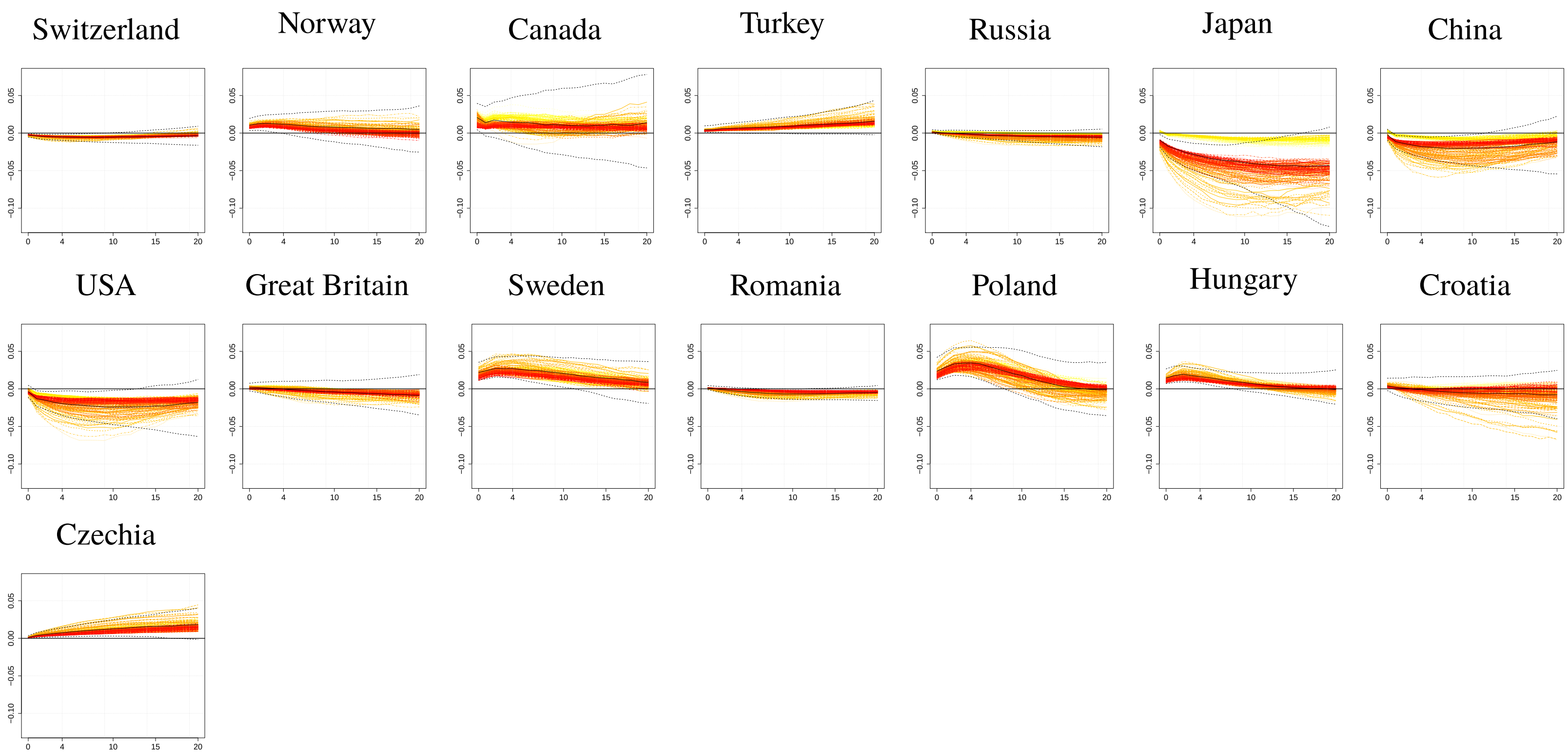
Notes: The plot shows posterior median responses over the sample period. Light yellow responses correspond to the beginning of the sample (i.e., 2001m1), dark red responses to the end (i.e., 2018m6). The black line corresponds to the posterior median of the time averaged response along with 68% credible bounds.

Fig. C.10: Responses of equity returns to a euro area forward guidance tigthening



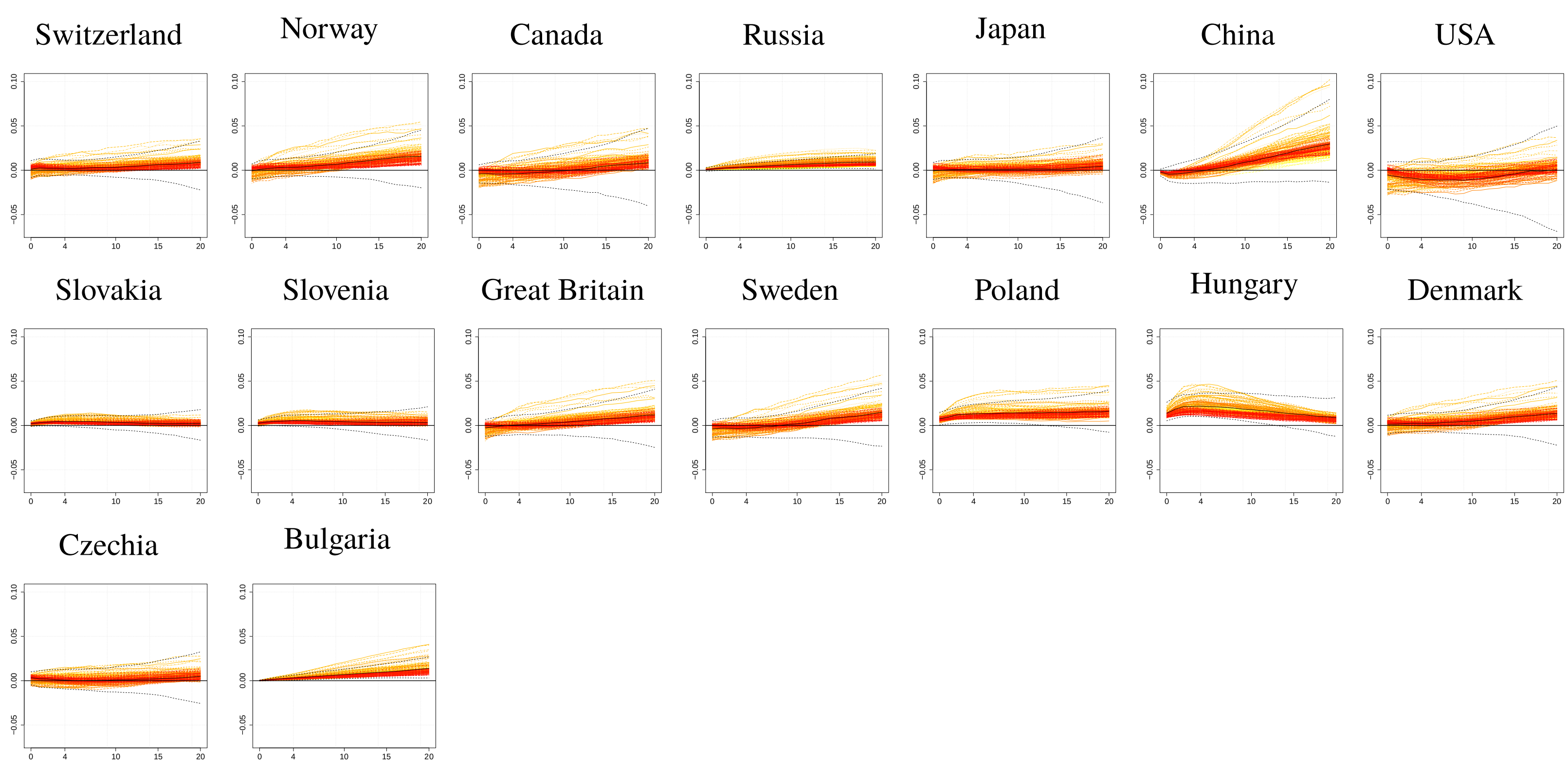
Notes: The plot shows posterior median responses over the sample period. Light yellow responses correspond to the beginning of the sample (i.e., 2001m1), dark red responses to the end (i.e., 2018m6). The black line corresponds to the posterior median of the time averaged response along with 68% credible bounds.

Fig. C.11: Responses of exchange rate changes to a euro area forward guidance tigthening



Notes: The plot shows posterior median responses over the sample period. Light yellow responses correspond to the beginning of the sample (i.e., 2001m1), dark red responses to the end (i.e., 2018m6). The black line corresponds to the posterior median of the time averaged response along with 68% credible bounds.

Fig. C.12: Responses of long-term interest rates to a euro area forward guidance tigthening



Notes: The plot shows posterior median responses over the sample period. Light yellow responses correspond to the beginning of the sample (i.e., 2001m1), dark red responses to the end (i.e., 2018m6). The black line corresponds to the posterior median of the time averaged response along with 68% credible bounds.