Uncertainty in Monetary Unions*

Benjamin Born, Luis Huxel,
Gernot J. Müller, and Johannes Pfeifer

November 2024

Abstract

A monetary union shapes how uncertainty impacts the economy: it does not alter the transmission of common uncertainty shocks, but significantly *dampens* the adverse effects of country-specific shocks. We establish this result based on time series data for 17 euro area countries and 13 countries with flexible exchange rates. To rationalize it, we propose a model of a monetary union where monetary policy responds to common shocks but not country-specific ones, as each member country is small. The union dampens the effect of country-specific shocks by providing a nominal anchor in the face of country-specific uncertainty, thereby eliminating price level risk.

Keywords: Uncertainty shocks, exchange rate regime, monetary policy, Monetary union, price level risk, nominal anchor, euro area
JEL-Codes: F41, F45, E44

*Born: Frankfurt School of Finance & Management, CEPR, CESifo, and ifo Institute, b.born@fs.de, Huxel: University of Tübingen, luis.huxel@uni-tuebingen.de, Müller: University of Tübingen, CEPR, and CESifo, gernot.mueller@uni-tuebingen.de, Pfeifer: Bundeswehr University Munich, johannes.pfeifer@unibw.de. We thank seminar participants at various conferences and seminars and our discussant Katrin Rabitsch for very helpful comments and suggestions. The usual disclaimer applies.

1 Introduction

Joining a monetary union has costs and benefits. The most significant benefit is often considered to be long-term or permanent: the elimination of a potential inflation bias through the 'nominal anchor' that the union provides (Alesina and Barro, 2002). Similarly, increased trade integration is a first-order benefit of a monetary union. On the other hand, costs are expected to materialize in a cyclical fashion because countries in a monetary union lack an independent currency and exchange rate flexibility vis-à-vis the other members of the union. This reduces their ability to cope with country-specific shocks, a central tenet of optimum currency area theory (Mundell, 1961). However, the distinction between permanent and cyclical implications of union membership only goes so far. For instance, as a monetary union fosters trade integration, it also alters business cycle co-movement and hence the incidence of country-specific shocks (Krugman, 1993).

Likewise, the nominal anchor provided by the union not only matters for inflation in the long run; it also shapes short-run fluctuations. We establish this point in the present paper as we contrast the effects of economic uncertainty and the associated price level risk on the countries of the euro area (EA) and on countries with flexible exchange rates. First, we present new evidence by estimating a structural Bayesian vector autoregression (VAR) on time series for 30 countries. Our main result is that the exchange rate regime does not matter for the effect of common uncertainty shocks, but it does matter for how country-specific uncertainty plays out. Somewhat surprisingly, the effect of country-specific shocks is much weaker in EA countries than in countries with floating exchange rates.

Second, to shed light on this result, we put forward a two-country model of a monetary union in which the domestic economy is small enough not to influence the common price level and estimate it by matching the empirical impulse responses to a common uncertainty shock. We evaluate the model predictions for the effect of a country-specific uncertainty shock: they are indeed weaker compared to a counterfactual in which we assume flexible exchange rates. The adjustment of the domestic price level explains this result. In a monetary union, it is anchored by the union-wide level, to which it converges to restore purchasing power parity. Therefore, increased uncertainty does not translate into long-run uncertainty about the price level—the union limits price level risk

The evidence we establish is based on quarterly data spanning the period from the euro's launch in 1999 to 2022. We allow for country-level heterogeneity by first estimating the VAR separately for each of the 30 countries in our sample. In a second step, we compute results for the median EA economy and the median floater by drawing from the posterior distribution of the estimates for each set of countries. We use realized national stock market volatility as the uncertainty indicator, following Bloom's (2009) seminal work. Our findings for the EA remain robust when we instead consider a forecast error-based uncertainty measure à la Jurado et al. (2015), compiled for the EA by Comunale and Nguyen (2023).

To identify uncertainty shocks, we follow Baker et al. (2016) and others by employing a recursive scheme, where the uncertainty indicator is ordered before the macroeconomic variables included in the VAR.[1] Our main interest is to identify country-specific and common uncertainty shocks separately. To do so, we first isolate the country-specific component in the uncertainty indicator for each country using principal component analysis. We then estimate the VAR with country-specific volatility ordered first and total volatility second. In this way, we allow total volatility to be driven by both types of shocks, but restrict country-specific volatility to being driven only by country-specific uncertainty shocks.

Uncertainty shocks matter: jointly they account for 20-30 percent of business cycle fluctuations, with common shocks responsible for the bulk. These common shocks adversely affect economic activity regardless of the exchange rate regime. They also lower inflation, as in previous work on "closed" economies (Leduc and Liu, 2016), and induce a decline in the policy rate, again independent of the exchange rate regime. In contrast, the exchange rate regime has a first-order effect on the transmission of country-specific uncertainty shocks. They lower only the output of floaters, not that of the median EA economy. And while floaters raise the policy rate in the face of higher inflation, the policy rate for the median EA economy is unresponsive—consistent with the notion that union membership constrains monetary policy. And yet, economic activity in the median EA economy is well insulated from country-specific uncertainty shocks.

[1]Several alternative strategies for identifying uncertainty shocks have been proposed, based, for instance, on narrative restrictions, sign restrictions, or external instruments (e.g. Piffer and Podstawski, 2018; Redl, 2020; Ludvigson et al., 2021). However, these approaches are more demanding regarding data and thus difficult to implement in our country panel. They are also not without their own caveats (see, for instance, Kilian et al., 2022).

To understand this result, we develop a model of a monetary union that extends the model of Basu and Bundick (2017) to a two-country setting. The countries are isomorphic except in two key respects. First, "Home" is small and does not affect the rest of the union, while "Foreign" is large and operates as a de facto closed economy. Second, the countries differ in the incidence of shocks. There are uncertainty shocks specific to Home and common uncertainty shocks that affect both countries alike. In each case, the shock widens the distribution from which (demand) shocks are drawn, without changing the mean. Monetary policy adjusts interest rates in response to union-wide inflation and output growth, for which Home is irrelevant due to its size.

We estimate the model by matching its predictions for the effects of common uncertainty shocks to the time series evidence for the EA. It turns out that the model is not only able to match the empirical impulse responses for the median EA economy to a common uncertainty shock under reasonable parameter values. It also predicts, consistent with the evidence, that country-specific uncertainty shocks have much weaker effects on countries in the monetary union. To understand how the monetary union shapes the transmission of both shocks, we compare the baseline scenario of a monetary union with a counterfactual in which Home operates outside the union and conducts monetary policy independently. In this case, the model correctly predicts that the effect of a country-specific shock is larger, while the effect of a common shock is unchanged relative to the baseline of union membership.

Thus, union membership is crucial for the transmission of country-specific uncertainty shocks, but not in the way traditionally expected. To see why, consider the transmission of the same shock when Home again conducts its own independent monetary policy outside the union. Suppose, however, that instead of targeting inflation and output growth, Home adjusts interest rates to stabilize the domestic price level. In this case, the response to country-specific shocks is basically indistinguishable from what happens under union membership. This testifies to the importance of the union-wide price level as an effective nominal anchor in the transmission of country-specific uncertainty shocks. Intuitively, when monetary policy targets inflation under flexible exchange rates, the price level exhibits a unit root and can drift arbitrarily far from its initial value, introducing price level risk. This risk is eliminated when Home is part of the monetary union.

Under union membership, the price level of the union provides a nominal anchor in the face of country-specific shocks due to purchasing power parity (PPP). And while the model allows for sizable deviations from PPP in the short run, PPP holds in steady state, consistent with evidence for the EA (Bergin et al., 2017). Since the nominal exchange rate is irrevocably fixed in the monetary union, domestic prices must eventually converge back to the union level to restore PPP after a country-specific shock.[2] Note that this mechanism will also be at play when it comes to policy uncertainty which impacts economic activity adversely (Mumtaz and Zanetti, 2013; Born and Pfeifer, 2014a; Fernández-Villaverde et al., 2015). Its adverse impact will also be reduced by the nominal anchor provided by the union—to the extent that it is country-specific.

We further analyze the impact of price level risk on the economy by decomposing the time-varying risk wedges due to uncertainty shocks in the spirit of Bianchi et al. (2023). We selectively shut off these wedges, solving variants of the model in which certain forward-looking equations are restricted to their log-linear approximation. Our simulation results suggest that the anchoring of price level expectations in a monetary union—and the resulting reduction in inflation risk—is particularly important for households' saving decisions.

The paper is organized as follows. In the remainder of the introduction, we place the paper in the context of the literature and outline its contribution. The next section contrasts time series evidence on the impact of uncertainty shocks on economic activity in a country operating in a monetary union with a country operating a floating exchange rate. Section 3 puts forward our theoretical DSGE model, which we estimate and use to run counterfactuals in Section 4. A final section concludes.

Related Literature. The idea that joining a monetary union removes the inflation bias by tying one's hands is already formalized by Giavazzi and Pagano (1988). Corsetti et al. (2013) and Groll and Monacelli (2020) emphasize that this is also important for the transmission of shocks. Our analysis differs in that it focuses on uncertainty shocks and is based on actual time series evidence. Likewise, we build on previous work that examines the macroeconomic effects of uncertainty, surveyed

[2] In fact, relative PPP is sufficient for the union price level to act as a nominal anchor in the face of shocks.

in Bloom (2014) and Castelnuovo (2023). This literature examines how the effects of uncertainty shocks are shaped by monetary policy (in a closed economy context), and how uncertainty plays out in the open economy. Against this background, our particular contribution is to highlight the importance of monetary policy in the open economy—and, in particular, the exchange rate regime—for the transmission of uncertainty shocks.

As such, our paper is distinct from, but related to, three strands of the literature. First, there is work on the relevance of constraints on monetary policy for the transmission of uncertainty shocks (e.g. Johannsen, 2014; Caggiano et al., 2017; Nakata, 2017). Andreasen et al. (2024), Pellegrino et al. (2023), and Fasani and Rossi (2018) show that the conduct of systematic monetary policy greatly matters for the transmission of uncertainty shocks to the economy in closed economy models. This strand of the literature also explores the role of endogenous uncertainty (Plante et al., 2018).

Second, there is work that focuses on the international dimension of uncertainty shocks. In particular, Mumtaz and Theodoridis (2017) and Caggiano and Castelnuovo (2023) decompose uncertainty measures into global, regional, and country-specific factors. Albagli et al. (2024) and Georgiadis et al. (2024) study the role of exchange rates in the transmission of global uncertainty shocks. Lakdawala et al. (2021) study the spillovers of U.S. monetary policy uncertainty, while Meinen and Roehe (2017) investigate the effect of uncertainty shocks on different EA countries. The transmission of uncertainty shocks, in particular interest rate shocks, has also been studied extensively in small open economy models (Fernández-Villaverde et al., 2011; Born and Pfeifer, 2014b; Başkaya et al., 2013; Kollmann, 2016; Johri et al., 2022). In a parallel strand, the literature also examines how uncertainty affects firms' export decisions (Handley and Limão, 2017; Carballo et al., 2022; Fernandes and Winters, 2021).

Finally, there is the question of whether and to what extent the exchange rate regime makes a difference for the transmission of shocks, both domestic and external (Bayoumi and Eichengreen, 1994; Broda, 2004; Giovanni and Shambaugh, 2008). The evidence presented in recent papers is inconclusive (Corsetti et al., 2021; Fukui et al., 2023), although there are cases where the exchange rate regime clearly matters, just like in our analysis—not only at the aggregate but also at the household level (Bayer et al., 2024; Born et al., 2013, 2024).

2 Time series evidence

In this section, we provide time series evidence on whether and, if so, how the exchange rate regime matters for the transmission of uncertainty shocks. To do so, we first estimate a small-scale panel (Bayesian) VAR on country-level data for 30 economies and compare the median effect of uncertainty shocks in the 17 countries that are members of the EA with the median effect in the 13 countries that float their exchange rate. To highlight the role of the exchange rate regime, we distinguish in each case the effect of common and country-specific uncertainty shocks. To further inform our model-based analysis, we zoom in on the adjustment dynamics in the EA and estimate a larger VAR with additional time series.

2.1 Sample and data

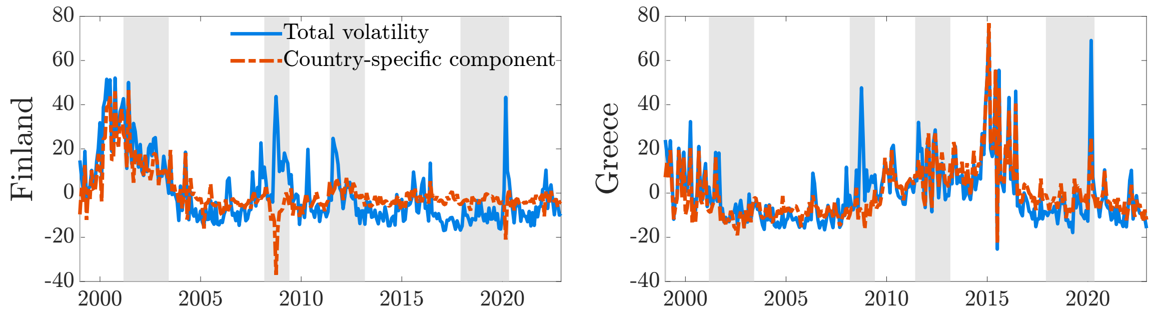
Our sample starts in 1999Q1 with the launch of the euro and runs until 2022Q4. For the EA, it covers time series data for 17 countries: Austria, Belgium, Cyprus, Estonia, Finland, France, Germany, Greece, Ireland, Italy, Latvia, Lithuania, Luxembourg, the Netherlands, Portugal, Slovenia, and Spain. Initially, the EA consisted of 11 of these countries. For the countries that joined later, we use data from the period when their currency was pegged to the euro in the run-up to accession.[3] The other part of our sample consists of 13 non-European countries with a floating exchange rate: Australia, Brazil, Canada, Chile, Colombia, India, Israel, Mexico, New Zealand, Russia, South Africa, South Korea, and the United States. For these countries, we restrict the sample to periods when they had a floating exchange rate. Appendix A provides a list of the starting dates for each country.

To identify uncertainty shocks, we use a time series of realized stock market volatility as a proxy for the underlying uncertainty.[4] Importantly, for each country, we distinguish a country-specific component from the total volatility observed in that country. We measure total volatility at a monthly frequency by the standard deviation of annualized daily returns for a country's Datastream stock market performance index. We compute the country-specific volatility component by

[3]Our sample does not include Malta and Slovakia because the stock market data necessary to measure volatility are not available for the full sample period. The 20th member country of the EA, Croatia, only joined after the end of our sample period in 2023.

[4]As a robustness check for the EA, we use a forecast error-based macroeconomic uncertainty measure (Jurado et al., 2015), compiled for the EA by Comunale and Nguyen (2023); see Sec. 2.4.

Figure 1: Monthly stock market volatility in Finland and Greece



Notes: Monthly stock market volatility of Finland (left panel) and Greece (right panel) in percent (demeaned); realized total (solid blue line) and country-specific (dashed red line) volatility, based on annualized volatility of daily returns of the market performance index (Datastream: TOTMK**(RI)). Shaded areas mark EA recessions according to OECD indicators. Sample: 1999M1–2022M12.

purging a country's total volatility of the first principal component extracted from the total volatility of a large panel of countries.[5] Finally, we aggregate the monthly series to a quarterly frequency by averaging the observations of a given quarter.[6]

Figure 1 visualizes the result for Finland (left panel) and Greece (right panel), contrasting total volatility (blue solid line) with the country-specific component (red dashed line) over our sample period. In both countries, total volatility spiked during the 2008/09 financial crisis and at the onset of the COVID-19 pandemic in 2020. In both instances, the country-specific components did not increase, consistent with the notion that these were global events. In Finland the country-specific component actually decreased in 2008/09, suggesting that the country was relatively less exposed to the financial crisis. However, the Finland-specific component was particularly high after the bursting of the dot-com bubble in the early 2000s, reflecting Nokia's prominent position in the Finish economy and stock market. Similarly, the increase in volatility in Greece during the debt crisis in the 2010s was almost entirely driven by the country-specific component. Similar patterns emerge for all countries in our EA sample; see Appendix-Figure B.1.

[5]In addition to the 30 countries on which we estimate the VAR, this panel includes countries with intermediate exchange rate regimes, see Appendix A for details.

[6]As an alternative to relying on realized returns, implied volatility can be extracted from option prices. While conceptually appealing, the data required for this approach are not consistently available across EA member countries. In practice, implied and realized volatility co-move strongly. Their correlation of 0.88 in U.S. data (Born and Pfeifer, 2021) allowed Bloom (2009) to concatenate realized and implied volatility measures in his seminal study.

We compare our volatility measure with the Economic Policy Uncertainty (EPU) index compiled by Baker et al. (2016) for selected EMU countries. For the four countries in our EA subsample for which the EPU country index is available, it comoves strongly with our stock market-based country-specific component, see Figure B.2. This suggests that country-specific volatility reflects country-specific uncertainty that is partly, but not exclusively, due to economic policy.

2.2 Time series framework

For each of the 30 countries in our sample, we estimate a parsimonious Bayesian VAR that includes, in addition to the two time series for total and country-specific volatility, the log of real GDP per capita, inflation, and the policy rate. For EA countries, we use the Wu and Xia (2016) shadow rate as a proxy for the policy rate, since the ECB was constrained by the zero lower bound for much of our sample period. Appendix A provides further details on the data.

Since we estimate the VAR country by country, we account for dynamic heterogeneity (Pesaran and Smith, 1995; Canova and Ciccarelli, 2013). Given the country-level posterior distributions, we synthesize the evidence by computing results for the 'median economy', following Degasperi et al. (2023): across the groups of EA countries and floaters, for each country we take a random draw from the posterior distribution and compute the cross-country median of the statistic of interest, such as the impulse response function of a given variable to a given shock at a given horizon. Repeating this procedure 1,000 times allows us to characterize the posterior distribution for the median EA country and the median floating exchange rate economy.

Formally, we estimate the following VAR model for each country:

$$Y_t = \mu_0 + \mu_1 D_t + \alpha_0 t + \alpha_1 t D_t + A(L) Y_{t-1} + \nu_t \ , \tag{2.1}$$

where Y_t is a 5×1-vector of endogenous variables, $A(L)$ is a lag polynomial of degree $p = 4$, and $\nu_t \overset{iid}{\sim} \mathcal{N}(0, \Sigma)$. D_t is a dummy variable equal to one starting with 2020:Q1 to capture the shifts in level and trend observed after the onset of the COVID pandemic; μ_0, μ_1 and α_0, α_1 are constants and time trends, respectively.

As discussed in the introduction, we follow Bloom (2009) and Jurado et al. (2015) and identify uncertainty shocks based on a recursive ordering. That is, we assume a lower-triangular matrix B that maps structural shocks ε_t into reduced-form innovations ν_t, $\varepsilon_t = B\nu_t$ such that $\Sigma = BB'$. In the spirit of Basu and Bundick (2017) and Baker et al. (2016), who order the uncertainty proxy first, we put the country-specific component first, followed by total volatility and then all other variables, with the shadow rate ordered last. We identify country-specific and common uncertainty shocks jointly, assuming that both shocks potentially drive total volatility, while country-specific volatility is driven only by country-specific uncertainty shocks. We also allow uncertainty shocks to affect the other variables in the VAR contemporaneously, but exclude other shocks from affecting volatility within the quarter.

We use a shrinking prior of the Independent Normal-Inverse Wishart type (Kadiyala and Karlsson, 1997), with mean and precision derived from a Minnesota-type prior (Litterman, 1986; Doan et al., 1984). We write the vector of stacked coefficients as $\beta = vec([\mu_0\ \mu_1\ \alpha_0\ \alpha_1\ A_1\ \ldots\ A_p]')$ and assume a normal prior: $\beta \sim N(\underline{\beta}, \underline{V})$. For its mean $\underline{\beta}$, we assume that the variables follow a univariate AR(1)-model with mean of 0.9, while all other coefficients are 0. The prior precision $\underline{V}$ is a diagonal matrix with the highest precision for the first lag and exponential decay for the remaining lags. The cross terms are weighted according to the relative size of the error terms in each equation. At the same time, a rather diffuse prior is used for the deterministic and exogenous terms. The diagonal element corresponding to the jth variable in equation i, $\underline{V}_{i,jj}$ is:

$$\underline{V}_{i,jj} = \begin{cases} \frac{\underline{a}_1}{r^2}, & \text{for coefficients on own lag } r \in \{1, \ldots, p\}\,, \\ \frac{\underline{a}_2 s_i^2}{r^2 s_j^2}, & \text{for coefficients on lag } r \in \{1, \ldots, p\} \text{ of variable } j \neq i\,, \\ \underline{a}_3 s_i^2, & \text{for coefficients on exogenous or deterministic variables}\,, \end{cases} \tag{2.2}$$

where s_i^2 is the OLS estimate of the error variance of an AR(p)-model with constant and trend estimated for the ith variable (see Litterman, 1986). We set $\underline{a}_1 = 0.1, \underline{a}_2 = 0.1$ and $\underline{a}_3 = 10^4$. The prior error covariance is assumed to follow $\underline{\Sigma} \sim IW(\underline{S}, \underline{\nu})$, where $\underline{\nu} = 10$ are "pseudo-observations", corresponding to ≈ 10 percent of the observations, and $\underline{S}$ is the OLS covariance matrix.

In the Gibbs sampler, we use 25,000 draws, discarding the first 5,000 draws as a burn-in.[7] Given the shortness of our sample, we prefer the 68% highest posterior

[7] Raftery and Lewis (1992) convergence diagnostics suggest that this is sufficient for convergence.

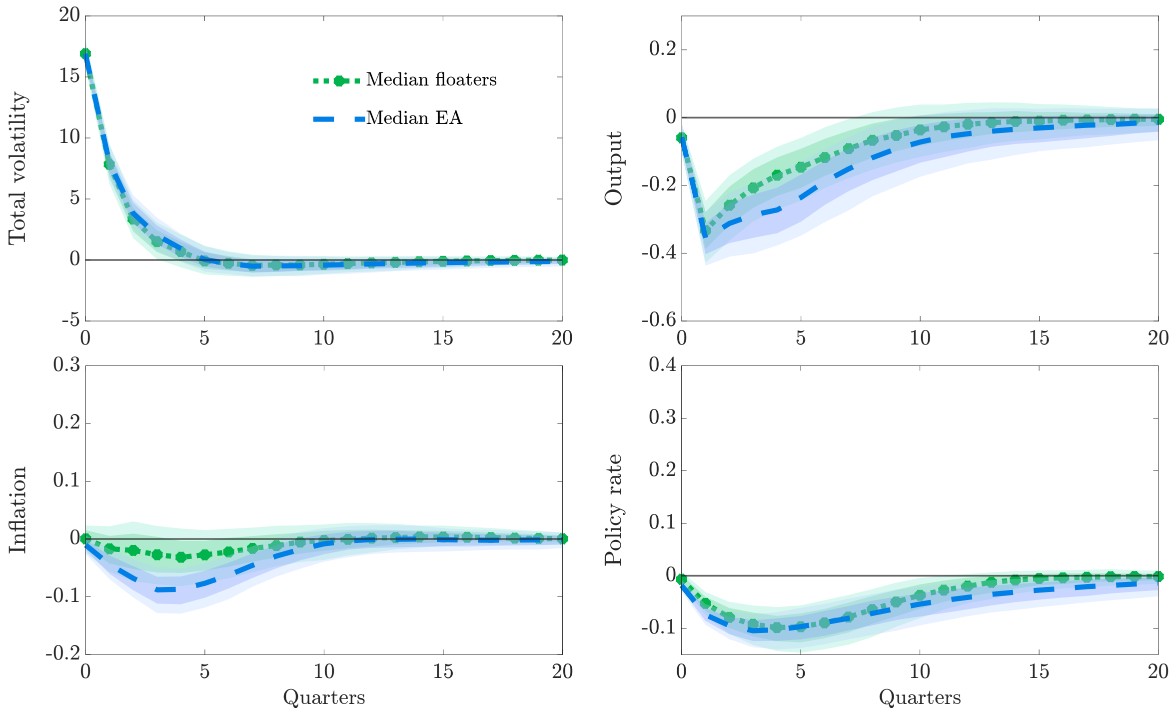
density intervals (HPDIs), but report 90% HPDIs as well. As a practical matter, we z-score the data (including the trend) to avoid numerical problems arising from under-/overflow in the posterior computations involving sums of squares. We also impose a stability condition on our VAR by drawing from the conditional distribution for β until the modulus of all eigenvalues of the companion form matrix is less than 1.

2.3 Results

In what follows, we focus on the results for the median economy, that is, we aggregate across countries based on the posterior distribution of the estimated VAR models—once for the countries within the EA and once for the countries that float their exchange rate. Figure 2 shows the adjustment dynamics to a common uncertainty shock, contrasting the impulse responses for the median EA economy (dashed blue line) and for the median floater (dotted green line with octagonal markers). The shaded areas indicate 68% and 90% HPDIs. Here and in what follows, we normalize the size of the shock so that its impact on total volatility is the same as that of a country-specific one-standard-deviation shock in the median EA economy. The response of total volatility is shown in the top-left panel of the figure, measured as percentage deviation from the unconditional mean, as in Basu and Bundick (2017). The horizontal axis measures quarters throughout.

The upper-right panel of Figure 2 shows the response of output. It is virtually identical for the median EA economy and the median floater. We observe the strongest effect after one quarter when output is reduced by about 0.35 percent, in both economies. Thereafter, output recovers, but only gradually. The bottom-left panel shows the response of inflation, which again displays similar adjustment dynamics in both cases, although the decline is somewhat stronger in the EA. A declining inflation response to an uncertainty shock is consistent with earlier evidence by Leduc and Liu (2016). Finally, looking at the bottom-right panel, we observe that the policy rate falls in response to the common uncertainty shock, also in the median EA economy—consistent with the notion that EA-wide monetary policy responds to a shock that is common to all countries in the EA. Overall, the pattern of adjustment is consistent with previous work based on aggregate data for the U.S. (Basu and Bundick, 2017). In the context of our analysis, it is noteworthy

Figure 2: Adjustment to common uncertainty shocks in median economy

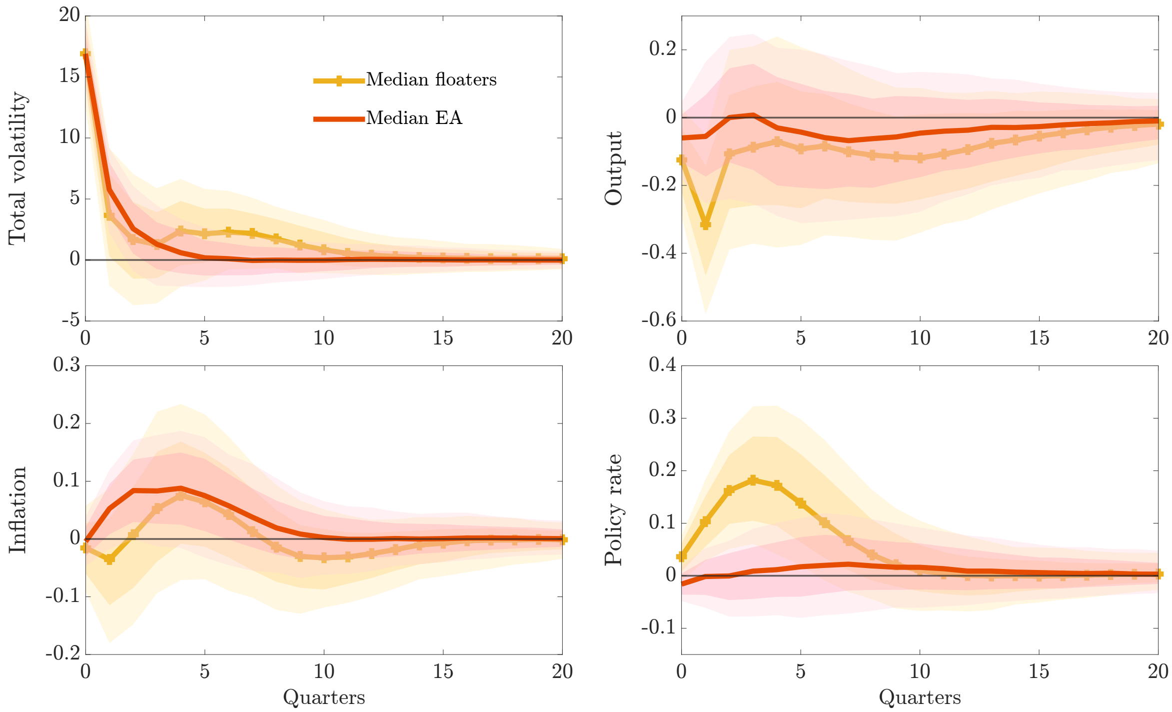


Notes: Impulse responses to common uncertainty shocks in the EA (blue dashed line) and among global floaters (green dotted line with octagonal markers). Shock sizes rescaled so that the median impact on total volatility equals that of one-standard deviation country-specific uncertainty shock in the EA. Shaded areas indicate point-wise 68% (dark) and 90% (light) HPDIs, respectively. Horizontal axis measures time in quarters, vertical axis measures deviations from pre-shock level in percent, except for inflation and policy rate (ppts). Country-specific volatility is included in the VAR, but not shown here.

that the adjustment dynamics to a common shock are essentially the same for the median EA economy and the median floater. This is intuitive and in line with theory: for countries that are not systematically different, the effect of common shocks will not depend on the exchange rate regime.

Country-specific shocks are a different matter. In this case, the exchange rate regime is bound to have a first-order effect. And indeed, as we shift focus and look at the adjustment induced by country-specific uncertainty shocks in Figure 3, we observe some notable differences in the median EA economy (solid red lines) and the median floater (solid yellow lines with plus-shaped markers). The figure is organized in the same way as Figure 2. The top-left panel shows the response of

Figure 3: Adjustment to country-specific uncertainty shocks in median economy



Notes: Impulse responses to country-specific uncertainty shocks in the EA (red solid line) and among global floaters (yellow solid line with plus-shaped markers). Shock size rescaled so that the median impact on total volatility equals that of one-standard deviation country-specific uncertainty shock in the EA. Shaded areas indicate point-wise 68% (dark) and 90% (light) HPDIs, respectively. Horizontal axis measures time in quarters, vertical axis measures deviations from pre-shock level in percent, except for inflation and policy rate (ppts). Country-specific volatility is included in the VAR, but not shown here.

total volatility, which is by construction the same as for the common shocks.

Our main result concerns the response of output, shown in the top-right panel. Here we observe a sharp, albeit transitory, decline in output for the median floater, comparable to, though somewhat weaker than, the case of a common shock. Instead, output does not move significantly in the median EA economy. This result is surprising given the received wisdom going back at least to Mundell (1961): after all, a country operating within a monetary union lacks the ability to adjust monetary policy in the face of country-specific shocks. Thus, if anything, one might have expected a stronger impact of country-specific uncertainty shocks in the median EA economy. And yet, we find no significant output response here.

The response of inflation, shown in the bottom-left panel, is quite similar in both cases. Finally, we note that the response of the policy rate in the median EA country is flat, supporting the notion that we are indeed capturing country-specific shocks to which the common monetary policy in the EA does not respond. Instead, the median floater raises its policy rate, possibly in response to the rise in inflation and suggestive of the monetary autonomy that floaters enjoy.

We now zoom in on the transmission of country-specific and common uncertainty shocks in the euro area and re-estimate the VAR with consumption and investment as additional variables on time series data for the EA countries. The result will serve as a vital input for the identification of some of our model parameters in Section 3 below. Figure 4 shows the results for this extended VAR. As before, the solid (red) line represents the adjustment to the country-specific uncertainty shock in the median EA economy. The dashed (blue) line shows the responses to a common shock.

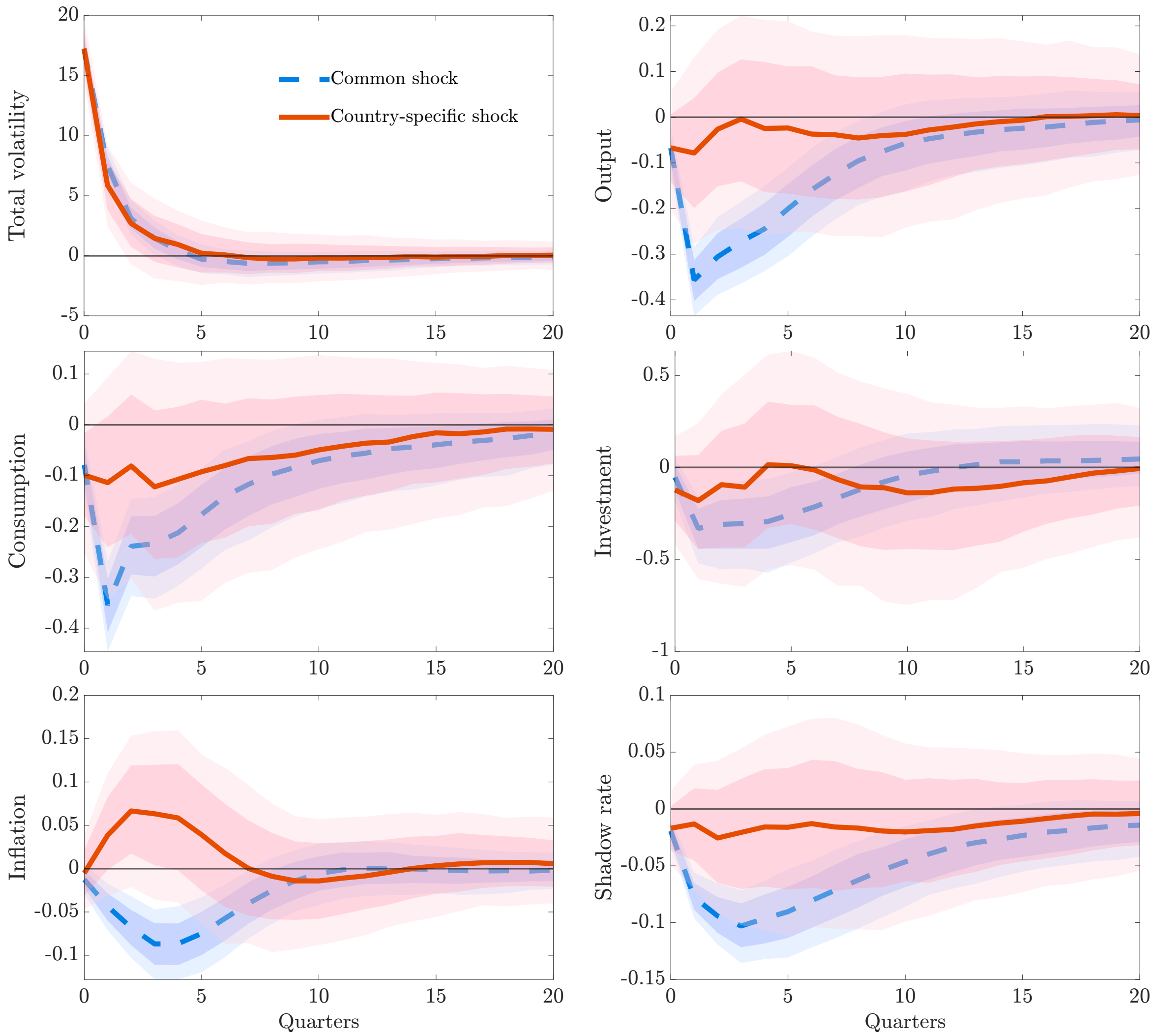
Consistent with the results above, we find that output (top-right panel) declines much more strongly after a common than after a country-specific uncertainty shock. We find similar patterns for consumption (middle-left panel) and investment (middle-right panel), although the difference is less extreme for the latter. Importantly, even when controlling for the two additional variables, the policy rate still does not respond significantly to a country-specific shock.[8]

We also assess the (relative) importance of uncertainty shocks for business cycle fluctuations in the median EA economy. To do so, we perform a forecast error variance decomposition (FEVD) based on the extended VAR, computed using the same sampling approach as for the IRFs, and report the results in Table 1. Focusing on the FEVD at a business cycle frequency of 20 quarters, we find that the two uncertainty shocks together account for more than 20 percent of the output fluctuations, with common shocks accounting for about two-thirds of this number. A similar pattern holds for the other variables, including the policy rate. Here, too, country-specific shocks appear to contribute to fluctuations. Note, however, that their effect on the shadow rate is generally insignificant, as Figure 4 shows.[9] In the appendix, we show the FEVD at the country level; see Figure B.7. As before, we

[8]Figures B.3 and B.4 in the appendix show the impulse responses for each of the 17 EA countries. Similarly, Figures B.5 and B.6 show the corresponding results for the floaters.

[9]In the FEVD, we nevertheless find a non-zero contribution because, while the country-level shocks have partly opposite effects that average out for our ‘median economy’ impulse responses, this is not the case for the FEVD, which does not consider the sign of responses to shocks.

Figure 4: The adjustment to uncertainty shocks in median EA economy



Notes: Impulse responses to one-standard-deviation country-specific (solid red line) and equally-scaled common (dashed blue line) uncertainty shock. Shaded areas indicate point-wise 68% (dark) and 90% (light) HPDIs, respectively. Horizontal axis measures time in quarters, vertical axis measures deviations from pre-shock level in percent, except for inflation and the shadow rate (ppts). The country-specific component of volatility is included in the VAR, but not shown.

find that the degree of heterogeneity at the country level is moderate. Table B.1 also shows the FEVD performed for the VAR of global floaters. We find that the role of uncertainty shocks is even larger there, accounting for about 29 percent of output fluctuations.

Table 1: Forecast error variance decomposition for median economy

	Country-specific uncertainty shock	Common uncertainty shock
Country-specific component	52.21 (46.8 , 57.58)	8.10 (6.33 , 10.28)
Total volatility	13.98 (11.98 , 16.04)	47.20 (43.14 , 51.46)
Output	7.69 (5.77 , 10.04)	15.43 (12.08 , 19.00)
Consumption	8.05 (6.18 , 10.53)	14.64 (11.52 , 18.13)
Investment	8.02 (6.19 , 10.29)	9.70 (7.48 , 12.61)
Inflation	8.39 (6.38 , 10.83)	11.94 (9.21 , 14.76)
Shadow rate	7.50 (5.53 , 9.89)	13.92 (10.48 , 18.19)

Notes: Contribution of country-specific (middle column) and common (right column) uncertainty shock to forecast error variance of each variable at horizon 20, in percent of total forecast error variance of that variable (with 68% HPDIs reported in parentheses).

2.4 Robustness

In what follows, we verify that our results for the euro area are robust to a number of alternative specifications. We briefly discuss these specifications and, to economize on space, present the results in the appendix. First, we address the concern that some countries in our sample are large enough to influence the common monetary policy in the EA by excluding the five countries that individually account for at least 5 percent of aggregate EA output (Germany, France, Italy, Spain, the Netherlands) when computing the median economy impulse responses. We find no meaningful difference from the baseline results, see Figure B.8.

Second, we want to consider the possibility that the milder effects of country-specific shocks are caused by fiscal stabilization. After all, in theory, countries in monetary unions may resort to fiscal policy to stabilize country-specific shocks (Galí and Monacelli, 2008). To explore whether a fiscal response can rationalize our findings, we include real per capita government consumption as an additional variable in the VAR. The result is clear: we do not find that government spending increases in response to uncertainty shocks. It tends to fall, but the response is generally insignificant, see Figure B.9.

Third, we include the level of the stock market index as an additional variable in the VAR model. In this way, we ensure that what we identify is a pure second-moment shock and rule out that our results are driven by a correlation between uncertainty shocks and a level shock of the opposite sign. We consider two different specifications. The first version follows Bloom (2009) and orders the log level of the stock market index as the first variable in the VAR. The second specification allows the stock market to react contemporaneously to uncertainty shocks by ordering it after the volatility measures. We do find that uncertainty shocks cause the stock market to fall, but our main results remain unaffected, see Figures B.10 and B.11.

Fourth, we verify that our results are not limited to the specific stock market-based measure of uncertainty and re-estimate our baseline VAR by replacing stock market volatility with the forecast error-based measure of macroeconomic uncertainty developed by Jurado et al. (2015) and provided by Comunale and Nguyen (2023) for all EA countries. Specifically, we use the 12-month-ahead uncertainty proxy for each country as the measure of total uncertainty and again proceed as above to isolate country-specific uncertainty. Due to the availability of the uncertainty measure, our sample here covers 2003Q3–2022Q4. This specification again confirms that the effects of country-specific shocks are weaker and not stronger than those of common shocks, and that only common shocks are accommodated by a significant monetary policy response, see Figure B.12.

Finally, we check the robustness of the floater response shown in Figures 2 and 3 by excluding the United States from the sample due to the unique position of the dollar as the dominant currency (Gopinath et al., 2020). As shown in Figures B.13 and B.14, this makes virtually no difference.

3 Model

We now put forward a model of monetary unions that features two countries: Home and Foreign. Home is small and does not affect the rest of the union as in Galí and Monacelli (2005, 2008). The rest of the union ("Foreign"), is a large economy that operates de facto as a closed economy but generates spillovers to

Home. The model structure in each country, and in particular the specification of uncertainty shocks, is based on Basu and Bundick (2017). We want to run a counterfactual with flexible exchange rates to understand how the monetary union affects the transmission of shocks. For this reason, we model the nominal exchange rate explicitly throughout, assuming that it is permanently fixed in the monetary union baseline. Home and Foreign are broadly symmetric, and we focus the exposition on Home, delegating details and derivations to Appendix C.1.

Formally, we develop our two-country setup by assuming that a fraction $n \in [0,1]$ of households and firms reside in Home and the rest in Foreign, with the global mass of firms and households normalized to unity. Later, we let $n \to 0$ as in Corsetti et al. (2021) to obtain a small open economy. We use the subscripts 'H' and 'F' to refer to domestic and foreign variables in Home, and an asterisk to refer to variables in Foreign.

3.1 Firms

Each firm in a given country produces a specific differentiated intermediate good, which is traded across borders and whose price is sticky in the producers' currency. A competitive final goods firm then uses a Dixit-Stiglitz technology to bundle these intermediate goods into a domestic composite, $Y_{H,t}$, and an imported composite $Y_{F,t}$. These, in turn, are combined to produce final goods, $\mathcal{F}_t$, which are used for consumption and investment:

$$\mathcal{F}_t = \left[[1-(1-n)v)^{\frac{1}{\eta}} \left(Y_{H,t}\right)^{\frac{\eta-1}{\eta}} + ((1-n)v)^{\frac{1}{\eta}} \left(Y_{F,t}\right)^{\frac{\eta-1}{\eta}} \right]^{\frac{\eta}{\eta-1}}, \tag{3.1}$$

where η is the trade price elasticity and $0 \leq v \leq 1$ measures the import content of final goods. To the extent that $v < 1$, there is 'home bias.' We refer to the price of final goods as the 'consumer price index' (CPI). It is given by

$$P_t = \left[(1-(1-n)v) \left(P_{H,t}\right)^{1-\eta} + ((1-n)v) \left(P_{F,t}\right)^{1-\eta} \right]^{\frac{1}{1-\eta}}. \tag{3.2}$$

Here, $P_{H,t}$ is the domestic producer price index (PPI), and $P_{F,t}$ is the price of imports which, under the law of one price, is given by $P^*_{F,t}\mathcal{E}_t$, where $P^*_{F,t}$ is the foreign currency price of imports and $\mathcal{E}_t$ is the nominal exchange rate, defined as the price of foreign currency in terms of domestic currency (equal to unity in the monetary union). We can then define the real exchange rate $\mathcal{Q}_t$ as the price of foreign goods in terms of the domestic final good (so that an increase amounts to a depreciation): $\mathcal{Q}_t \equiv \mathcal{E}_t P_t^* / P_t$.

Assuming that Home is small ($n \to 0$), expenditure minimization in Home and Foreign implies that the demand for a generic intermediate good $i \in [0, n]$ in Home, which sells at price $P_t(i)$, is given by

$$Y_t^d(i) = \left(\frac{P_t(i)}{P_{H,t}}\right)^{-\epsilon} \left(\frac{P_{H,t}}{P_t}\right)^{-\eta} \left[(1-v)\mathcal{F}_t + v\mathcal{Q}_t^{\eta}\mathcal{F}_t^*\right] \,, \tag{3.3}$$

where $\epsilon > 1$ is the elasticity of substitution between intermediate goods and $\mathcal{F}_t^*$ denotes the foreign final goods production.

To produce a differentiated intermediate good, $Y_t(i)$, a generic monopolistically competitive firm i executes the following production function:

$$Y_t(i) = \left(u_t(i) K_{t-1}(i)\right)^{\alpha} \left(N_t(i)\right)^{1-\alpha} - \Phi \,, \tag{3.4}$$

where $K_{t-1}(i)$ is the predetermined capital stock, $u_t(i)$ capital utilization, and $N_t(i)$ denotes labor input. $0 \leq \alpha \leq 1$ parameterizes the capital share. The fixed costs of production Φ ensure that profits are zero in the steady state. Adjusting investment is costly for firms and incurs a quadratic adjustment cost parameterized by $\phi_k > 0$. The law of motion for the capital stock is given by

$$K_t(i) = \left(1 - \delta_t(u_t)\right) K_{t-1}(i) + \left(1 - \frac{\phi_K}{2}\left(\frac{I_t(i)}{I_{t-1}(i)} - \delta_0\right)^2\right) I_t(i) \,,\ \phi_K \geq 0 \,. \tag{3.5}$$

The depreciation rate δ_t depends on the rate of capital utilization $u_t(i)$:

$$\delta_t(u_t) = \delta_0 + \delta_1\left(u_t - 1\right) + \frac{\delta_2}{2}\left(u_t - 1\right)^2 \,, \tag{3.6}$$

where $\delta_i \geq 0$ are parameters.

Intermediate good firms are owned by domestic households and maximize the expected sum of discounted cash flows $D_t(i)$,

$$\mathbb{E}_t \sum_{s=0}^{\infty} M_{t,t+s} \frac{D_{t+s}(i)}{P_{t+s}} \,, \tag{3.7}$$

subject to (3.3), (3.4), and (3.5), by choosing $N_t(i), u_t(i), I_t(i)$, and $P_t(i)$. Here,

$M_{t,t+s}$ denotes the real stochastic discount factor derived in Appendix C.2 and $\mathbb{E}_t$ is the conditional expectation operator. Real cash flows are given by

$$\frac{D_t(i)}{P_t} = \frac{P_t(i)}{P_t} Y_t(i) - \frac{W_t}{P_t} N_t(i) - I_t(i) - \frac{\phi_p}{2} \left(\frac{P_t(i)}{P_{t-1}(i)} - \Pi_H \right)^2 Y_t(i) \ . \tag{3.8}$$

Here, W_t is the wage, Π_H is steady-state PPI inflation, and the last term is the price adjustment cost as in Rotemberg (1982), measured in terms of domestic goods.

As in Basu and Bundick (2017), firms finance their operations by issuing shares S_t at price P_t^E and real risk-free discount bonds B_t^{rf}, paying the real interest rate R_t^R. We normalize the number of shares to one. Since the Modigliani-Miller theorem holds in our model, the financing structure does not matter for the value of the firm and the real economy. We can, therefore, set the face value of risk-free bonds that fund the capital stock to νK_t, where the parameter $0 < \nu < 1$ determines the leverage, which in turn determines the volatility of equity returns. Shareholders receive the residual cash flows as dividends:

$$\frac{D_t^E(i)}{P_t} = \frac{D_t(i)}{P_t} - \nu \left(K_{t-1}(i) - \frac{1}{{R^R}_t} K_t(i) \right) \ . \tag{3.9}$$

Foreign firms operate in an isomorphic environment, except for the fact that the export composite Y_t^H has infinitesimal weight in the composition of foreign final goods.

3.2 Households

For the representative household, following Epstein and Zin (1989) and Weil (1989), we assume preferences that allow risk aversion to be independent of the elasticity of intertemporal substitution.[10] Specifically, we write the household's expected lifetime utility recursively as

$$V_t = \max \left[(1-\beta_t) \left(\xi_{H,t} \xi_{C,t} C_t^{\varphi} (1-N_t)^{1-\varphi} \right)^{\frac{1-\sigma}{\theta_V}} + \beta_t \left(\mathbb{E}_t V_{t+1}^{1-\sigma} \right)^{\frac{1}{\theta_V}} \right]^{\frac{\theta_V}{1-\sigma}} . \tag{3.10}$$

Here C_t is consumption and N_t is hours worked, which are supplied to domestic intermediate goods firms. The parameter $\sigma \geq 0$ measures risk aversion, while ψ is

[10] We follow the specification in Basu and Bundick (2018), which does not lead to an asymptote of the model responses when the intertemporal elasticity of substitution approaches unity.

the intertemporal elasticity of substitution with $\theta_V \equiv \frac{1-\sigma}{1-\psi^{-1}}$. $0 \leq \varphi \leq 1$ denotes the share of the consumption good in the consumption-leisure bundle.[11] $\xi_{H,t}$ and $\xi_{C,t}$ denote home-specific and common shocks to the discount factor, "demand shocks" for short. They follow AR(1)-processes with stochastic volatility:

$$\xi_t^i = (1-\rho_{pref}) + \rho_{pref}\xi_{t-1}^i + \sigma_t^i \varepsilon_t^i \tag{3.11}$$

$$\sigma_t^i = (1-\rho_{\sigma^{pref}})\bar{\sigma}^{pref} + \rho_{\sigma^{pref}}\sigma_{t-1}^i + \sigma^{\sigma^{pref}}\varepsilon_t^{\sigma^i}\ , \tag{3.12}$$

where $i \in \{H, C\}$. The $\varepsilon_t^j, j \in \{H, C, \sigma^H, \sigma^C\}$ are standard normally distributed i.i.d. shock processes.[12] We henceforth refer to $\varepsilon_t^{\sigma_C}, \varepsilon_t^{\sigma_H}$ as common and country-specific uncertainty shocks, respectively.

Capital is mobile across borders, and households have access to a domestic (currency) bond B_t paying the nominal interest rate R_t and a foreign (currency) bond B_t^* that pays R_t^*. We assume an endogenous discount factor that decreases in the consumption-to-output ratio to ensure stationarity of the net foreign asset position. The foreign household has identical preferences, except for the absence of country-specific demand shocks, see Appendix C.1.

The household's period budget constraint reads, in nominal terms, as follows:

$$\begin{aligned} &B_t + \mathcal{E}_t B_t^* + P_t^E S_t + P_t\left(\frac{1}{R_t^R}B_t^{rf} + C_t\right) \\ &\leq W_t N_t + \left(P_t^E + D_t^E\right) S_{t-1} + P_t B_{t-1}^{rf} + R_{t-1}B_{t-1} + \mathcal{E}_t R_{t-1}^* B_{t-1}^*\ . \end{aligned} \tag{3.13}$$

The household maximizes (3.10), subject to the budget constraint (3.13). The first-order conditions for bonds can be combined into an uncovered interest parity condition that links domestic and foreign interest rates and monetary policy; see Appendix C.1.

The household in Foreign faces a nominal per-period budget constraint analogous to (3.13), but since Foreign acts as a closed economy from its point of view, it does not trade bonds of the Home country B_t.

[11]As a numerical matter, we introduce a normalizing constant to scale the discounted lifetime utility in the deterministic steady state to 1. While inconsequential for the results, it improves the numerical behavior of the model solution (e.g. Rudebusch and Swanson, 2012).

[12]We use a level specification in both the level and volatility equations rather than a log-log specification to avoid the problem of non-existent moments implied by the latter (Andreasen, 2010).

3.3 Monetary Policy

For the baseline, we assume that Home operates in a monetary union with Foreign. Since it has zero weight ($n \to 0$), its economic conditions are not included in the "union-wide" policy rule. Since monetary policy sets the nominal interest rate R_t^*, it follows a conventional interest rate feedback rule, responding only to "Foreign" inflation and output growth:

$$\frac{R_t^*}{R^*} = \left(\frac{R_{t-1}^*}{R^*}\right)^{\rho_R} \left(\left(\frac{\Pi_t^*}{\Pi^*}\right)^{\phi_{R\pi}} \left(\frac{Y_t^*}{Y_{t-1}^*}\right)^{\phi_{Ry}}\right)^{1-\rho_R} . \tag{3.14}$$

Here, R^* is the steady-state nominal interest rate, ρ_R is a smoothing parameter introduced to capture the empirical evidence of gradual movements in interest rates, Π^* is the inflation target set by the central bank, and the parameters $\phi_{R\pi}$ and ϕ_{Ry} capture the responsiveness of the nominal interest rate to deviations of inflation from its steady-state value and output growth, respectively. If a monetary union is in place, the nominal exchange rate is fixed at unity, and interest rates are perfectly aligned:

$$\mathcal{E}_t = 1 \text{ and } R_t = R_t^* \; \forall \, t. \tag{3.15}$$

As a counterfactual, we consider the case of monetary autonomy, assuming an interest rate for Home that mirrors that of Foreign:

$$\frac{R_t}{R} = \left(\frac{R_{t-1}}{R}\right)^{\rho_R} \left(\left(\frac{\Pi_{H,t}}{\Pi_H}\right)^{\phi_{R\pi}} \left(\frac{Y_t}{Y_{t-1}}\right)^{\phi_{Ry}}\right)^{1-\rho_R} , \tag{3.16}$$

with $\Pi^* = \Pi_H = P_t^H / P_{t-1}^H$, so that monetary policy has the same inflation target in Home and Foreign. Then, $\mathcal{E}_t$ adjusts to clear the foreign exchange market.

3.4 Equilibrium

Under Rotemberg price adjustment costs, there is a symmetric equilibrium in which the representative intermediate firm in each country charges the same price, uses the same amount of inputs, and the labor market clears $N_t = 1/n \int_0^n N_t(i) di$. Home and Foreign bonds each are in zero net supply in equilibrium. The resource constraint for Home implies that domestic output is used for consumption, investment, to pay for price adjustment costs, and for exports:

$$Y_t = (1-v)\left(\frac{\mathcal{Q}_t}{\mathcal{S}_t}\right)^{-\eta}(C_t + I_t) + \frac{\phi_p}{2}\left(\Pi_{H,t} - \Pi_H\right)^2 Y_t + v\mathcal{S}_t^{\eta} Y_t^* \,, \tag{3.17}$$

where $\mathcal{S}_t = P_{F,t}/P_{H,t}$ denotes the terms of trade.

Since Foreign behaves like a closed economy from its own point of view, the resource constraint there implies that all output is used for consumption, investment, and price adjustment:

$$Y_t^* = C_t^* + I_t^* + \frac{\phi_p}{2}\left(\Pi_t^* - \Pi^*\right)^2 Y_t^* \,. \tag{3.18}$$

4 Model-based analysis

We now bring the model to the data to provide a structural account of the time series evidence established in Section 2. We focus on the median EA economy and estimate model parameters by matching the model predictions to the time series evidence for the effects of a common uncertainty shock shown in Figure 4. We target the effects of a common shock rather than a country-specific shock because the latter has little or no effect on key variables. However, to provide an external validation of the model, we check that the model's predictions are consistent with the evidence along this dimension as well. Finally, we assess the transmission of country-specific uncertainty shocks in counterfactual scenarios, assuming that Home enjoys monetary autonomy and lets its exchange rate float freely.

4.1 Estimation

Prior to estimation, we fix a first set of parameters that are poorly identified or pinned down by long-run observations. In this respect, we mostly assume values in line with Basu and Bundick (2017, 2018), but make some adjustments where necessary to account for the open-economy dimension and the specificities of the EA. The second set of parameters is estimated by matching impulse response functions. Table 2 shows the parameters that are fixed before the estimation. The capital share α is set to 1/3, the discount factor β to 0.99, and the quarterly steady-state depreciation rate δ_0 to 0.025; δ_1 is set such that steady-state capital utilization

Table 2: Parameters fixed prior to estimation

Parameter	Description	Value	Target / Source
α	capital share parameter	0.3333	Basu and Bundick (2017)
β	discount factor	0.9900	4% interest rate per year
δ_0	depreciation rate steady state	0.0250	Basu and Bundick (2017)
δ_1	linear utilization cost	0.0351	steady-state utiliz. of 1
σ	risk aversion	100.00	Basu and Bundick (2018)
ψ	intertemp. elast. of subst.	0.5000	Basu and Bundick (2018)
φ	leisure share	0.2658	Frisch elasticity of 2
ϵ	intermed. goods subst. elast.	11.000	steady-state markup 10%
η	trade price elasticity	0.9000	Heathcote and Perri (2002)
ϕ_p	price adjustment costs	116.50	equivalent to Calvo (1983) parameter of 0.75
Π_H/Π^*	steady state inflation	1.0000	no trend inflation
υ	import share	0.4650	Gunnella et al. (2021)
ϕ_B	slope endog. discount factor	0.0010	small positive number
ν	leverage	0.9000	Basu and Bundick (2017)
Φ	fixed costs	0.1111	steady-state profits of 0

is 1. We set the risk aversion parameter $\sigma = 100$ and the intertemporal elasticity of substitution $\psi = 0.5$, following Basu and Bundick (2018). The leisure share in the Cobb-Douglas utility bundle φ is set to imply a Frisch elasticity of 2.[13] For the elasticity of substitution ϵ we assume a value of 11, corresponding to a steady-state markup of 10%. For the trade price elasticity η, we use the point estimate of 0.9 reported by Heathcote and Perri (2002). We set the Rotemberg price adjustment cost parameter ϕ_p to 116.5049, which corresponds to a slope of the linearized New Keynesian Phillips Curve consistent with a mean price duration of one year in a Calvo model. We assume the absence of trend inflation: $\Pi_H = \Pi^* = 1$. For the openness parameter υ, we chose a value of 0.465, in line with the estimate of average EA openness by Gunnella et al. (2021). The slope of the endogenous discount factor ϕ_B is set to a small positive number, sufficient to ensure stationarity (Schmitt-Grohé and Uribe, 2003). Leverage is set to 90% of assets, following Basu and Bundick (2017). Finally, $\Phi = 0.1111$ ensures zero profits in steady state.

[13] See Appendix A.2.1 of Born and Pfeifer (2021) for details.

The remaining nine parameters are estimated by matching the model impulse responses to a common uncertainty shock to their empirical counterparts from the VAR model.[14] We solve the model using third-order perturbation techniques in Dynare 6.1 and compute generalized impulse responses at the stochastic steady state while pruning the decision rules (Adjemian et al., 2024).[15] The vector of estimated parameters θ contains the parameters governing the exogenous processes (3.11) and (3.12), the capital adjustment costs, the quadratic capital utilization costs, and the three coefficients of the interest rate feedback rule. The point estimate $\hat{\theta}$ solves the following optimization problem:

$$\hat{\theta} = \underset{\theta}{\operatorname{argmin}} \left(\Psi^{Model}(\theta) - \Psi^{VAR} \right)' W \left(\Psi^{Model}(\theta) - \Psi^{VAR} \right) \quad . \tag{4.1}$$

Here, Ψ^{VAR} is a column vector stacking the VAR impulse responses to a one-standard-deviation common uncertainty shock in the EA sample up to period 20 after the shock, as shown in Figure 4 above. We include the responses of all variables in Ψ^{VAR}, except for the country-specific volatility component, as there is no counterpart in the model.[16] Ψ^{Model} includes the corresponding structural model impulse responses. W is a diagonal weighting matrix with the squared inverse of the width of the 68% HPDIs of the VAR responses on the diagonal. We put additional emphasis on matching the impact response by multiplying the weights of the first four quarters by 4^2. To avoid numerical problems, we truncate the weights at 150 for the variables measured in percent and at 1000 for the variables measured in percentage points. Following Ruge-Murcia (2010) and Born and Pfeifer (2014a), we formally incorporate our prior knowledge about plausible parameter ranges into our estimation. We do this by adding the quadratic distance of the parameters from their prior mean, standardized by the prior variance, to

[14]We match impulse responses directly, rather than employing an indirect inference approach. This allows us to remain agnostic about additional structural shocks required to avoid stochastic singularity in the VAR estimation.

[15]When computing impulse responses, we consider only a single preference shifter at a time. I.e., we set $\xi_t^H = 1$ when examining the transmission of a common uncertainty shock and $\xi_t^C = 1$ for a country-specific shock. This does not change the impulse responses of the first-moment variables but affects the VXO in the stochastic steady state. Our strategy ensures that the relative VXO response to a common shock is the same in both countries (which is desirable since Foreign should be seen as an aggregation of many small countries) and that the relative VXO response in Home is the same for both shocks, consistent with our normalization of VAR impulse responses in Figure 4.

[16]For the derivation of the model-implied stock market volatility index ("VXO"), see App. C.3.

Table 3: Estimated parameters

Parameter	Description	Prior Mean	Prior Std.	Point Estimate	Standard Error
ρ_{pref}	pref. shock autocorr.	0.90	0.20	0.8823	0.0154
$\rho_{\sigma^{pref}}$	pref. shock volatility autocorr.	0.01	∞	0.4672	0.0383
$\bar{\sigma}_{pref}$	pref. shock volatility	0.90	0.20	0.0387	0.0038
$\sigma_{\sigma^{pref}}$	pref. volatility shock volatility	0.30	∞	0.0153	0.0019
δ_2	quadratic utiliz. costs	0.01	0.20	0.0928	0.0487
ϕ_K	capital adjustment costs	4.00	1.50	3.6056	0.2174
ρ_r	interest rate smoothing	0.75	0.15	0.7455	0.0141
$\phi_{R\pi}$	inflation feedback	1.50	0.25	1.4374	0.0549
ϕ_{Ry}	output feedback	0.50	0.50	0.7323	0.0830

the objective function to be minimized. Table 3 shows the prior moments, which are mostly standard in the literature (e.g. Smets and Wouters, 2007). For the shock volatilities, we have no prior knowledge and use a uniform prior.

To calculate standard errors, we use a bootstrapping procedure that involves repeating the matching process described above for i.i.d. draws from the posterior distribution of the "median economy" impulse responses. The standard deviation of each parameter across these iterations is then computed.

The last two columns of Table 3 depict the point estimates and standard errors for the estimated parameters. Overall, the parameters are precisely estimated. The preference level shock is quite persistent, with an autocorrelation parameter of 0.88, and exhibits a quarterly standard deviation of 3.9 percent. The volatility process has a much lower persistence of 0.47. A one-standard-deviation shock increases the volatility of the level shock by 1.53 percentage points, or about 39 percent. Although we do not target the steady-state level of the VXO, it is roughly similar in the model, at 7.23 percentage points, to the data, where the median unconditional within-country mean is 17.6 percentage points. The fact that volatility is higher in the data is consistent with the notion that the data are driven by more shocks than the model, and that the magnitude of our reported output responses is not driven by an unrealistically high degree of volatility in the model. Capital utilization costs show a low curvature of $\delta_2 = 0.0928$, while investment adjustment costs are moderate with $\phi_K = 3.61$. The Taylor-rule coefficients for interest rate smoothing and inflation feedback, estimated at 0.75 and 1.44 respectively, are slightly higher than those estimated for the euro area by Enders et al. (2013), while the output feedback is estimated to be relatively strong at 0.73.

4.2 Model performance

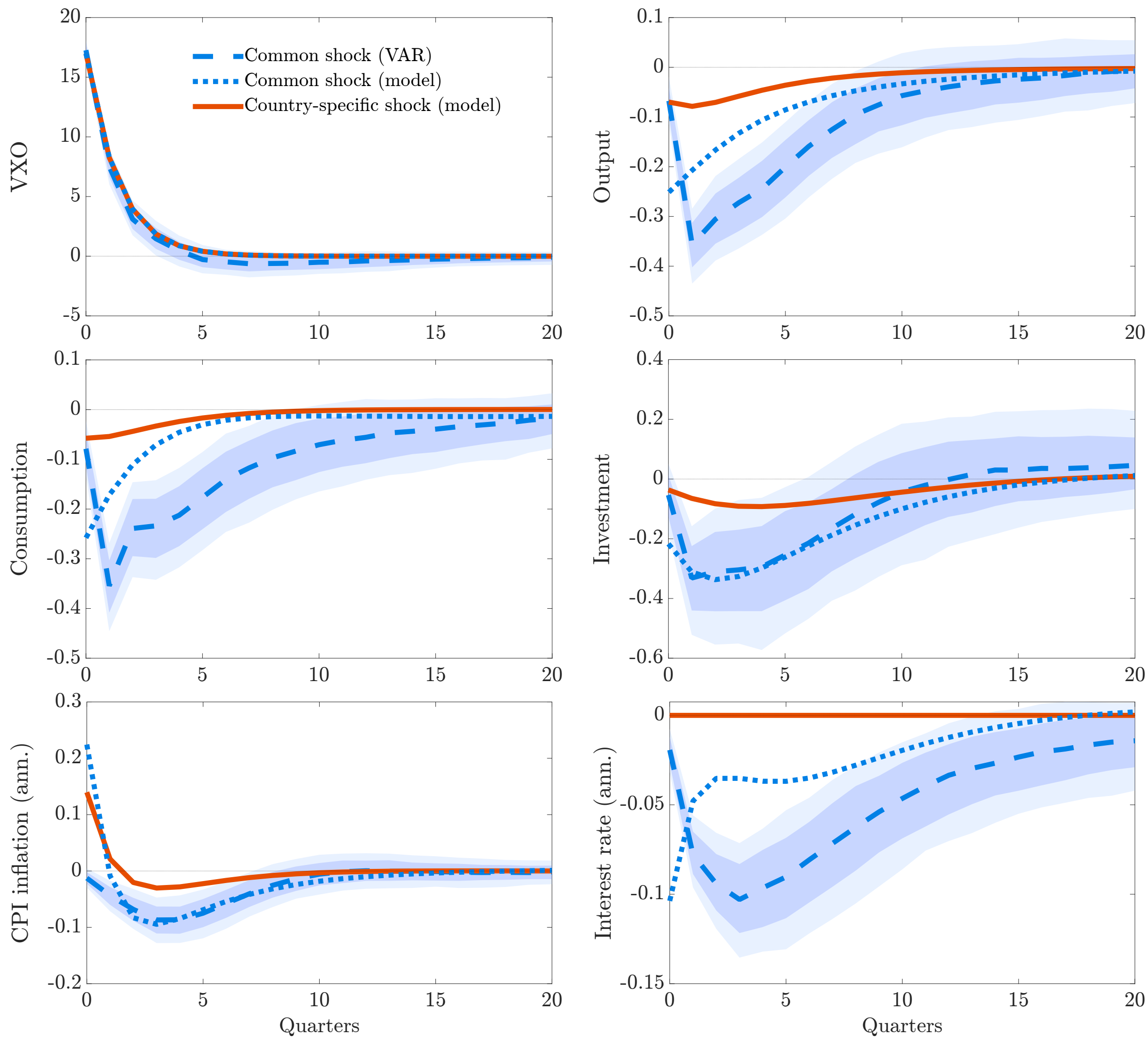
The estimated model is able to account for the time series evidence presented in Section 2 along a number of key dimensions. First, consider the model fit shown in Figure 5: The dotted (blue) lines show the responses of the calibrated model to a common uncertainty shock equal to one standard deviation. We contrast these responses with those for the VAR estimated on the EA countries, reproduced from Figure 4 above. They are shown by the dashed (blue) lines, and the shaded area indicates, as before, the pointwise 68% (dark) and 90% (light) HPDIs. Even as the model is over-identified, its predictions match the empirical responses quite well. As in the data, a common uncertainty shock has a substantial contractionary effect on the economy. Output, consumption, and investment all fall. The model predicts, somewhat counterfactually, inflation to increase on impact, but matches its medium to long-term adjustment well.

Figure 5 also shows the model's predictions for the effect of a country-specific uncertainty shock. They are represented by the solid (red) lines. In this respect, the model prediction is also quite consistent with the VAR evidence in Figure 4, both qualitatively and quantitatively.[17] The effect of a country-specific shock tends to be markedly weaker than that of a similarly sized common shock. This prediction is noteworthy because the empirical responses to a country-specific shock were not used to estimate the model.

Turning to the bottom-right panel of Figure 5, we observe that the response of the interest rate to a country-specific shock is flat. This is because Home is a small country in the monetary union and therefore has zero weight in the union-wide interest rate rule. In contrast, monetary policy lowers interest rates in response to a common uncertainty shock to dampen its recessionary impact. Still, as the responses in Figure 5 show, the monetary accommodation in response to the

[17] A detailed comparison between the model and empirical responses to a country-specific shock is shown in Figure C.1 in the appendix.

Figure 5: Adjustment in EA countries—estimated model v time series evidence



Notes: Adjustment to common (dotted blue line) and country-specific (red solid line) uncertainty shock (normalized to one standard deviation) in a monetary union according to model; empirical responses to common shock (dashed blue line) reproduced from Figure 4. Horizontal axis: quarters, vertical axis: deviations from pre-shock level in percent, except for inflation and interest rate (ppts). Shaded area: pointwise 68% (dark) and 90% (light) HPDIs.

common shock is insufficient: Its recessionary impact is actually larger than that of a country-specific shock.

4.3 Union v flexible exchange rates: the role of price level risk

We are finally able to shed light on our main finding: that the effects of country-specific uncertainty shocks are weaker in a country operating in a monetary union, even though countries outside the union enjoy monetary autonomy. To do so, we compare the adjustment dynamics in the union with a scenario in which Home operates outside the union, allows the exchange rate to adjust freely in response to shocks, and adjusts the interest rate according to (3.16). This rule differs from the union-wide monetary policy rule only in that, with flexible exchange rates, monetary policy responds to domestic developments rather than to union-wide developments (which are dominated by Foreign).

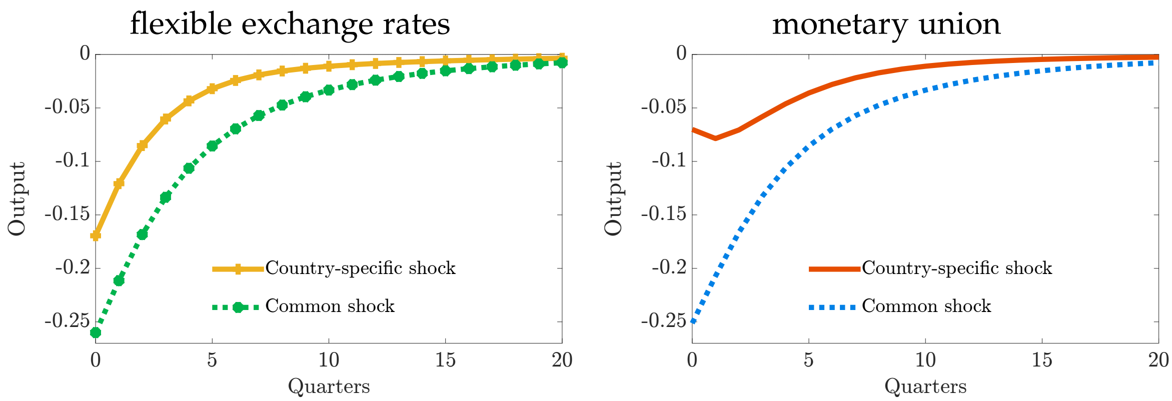
We focus on the output response, which is shown in Figure 6. The left panel shows the flexible exchange rate scenario, contrasting the effect of a common uncertainty shock (dotted green line with octagonal markers) and a country-specific shock (solid yellow line with plus-shaped markers). While both shocks are normalized to have the same effect on volatility (not shown), the effect of the common shock is about 50 percent larger. Intuitively, when both Home and Foreign experience the shock, there is a global contraction, which features adverse spillovers from Foreign (large) to Home (small).[18] Such spillovers are absent when the shock is specific to Home.

The adjustment under flexible exchange rates serves as a natural benchmark for interpreting the adjustment dynamics when Home operates in a monetary union. We consider this case in the right panel of Figure 6, reproducing the output responses already shown in Figure 5 above. The response to a common shock is the same as in the left panel. This result is consistent with the evidence for the median EA economy and the median floater, shown in Figure 2 and, as discussed above, is to be expected: if Home and Foreign are symmetric and exposed to the same shock, the exchange rate regime does not affect the outcome.

However, the exchange rate regime matters a great deal for how a country-specific shock plays out. When Home operates in a monetary union (right panel), the output response is only about half as large as under flexible exchange rates (left panel). In other words, the model predicts not only that the effects of country-specific shocks are weaker in a monetary union than those of a common shock; it

[18]Figure C.2 shows a Foreign-only shock, which also lowers economic activity in Home.

Figure 6: Output response to uncertainty shocks under ...



Notes: Model impulse responses to a one-standard-deviation country-specific (left panel: yellow solid line with plus-shaped markers, right panel: red solid line) and a comparably sized common uncertainty shock (left panel: green dotted line with octagonal markers, right panel: blue dotted line). Quarterly responses are in percentage deviations from the stochastic steady state. Left panel assumes flexible exchange rates and that monetary policy follows rule (3.16).

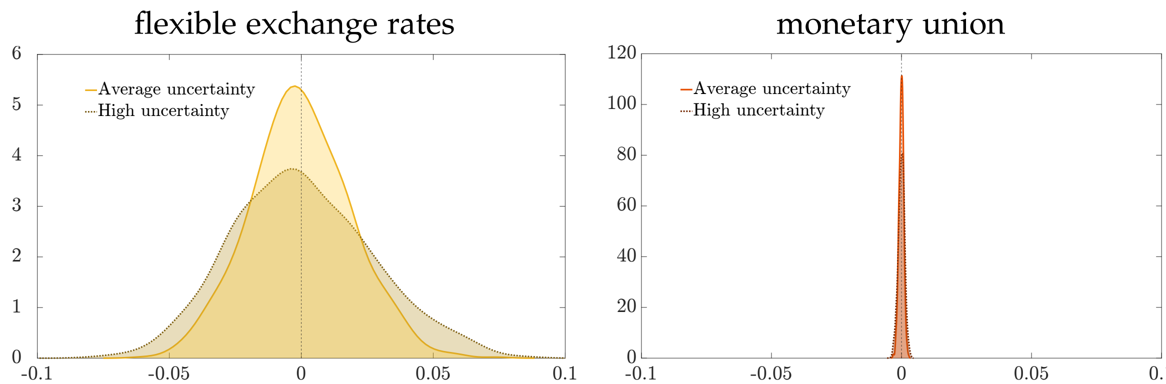
also predicts that they are much weaker than under flexible exchange rates, both predictions being consistent with the evidence established in Section 2 above.

This result is remarkable because, in a monetary union, union-wide monetary policy does not accommodate Home shocks.[19] At the same time, monetary policy in Home can no longer react because by joining the union, Home gives up its monetary autonomy in exchange for anchoring its price level to that of the union. This notion is formalized in earlier work by Giavazzi and Pagano (1988) and Alesina and Barro (2002) with a focus on how the anchor removes the inflation bias that raises average inflation independently of the business cycle.

However, it turns out that the nominal anchor also plays a crucial role for business cycle dynamics, in particular for the transmission of uncertainty shocks. To illustrate this, we compute the distribution of the price level 100 periods after a first-moment shock, considering both the flexible exchange rate scenario and the monetary union case. Figure 7 shows the results, contrasting the distribution

[19]According to the VAR evidence, monetary policy under a float raises interest rates in response to country-specific shocks. However, this in itself does not explain why the contraction is larger there under a float. As we show in Figure C.3, monetary policy cuts interest rates in response to a country-specific uncertainty shock in our estimated model when Home operates under flexible exchange rates. And yet, economic activity contracts more than in the union case.

Figure 7: Simulated distributions of the long-run price level under . . .



Kernel densities of long-run price level (CPI) after random one-time level demand shock drawn from distribution with average uncertainty (solid line) and widened distribution after one-standard deviation uncertainty shock (dotted line) under floating exchange rate (left panel) and in monetary union (right panel). We draw 1000 shock realizations from each distribution and then take the cumulated inflation response 100 periods after the shock.

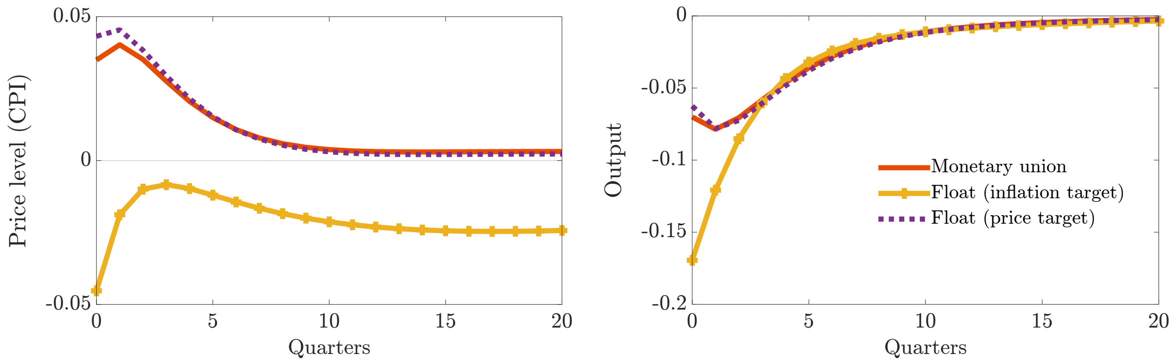
with average uncertainty, i.e., uncertainty at the unconditional mean (light-shaded area), with the case when a one-standard deviation uncertainty shock widens the shock distribution (dark-shaded area). The left panel illustrates how increased uncertainty translates into greater price level risk, as the long-run distribution shifts with the distribution of shocks. In contrast, the monetary union eliminates price level risk, as shown in the right panel. Since the domestic price level is anchored to the union-wide price level through purchasing power parity, it no longer exhibits a unit root. Consequently, the long-run distribution of the price level remains unaffected by the shock.[20]

In light of this result, we consider an alternative scenario for monetary policy under flexible exchange rates: We assume that Home runs an independent monetary policy, but targets the domestic price level instead of inflation in the interest rate rule (3.16). Specifically, it weakly adjusts interest rates whenever domestic prices deviate from their steady-state level:

$$R_t = R \left(\frac{P_{H,t}}{P_H} \right)^{0.0005} . \tag{4.2}$$

[20]Note that we show the distribution in period 100 after a possible first-moment shock. Asymptotically, the distribution collapses to zero.

Figure 8: Effects of country-specific uncertainty shocks w/ alternative policies



Model IRFs to a one-standard-deviation country-specific preference uncertainty shock in a monetary union (red solid line), a float with inflation targeting (yellow solid line with plus-shaped markers), and a float with price targeting (purple dotted line). Quarterly responses are in percentage deviations from the stochastic steady state.

This rule implies a very mild monetary response to the shock; in fact, there is hardly any visible interest rate response, as we show in Figure C.3 along with the response of additional variables. However, the price level rule (4.2) anchors the price level in the long run, similar to what union membership does. We can see this in the left panel of Figure 8. The dotted (purple) line represents the case of flexible exchange rates when monetary policy follows the price level rule (4.2). It removes the unit root in the price level, and the adjustment mechanism is almost identical to what we observe for the monetary union, shown by the solid red line. This is true not only for the price level but also for output, shown in the right panel of the same figure. The inflation targeting case is shown by the solid (yellow) line with plus-shaped markers. Here, the price level features a unit root (left) and output falls much more, as discussed above.

In sum, membership in a monetary union anchors the price level in a way that is comparable to what happens when an independent monetary policy targets the price level—in either case, the long-run risk to the price level is eliminated. Note, however, that as far as the monetary union is concerned, this is only true for country-specific shocks: in this case, the price level temporarily deviates from that of the rest of the union, but it adjusts over time, ensuring that purchasing parity is satisfied in the long run, consistent with the evidence (Bergin et al., 2017).[21]

[21]Computing the long-run response of the price level implied by the inflation response to a

4.4 Households v firms

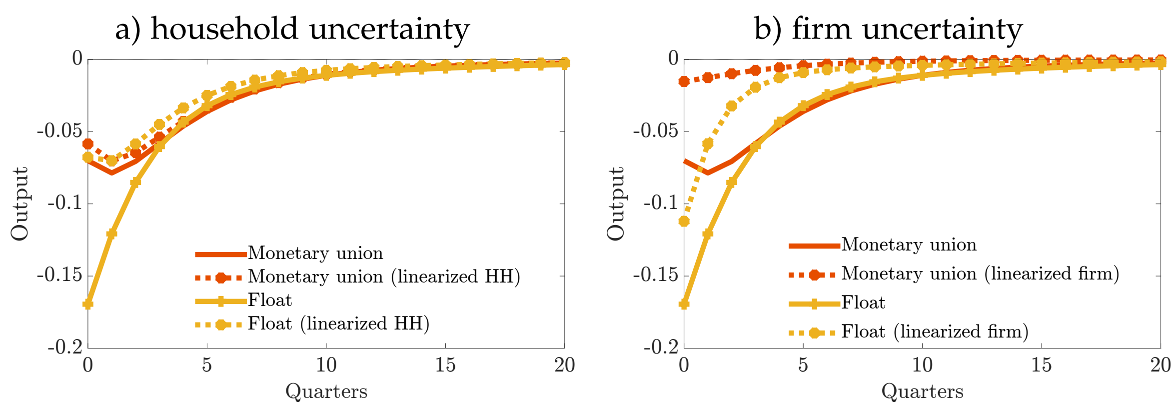
Price level risk is potentially important for both firm and household decisions because of precautionary saving and precautionary pricing (Born and Pfeifer, 2014a; Fernández-Villaverde et al., 2015; Basu and Bundick, 2017; Bianchi et al., 2023). To quantify the extent to which price level risk affects households and firms, we conduct further model simulations. They are centered on the insight that the effect of uncertainty shocks operates through time-varying, endogenous risk wedges that arise in forward-looking expectations equations. Intuitively, suppose that the probability distribution of shocks widens due to an uncertainty shock. In this case, the expected values in the optimality conditions of households and firms change due to Jensen's inequality—in contrast to what happens in a linear world. To identify the quantitatively important margins in this regard, we perform a decomposition à la Bianchi et al. (2023): We selectively shut off these endogenous risk wedges by solving variants of the model in which certain forward-looking equations are restricted to their log-linear approximation.

For our analysis, we split the overall effect of uncertainty shocks into two main categories: the household's consumption-saving decision and the firm's pricing decision. On the household side, we shut off the wedges associated with precautionary saving and price level risk embedded in the Euler equation, the investment adjustment wedge embedded in the investment first-order condition, and the investment risk wedge embedded in the first-order condition for capital. The results are shown in the left panel of Figure 9. It contrasts the results for the baseline scenario (solid lines), reproduced from Figure 8 above, with the results obtained when uncertainty is assumed to play no role in households' intertemporal decisions (dotted lines with octagonal markers). Recall that in the baseline, a country-specific uncertainty shock has much stronger output effects in the case of a floating exchange rate. We now see that consumption-saving wedges are responsible for this difference: the difference between float and union membership disappears almost entirely.

Next, we isolate the endogenous risk wedge associated with the firm's pricing decision and show the results in the right panel of Figure 9. In this case, we remove the nominal pricing bias by log-linearizing the firm's recursive pricing

country-specific shock in the VAR (Figure 4), we find that it is not significantly different from zero.

Figure 9: Output effect of country-specific uncertainty w/o ...



Model responses to a one-standard-deviation country-specific uncertainty shock in a monetary union (red solid line) and under a float with inflation targeting (yellow solid line with plus-shaped markers). Left panel: no risk associated with the household's consumption-saving decision (dotted lines with octagonal markers). Right panel: no risk associated with the firms' pricing decision (dotted lines with octagonal markers). Quarterly responses are in percentage deviations from the stochastic steady state.

equation. Also in this case, the output effects of country-specific uncertainty shocks are substantially reduced relative to the baseline. However, the reduction is comparable for both the float and union membership cases. Thus, it price level risk is quantitatively less important on the firm side. To sum up, our simulation shows that the anchoring of price-level expectations in a monetary union—and the resulting reduction in price level risk—is particularly important for households' saving decisions.

Finally, we note that the reduction in price level risk in the monetary union does not necessarily dampen the impact of all shocks. For instance, we find that a country-specific preference (level) shock has stronger adverse output effects in the monetary union than under a float (when monetary policy lowers interest rates and thus stimulates investment, see Figure C.4). It is, therefore, possible that the containment of price level risk is particularly important for the transmission of uncertainty shocks. We leave it to future work, both empirical and in quantitative business cycle models, to explore this conjecture.

5 Conclusion

A monetary union provides a nominal anchor for the price levels of its members. This not only eliminates potential inflation biases but also reduces price level risk, as shown in this paper. As such, it also dampens the effects of country-specific uncertainty shocks within the monetary union, making them weaker, rather than stronger than those experienced by countries with flexible exchange rates.

This is particularly relevant in a context of heightened uncertainty. While economic uncertainty is often heightened by global events, such as Russia's invasion of Ukraine, they tend to load differently on different countries (Federle et al., 2024), giving rise to country-specific uncertainty shocks. In such an environment, the lack of monetary independence may prove less costly than is commonly perceived.

References

Adjemian, Stéphane et al. (2024). "Dynare: reference manual, version 6". Dynare Working Papers 80. CEPREMAP.

Albagli, Elias, Luis Ceballos, Sebastian Claro, and Damian Romero (2024). "UIP deviations: insights from event studies". *Journal of International Economics 148*, 103877.

Alesina, Alberto and Robert J. Barro (2002). "Currency unions". *Quarterly Journal of Economics 117* (2), 409–436.

Andreasen, Martin M. (2010). "Stochastic volatility and DSGE models". *Economics Letters 108* (1), 7–9.

Andreasen, Martin M., Giovanni Caggiano, Efrem Castelnuovo, and Giovanni Pellegrino (2024). "Does risk matter more in recessions than in expansions? Implications for monetary policy". *Journal of Monetary Economics 143*, 103533.

Baker, Scott R., Nicholas Bloom, and Steven J. Davis (2016). "Measuring economic policy uncertainty". *Quarterly Journal of Economics 131* (4), 1593–1636.

Başkaya, Yusuf Soner, Timur Hülagü, and Hande Küşük (2013). "Oil price uncertainty in a small open economy". *IMF Economic Review 61* (1), 168–198.

Basu, Susanto and Brent Bundick (2017). "Uncertainty shocks in a model of effective demand". *Econometrica 85* (3), 937–958.

Basu, Susanto and Brent Bundick (2018). "Uncertainty shocks in a model of effective demand: Reply". *Econometrica 86* (4), 1527–1531.

Bayer, Christian, Alexander Kriwoluzky, Gernot J. Müller, and Fabian Seyrich (2024). "A HANK^2 model of monetary unions". *Journal of Monetary Economics*, 103579.

Bayoumi, Tamim and Barry Eichengreen (1994). "Macroeconomic adjustment under Bretton Woods and the post-Bretton-Woods float: an impulse-response analysis". *Economic Journal 104* (425), 813–827.

Bergin, Paul R., Reuven Glick, and Jyh-Lin Wu (2017). ""Conditional PPP" and real exchange rate convergence in the euro area". *Journal of International Money and Finance 73*, 78–92.

Bianchi, Francesco, Howard Kung, and Mikhail Tirskikh (2023). "The origins and effects of macroeconomic uncertainty". *Quantitative Economics 14* (3), 855–896.

Bloom, Nicholas (2009). "The impact of uncertainty shocks". *Econometrica 77* (3), 623–685.

——— (2014). "Fluctuations in uncertainty". *Journal of Economic Perspectives 28* (2), 153–76.

Born, Benjamin, Francesco D'Ascanio, Gernot J. Müller, and Johannes Pfeifer (2024). "Mr. Keynes meets the Classics: government spending and the real exchange rate". *Journal of Political Economy 132* (2), 1642–1683.

Born, Benjamin, Falko Juessen, and Gernot J. Müller (2013). "Exchange rate regimes and fiscal multipliers". *Journal of Economic Dynamics and Control 37* (2), 446–465.

Born, Benjamin and Johannes Pfeifer (2014a). "Policy risk and the business cycle". *Journal of Monetary Economics 68*, 68–85.

——— (2014b). "Risk matters: the real effects of volatility shocks: Comment". *American Economic Review 104* (12), 4231–4239.

——— (2021). "Uncertainty-driven business cycles: assessing the markup channel". *Quantitative Economics 12* (2), 587–623.

Broda, Christian (2004). "Terms of trade and exchange rate regimes in developing countries". *Journal of International Economics 63* (1), 31–58.

Caggiano, Giovanni and Efrem Castelnuovo (2023). "Global financial uncertainty". *Journal of Applied Econometrics 38* (3), 432–449.

Caggiano, Giovanni, Efrem Castelnuovo, and Giovanni Pellegrino (2017). "Estimating the real effects of uncertainty shocks at the Zero Lower Bound". *European Economic Review 100*, 257–272.

Calvo, Guillermo A. (1983). "Staggered prices in a utility-maximizing framework". *Journal of Monetary Economics 12* (3), 383–398.

Canova, Fabio and Matteo Ciccarelli (2013). "Panel vector autoregressive models: a survey". *Advances in Econometrics 32* (15), 205–246.

Carballo, Jeronimo, Kyle Handley, and Nuno Limão (2022). "Economic and policy uncertainty: aggregate export dynamics and the value of agreements". *Journal of International Economics 139*, 103661.

Castelnuovo, Efrem (2023). "Uncertainty before and during COVID-19: a survey". *Journal of Economic Surveys 37* (3), 821–864.

Comunale, Mariarosaria and Anh Dinh Minh Nguyen (2023). "A comprehensive macroeconomic uncertainty measure for the euro area and its implications to COVID-19". IMF Working Paper 23/229.

Corsetti, Giancarlo, Keith Kuester, and Gernot J. Müller (2013). "Floats, pegs, and the transmission of fiscal policy". *Fiscal policy and macroeconomic performance*. Ed. by Jordi Galí and Luis Cespedes. Vol. 17. Central Banking, Analysis, and Economic Policies. Central Bank of Chile. Chap. 7, 253–287.

Corsetti, Giancarlo, Keith Kuester, Gernot J. Müller, and Sebastian Schmidt (2021). "The exchange rate insulation puzzle". CEPR Discussion Paper 15689.

Degasperi, Riccardo, Seokki Simon Hong, and Giovanni Ricco (2023). "The global transmission of U.S. monetary policy". Working Papers 2023-02. Center for Research in Economics and Statistics.

Doan, Thomas, Robert Litterman, and Christopher Sims (1984). "Forecasting and conditional projection using realistic prior distributions". *Econometric Reviews 3* (1), 1–100.

Enders, Zeno, Philip Jung, and Gernot J. Müller (2013). "Has the Euro changed the business cycle?" *European Economic Review 59*, 189–211.

Epstein, Larry G. and Stanley E. Zin (1989). "Substitution, risk aversion, and the temporal behavior of consumption and asset returns: a theoretical framework". *Econometrica 57* (4), 937–969.

Fasani, Stefano and Lorenza Rossi (2018). "Are uncertainty shocks aggregate demand shocks?" *Economics Letters 167*, 142–146.

Federle, Jonathan, André Meier, Gernot J. Müller, Willi Mutschler, and Moritz Schularick (2024). "The price of war". CEPR Discussion paper 18834.

Fernandes, Ana P. and L. Alan Winters (2021). "Exporters and shocks: the impact of the Brexit vote shock on bilateral exports to the UK". *Journal of International Economics 131*, 103489.

Fernández-Villaverde, Jesús, Pablo A. Guerrón-Quintana, Keith Kuester, and Juan F. Rubio-Ramírez (2015). "Fiscal volatility shocks and economic activity". *American Economic Review 105* (11), 3352–3384.

Fernández-Villaverde, Jesús, Pablo A. Guerrón-Quintana, Juan F. Rubio-Ramírez, and Martín Uribe (2011). "Risk matters: the real effects of volatility shocks". *American Economic Review 101* (6), 2530–61.

Fukui, Masao, Emi Nakamura, and Jón Steinsson (2023). "The macroeconomic consequences of exchange rate depreciations". NBER Working Paper 31279.

Galí, Jordi and Tommaso Monacelli (2005). "Monetary policy and exchange rate volatility in a small open economy". *Review of Economic Studies 72* (3), 707–734.

——— (2008). "Optimal monetary and fiscal policy in a currency union". *Journal of International Economics 76* (1), 116–132.

Georgiadis, Georgios, Gernot J. Müller, and Ben Schumann (2024). "Global risk and the dollar". *Journal of Monetary Economics*, 103549.

Giavazzi, Francesco and Marco Pagano (1988). "The advantage of tying one's hands: EMS discipline and central bank credibility". *European Economic Review 32* (5), 1055–1075.

Giovanni, Julian di and Jay C. Shambaugh (2008). "The impact of foreign interest rates on the economy: the role of the exchange rate regime". *Journal of International Economics 74* (2), 341–361.

Gopinath, Gita et al. (2020). "Dominant currency paradigm". *American Economic Review 110* (3), 677–719.

Groll, Dominik and Tommaso Monacelli (2020). "The inherent benefit of monetary unions". *Journal of Monetary Economics 111*, 63–79.

Gunnella, Vanessa, Laura Lebastard, Paloma López-García, Roberta Serafini, and Alessandro Zona Mattioli (2021). "The impact of the euro on trade: two decades into monetary union". *ECB Occasional Paper* (2021/283).

Handley, Kyle and Nuno Limão (2017). "Policy uncertainty, trade, and welfare: theory and evidence for China and the United States". *American Economic Review 107* (9), 2731–83.

Heathcote, Jonathan and Fabrizio Perri (2002). "Financial autarky and international business cycles". *Journal of Monetary Economics 49* (3), 601–627.

Ilzetzki, Ethan, Carmen M. Reinhart, and Kenneth S. Rogoff (2022). "Rethinking exchange rate regimes". *Handbook of International Economics*. Ed. by Gita Gopinath, Elhanan Helpman, and Kenneth S. Rogoff. Vol. 6. North Holland: Elsevier. Chap. 3, 91–145.

Johannsen, Benjamin K. (2014). "When are the effects of fiscal policy uncertainty large?" Finance & Economics Discussion Series 2014-40. Federal Reserve Board.

Johri, Alok, Shahed Khan, and César Sosa-Padilla (2022). "Interest rate uncertainty and sovereign default risk". *Journal of International Economics 139*, 103681.

Jurado, Kyle, Sydney C. Ludvigson, and Serena Ng (2015). "Measuring uncertainty". *American Economic Review 105* (3), 1177–1216.

Kadiyala, K. Rao and Sune Karlsson (1997). "Numerical methods for estimation and inference in Bayesian VAR-models". *Journal of Applied Econometrics 12* (2), 99–132.

Kilian, Lutz, Michael Plante, and Alexander W. Richter (2022). "Macroeconomic responses to uncertainty shocks: the perils of recursive orderings". FRB of Dallas Working Paper 2223.

Kollmann, Robert (2016). "International business cycles and risk sharing with uncertainty shocks and recursive preferences". *Journal of Economic Dynamics and Control 72*, 115–124.

Krugman, Paul (1993). "Lessons of Massachusetts for EMU". *Adjustment and growth in the European Monetary Union*. Ed. by Francisco Torres and Francesco Giavazzi. Cambridge University Press. Chap. 8, 241–261.

Lakdawala, Aeimit, Timothy Moreland, and Matthew Schaffer (2021). "The international spillover effects of US monetary policy uncertainty". *Journal of International Economics 133*, 103525.

Leduc, Sylvain and Zheng Liu (2016). "Uncertainty shocks are aggregate demand shocks". *Journal of Monetary Economics 82*, 20–35.

Litterman, Robert B. (1986). "Forecasting with Bayesian vector autoregressions: five years of experience". *Journal of Business & Economic Statistics 4* (1), 25–38.

Ludvigson, Sydney C., Sai Ma, and Serena Ng (2021). "Uncertainty and business cycles: exogenous impulse or endogenous response?" *American Economic Journal: Macroeconomics 13* (4), 369–410.

Meinen, Philipp and Oke Roehe (2017). "On measuring uncertainty and its impact on investment: cross-country evidence from the euro area". *European Economic Review 92* (C), 161–179.

Mumtaz, Haroon and Konstantinos Theodoridis (2017). "Common and country specific economic uncertainty". *Journal of International Economics 105*, 205–216.

Mumtaz, Haroon and Francesco Zanetti (2013). "The impact of the volatility of monetary policy shocks". *Journal of Money, Credit and Banking 45* (4), 535–558.

Mundell, Robert A. (1961). "A theory of optimum currency areas". *American Economic Review 51* (4), 657–665.

Nakata, Taisuke (2017). "Uncertainty at the zero lower bound". *American Economic Journal: Macroeconomics 9* (3), 186–221.

Pellegrino, Giovanni, Efrem Castelnuovo, and Giovanni Caggiano (2023). "Uncertainty and monetary policy during the Great Recession". *International Economic Review 64* (2), 577–606.

Pesaran, M. Hashem and Ron Smith (1995). "Estimating long-run relationships from dynamic heterogeneous panels". *Journal of Econometrics 68*, 79–113.

Piffer, Michele and Maximilian Podstawski (2018). "Identifying uncertainty shocks using the price of gold". *Economic Journal 128* (616), 3266–3284.

Plante, Michael, Alexander W. Richter, and Nathaniel A. Throckmorton (2018). "The zero lower bound and endogenous uncertainty". *Economic Journal 128* (611), 1730–1757.

Raftery, Adrian E. and Steven Lewis (1992). "How many iterations in the Gibbs sampler?" *Bayesian Statistics*. Ed. by José M. Bernardo, James O. Berger, A. Philip Dawid, and Adrian F. M. Smith. Vol. 4. Oxford: Clarendon Press, 763–773.

Redl, Chris (2020). "Uncertainty matters: evidence from close elections". *Journal of International Economics 124*, 103296.

Rotemberg, Julio J. (1982). "Sticky prices in the United States". *Journal of Political Economy 90* (6), 1187–1211.

Rudebusch, Glenn D. and Eric T. Swanson (2012). "The bond premium in a DSGE model with long-run real and nominal risks". *American Economic Journal: Macroeconomics 4* (1), 105–143.

Ruge-Murcia, Francisco J. (2010). “Estimating nonlinear DSGE models by the simulated method of moments”. CIREQ Working Paper Series 19-2010.

Schmitt-Grohé, Stephanie and Martín Uribe (2003). “Closing small open economy models”. *Journal of International Economics 61*, 163–185.

Smets, Frank and Rafael Wouters (2007). “Shocks and frictions in US business cycles: a Bayesian DSGE approach”. *American Economic Review 97* (3), 586–606.

Weil, Philippe (1989). “The equity premium puzzle and the risk-free rate puzzle”. *Journal of Monetary Economics 24* (3), 401–421.

Wu, Jing Cynthia and Fan Dora Xia (2016). “Measuring the macroeconomic impact of monetary policy at the zero lower bound”. *Journal of Money, Credit and Banking 48* (2-3), 253–291.

A Data appendix

In our euro area VARs, we use the following country-level data series:

1. Total and country-specific volatility: the average realized return volatility of the Datastream Country Market Total Return Index during the quarter, computed as the average annualized standard deviation of daily returns of the performance index (obtained from Datastream: `TOTMK*(RI)`); for the country-specific realized volatility, we remove the first principal component of the constructed volatility indices of a sample of 48 countries (see below). We use the alternating least squares (ALS) algorithm of the Matlab `pca`-function in the R2023a version to deal with missing values. We first conduct the PCA at the monthly frequency and then aggregate the resulting series to quarterly data.

2. log GDP: real GDP, Million euro, chain-linked volumes, reference year 2010 (Eurostat table `namq_10_gdp`, series `B1GQ`), divided by population

3. log Consumption: real personal consumption expenditures, Million euro, chain-linked volumes, reference year 2010 (Eurostat table `namq_10_gdp`, Household and NPISH final consumption expenditure series `P31_S14_S15`), divided by population

4. log Investment: real private investment, Million euro, chain-linked volumes, reference year 2010 (Eurostat table `namq_10_gdp`, series `P51G`), divided by population

5. Inflation: Harmonized index of consumer prices (annual rate of change), all-items HICP (Eurostat table `prc_hicp_manr`). We aggregate the monthly data to quarterly frequency by using the geometric mean of all monthly observations within in a given quarter.

6. Policy rate: ECB interest rate for main refinancing operations / End of month (ECB Data Warehouse, `BBK01.SU0202`), complemented by the Wu and Xia (2016) shadow rate for the Euro Area

7. log Government spending: real final consumption expenditures of general

government, Million euro, chain-linked volumes, reference year 2010 (Eurostat table namq_10_gdp, series P3_S13), divided by population

8. log Stock market index: Quarterly average of Datastream Country Market Total Return Index (TOTMK*(RI)), same as for volatility measures). We first average the daily values within each month and subsequently aggregate to quarterly averages.

9. Jurado et al. (2015) uncertainty measure: 12-month-ahead macroeconomic uncertainty compiled by Comunale and Nguyen (2023). For the isolation of country-specific uncertainty, we remove the first principal component of that uncertainty index for all 17 EA countries in Table A.1.

To construct per capita values, we use Total population national concept, Seasonally and calendar adjusted data, Thousand persons (Eurostat table namq_10_gdp, series POP_NC).

All national account series employed seasonally and calendar-adjusted data. For countries where not all series (apart from volatility and the shadow rate) are available with the same seasonal and calendar adjustment (Italy, France, Greece, Portugal), we use the Matlab x13-function of the X-13 Toolbox for Seasonal Filtering, version 1.58, to remove seasonal fluctuations.

In the VARs, we use the countries listed in Table A.1. The second column reports the start of the sample. We exclude Cyprus from the VAR with the Jurado et al. (2015) uncertainty measure because its coefficient for the first principal component is negative, leading to its common component being inversely related to that of all other countries.

PCA sample

For the principal component analysis, we use a sample of 48 countries made up of OECD members, participating partners, and countries negotiating OECD membership. Two exceptions are the OECD members Costa Rica, for which the Datastream stock market index is unavailable, and Iceland, whose stock market crash following the financial crisis in 2008-2011 significantly distorts all PCA outcomes.

Table A.1: Country starting dates EA sample

Country	First Quarter Used
Austria	1999Q1
Belgium	1999Q1
Cyprus	2008Q1
Estonia	1999Q1
Finland	1999Q1
France	1999Q1
Germany	1999Q1
Greece	1999Q1
Ireland	1999Q1
Italy	1999Q1
Latvia	2009Q3
Lithuania	1999Q1
Luxembourg	1999Q1
Netherlands	1999Q1
Portugal	1999Q1
Slovenia	2007Q1
Spain	1999Q1

Notes: For countries that did not adopt the euro at its inception in 1999Q1, we start the sample when Ilzetzki et al. (2022) classify the country's exchange rate arrangement as a 2 ("Pre announced peg or currency board arrangement") or 1 ("No separate legal tender or currency union") in their fine classification.

Countries included in the PCA sample but in neither of the VARs are Argentina, Bulgaria, China, Croatia, Czechia, Denmark, Hungary, Indonesia, Japan, Norway, Peru, Poland, Romania, Saudi Arabia, Sweden, Switzerland, Turkey, and United Kingdom. These countries either had pegged exchange rates for most or all of the sample duration or are European floaters not insulated from EA-wide shocks (see Corsetti et al. (2021)). Another exception is Japan, which we dropped from the VAR due to its 20 years at the zero lower bound, which creates stochastic singularity issues in the VAR once we include the policy rate.

Global float sample

In our global sample of countries with floating exchange rates, we use the following variables in addition to the two volatility measures:

1. log GDP: OECD quarterly national accounts data. Gross domestic product at market prices - output approach (subject `B1_GA`) measured in national currency, chained volume estimates, national reference year, quarterly levels, seasonally adjusted (measure `LNBQRSA`), when available, otherwise gross domestic product - expenditure approach (subject `B1_GE`) measured in national currency, constant prices, national base year, quarterly levels, seasonally adjusted (measure `VNBQRSA`), divided by population.

2. Inflation: OECD Data Archive. Indicator: `Inflation (CPI)`. Subject: `Total`. Measure: `Annual growth rate (%)`.

3. Policy rate: Central bank policy rate from the BIS data portal, data set `BIS WS_CBPOL 1.0`. To get quarterly values, we average over monthly observations.

To construct per capita values, we use OECD historical population data (table `HISTPOP`). To obtain quarterly values, we linearly interpolate the annual data.

In the VAR, we use the countries listed in Table A.2. The second column reports the periods for which all variables are available.

Dealing with Outliers

We have two cases of large outliers in the stock market data that significantly distort the decomposition into country-specific and common components. These distortions arise not only for the country with the outlier observation but also lead to the country-specific component being orders of magnitude larger (in absolute value) than in the period with the second-largest value for some other countries.

The first of these cases is a series of extreme spikes in annualized volatility for Latvia from August to October 2001. Since our VAR for Latvia only uses data starting in 2005Q3 and these observations are only used for computing the first PC, we winsorize the data and set the corresponding values to the next-highest value of Latvian volatility.

The second is a permanent drop in the performance index of Cyprus by around 2/3 of the index occurring between two daily observations in August 2020. We are unaware of any event that could explain this permanent drop and do not find a collapse like this in other available Cypriot performance indices. In contrast to the

Table A.2: Country sample periods global sample

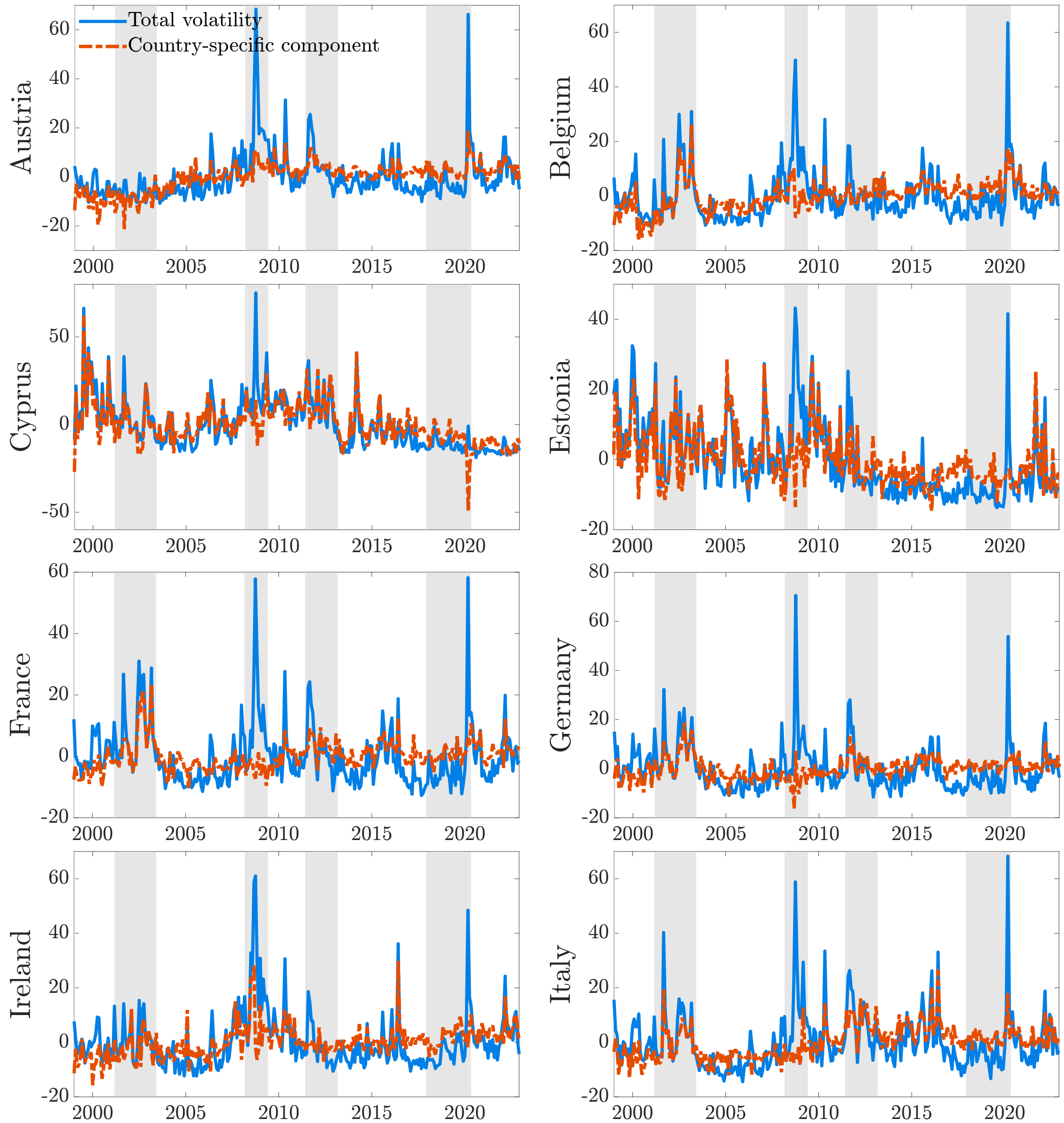
Country	Sample periods
Australia	1999Q1-2022Q4
Brazil	1999Q1-2022Q4
Canada	2002Q3-2022Q4
Chile	1999Q4-2022Q4
Colombia	2005Q1-2022Q4
India	2013Q1-2022Q4
Israel	1999Q1-2022Q4
Mexico	1999Q1-2022Q4
New Zealand	1999Q1-2022Q4
Russia	2003Q1-2021Q3
South Africa	1999Q1-2022Q4
South Korea	1999Q2-2022Q4
United States	1999Q1-2022Q4

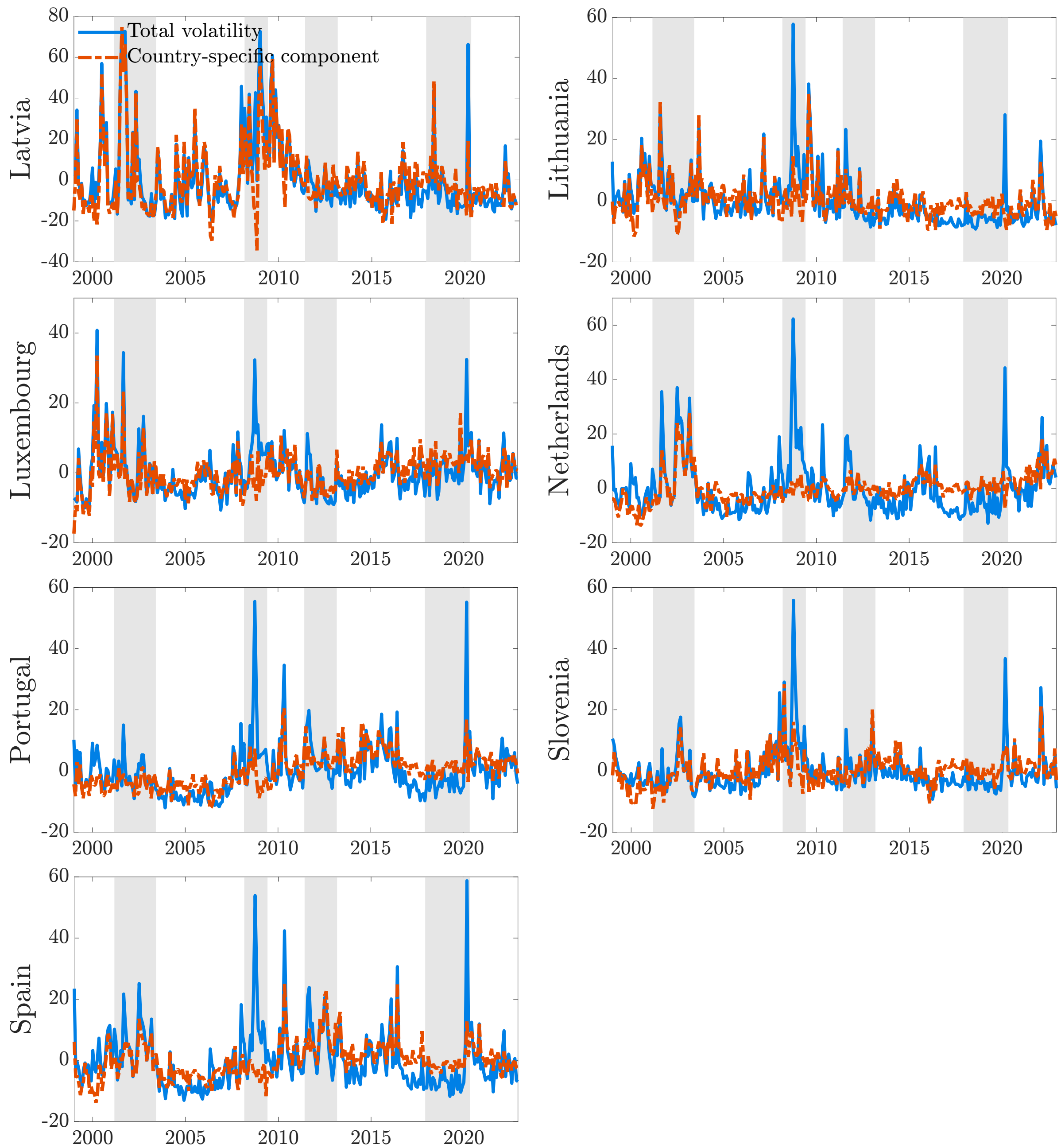
Notes: For each country listed, we use the periods where all variables are available and the country's exchange rate arrangement is classified as an 11 ("Moving band that is narrower than or equal to +/-2%") or higher by Ilzetzki et al. (2022). An exception is Russia, which is classified as a 10 for most of 2013 and 2014, but we use all periods as reported in the table.

Latvian case above, we actually need the value of this observation for our VAR, so we add the value of the observed drop-off to all days after the 'crash' in the rest of the month. The monthly observation that is used in the PCA is then the annualized return volatility over these partly fixed daily observations. We exclude Cyprus in the VARs, which include the stock market level (Figures B.10 and B.11).

B Further evidence

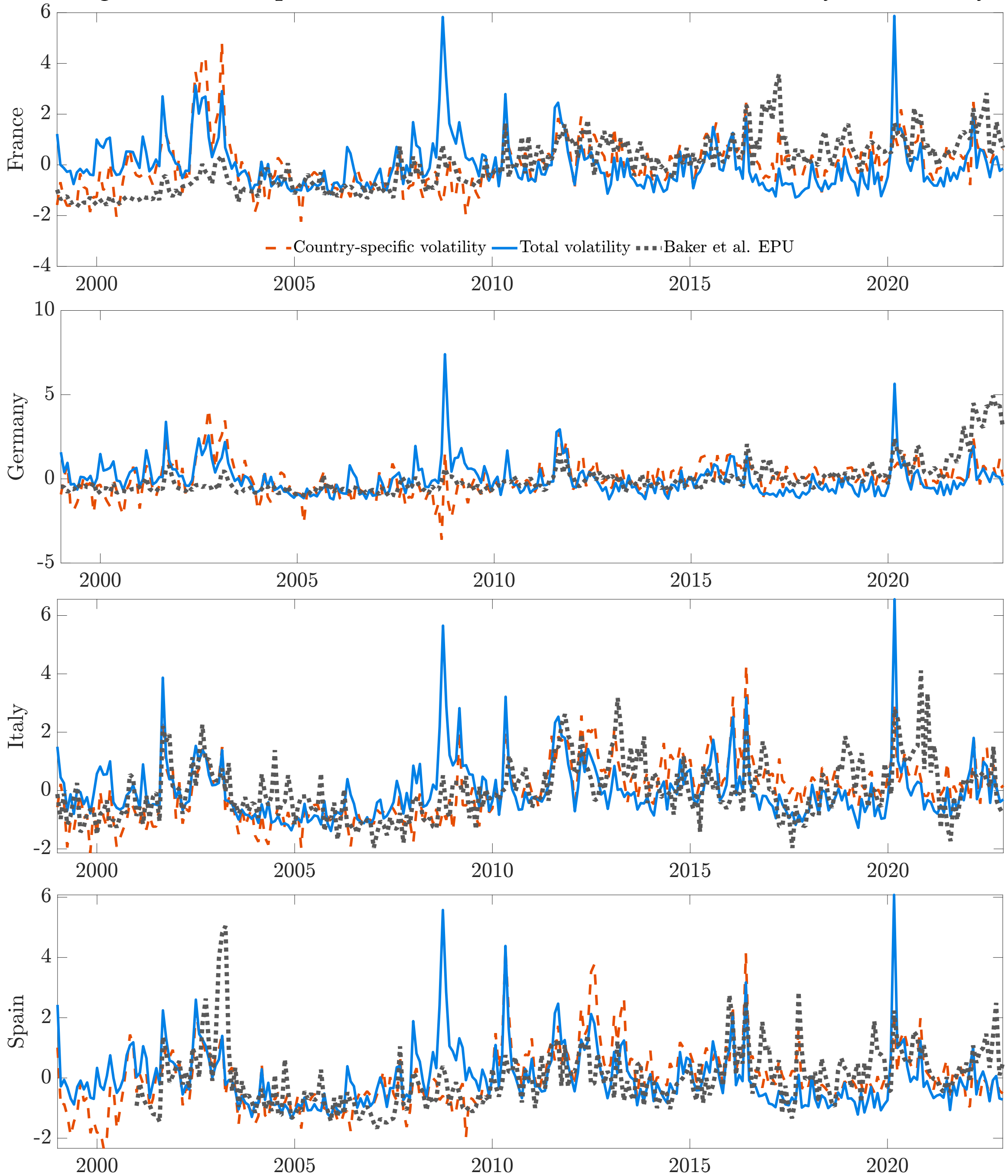
Figure B.1: Demeaned annualized stock market volatility components





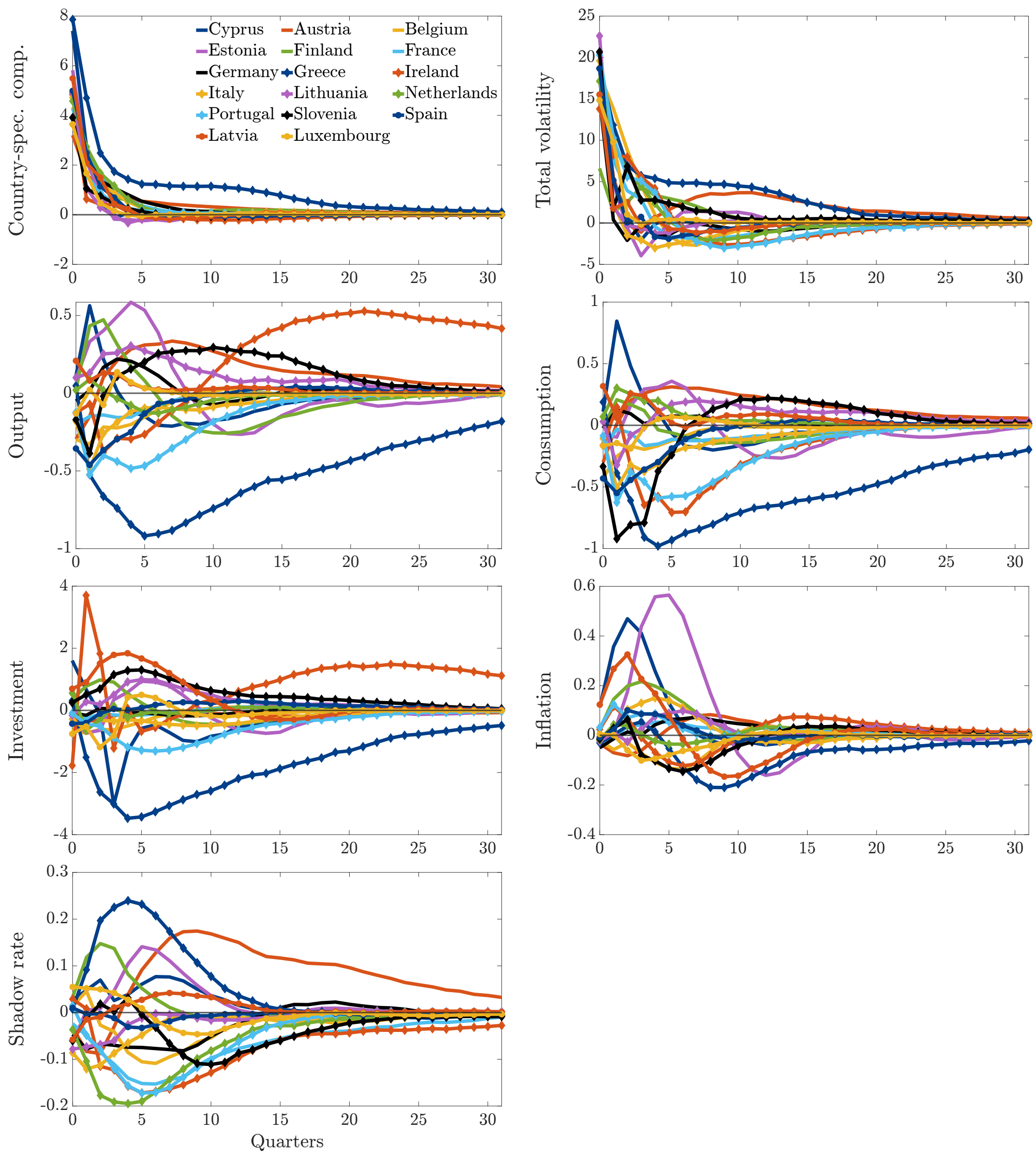
Notes: Monthly country-specific component (red dashed line) and total (blue solid line) realized volatility of annualized stock market returns in percent. All time series are demeaned. Shaded areas denote EA recession as dated by OECD-based recession indicators.

Figure B.2: Comparison to Baker et al. (2016) Economic Policy Uncertainty



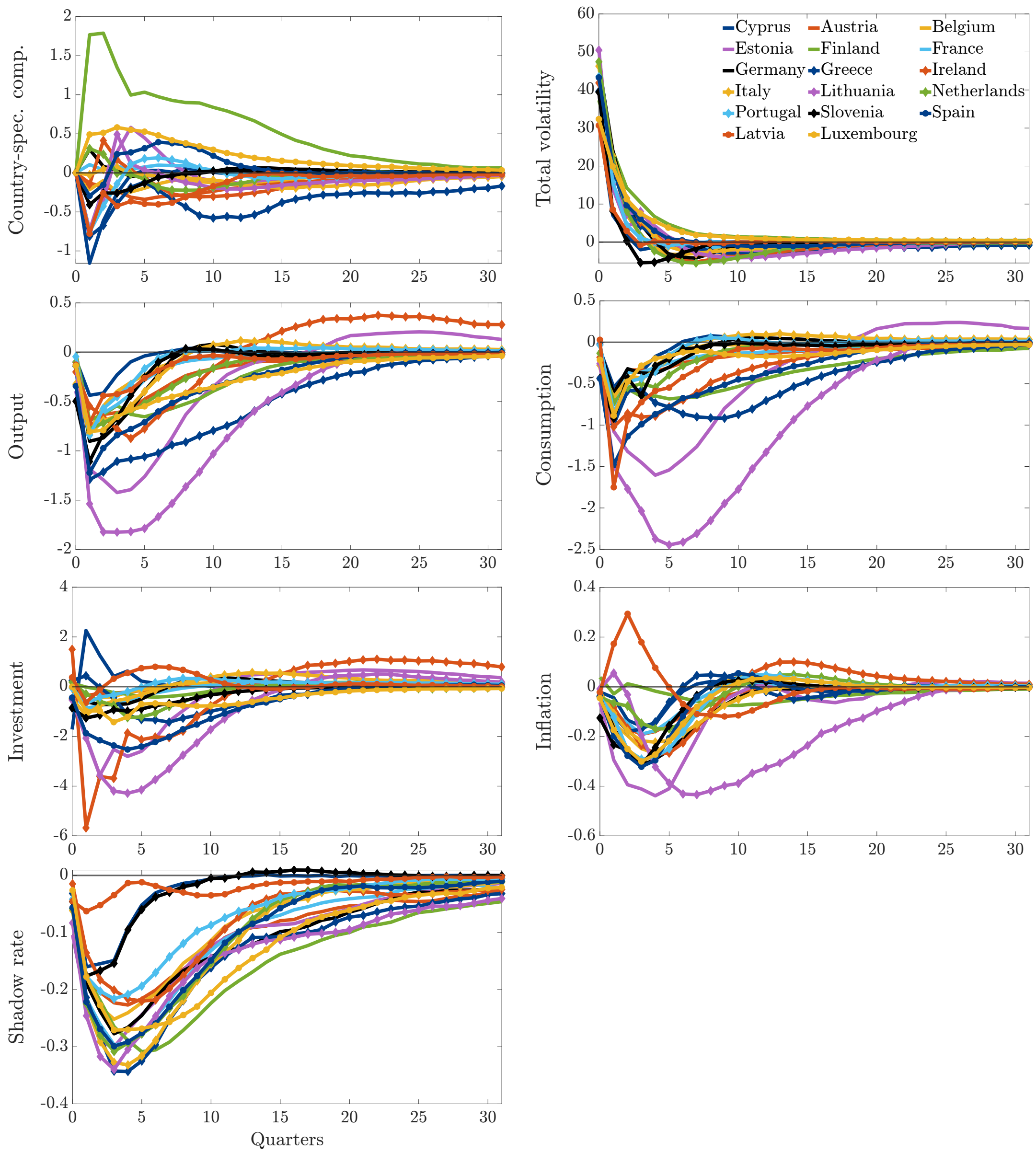
Notes: Comparison of our country-specific (red dashed line) and total volatility (blue solid line) measures with Baker et al. (2016) economic policy uncertainty index (grey dotted line), when available. All time series are z-scored.

Figure B.3: IRFs to country-specific uncertainty shock: country-level evidence (EA sample)



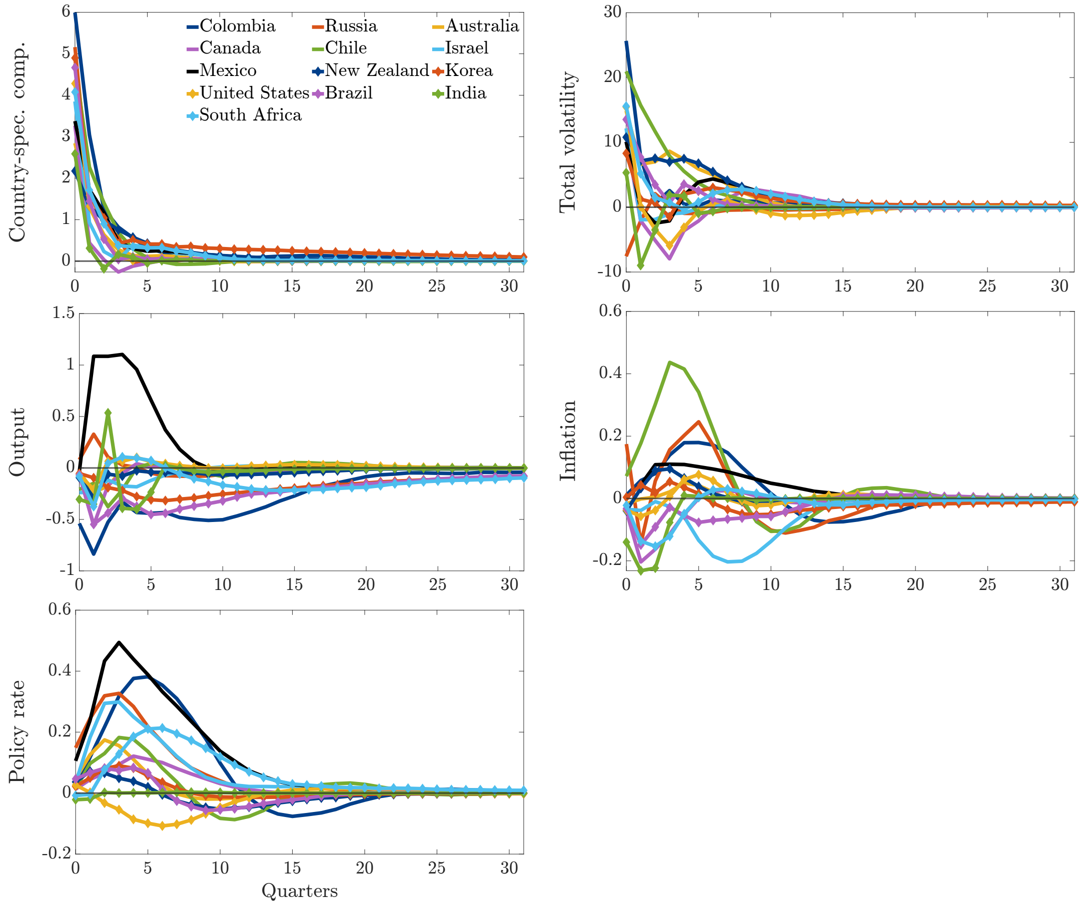
Notes: IRFs to one-standard-deviation country-specific uncertainty shock. Deviations in percent, except for country-specific component, inflation, and the shadow rate, which are in ppts.

Figure B.4: IRFs to common uncertainty shock: country-level evidence (EA sample)



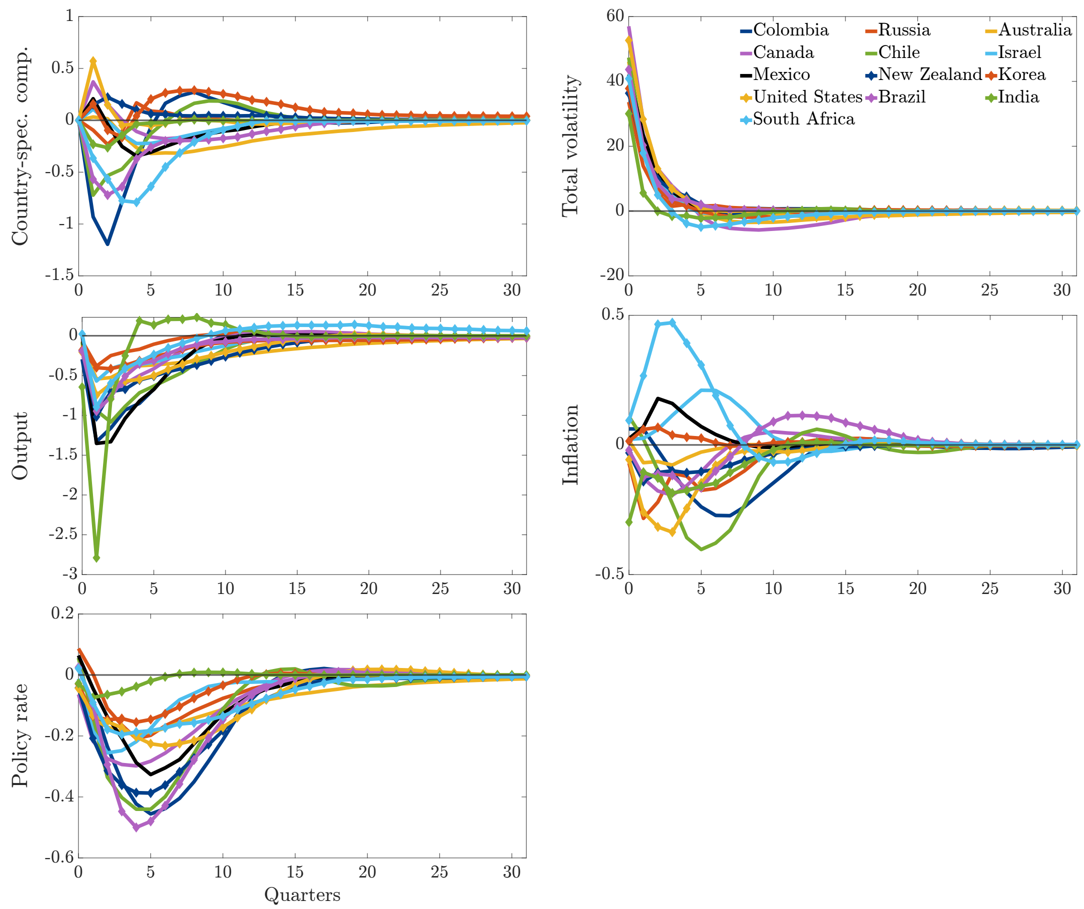
Notes: IRFs to one-standard-deviation common uncertainty shock. Deviations in percent, except for country-specific component, inflation, and the shadow rate, which are in ppts.

Figure B.5: IRFs to country-specific uncertainty shock: country-level evidence (floaters sample)



Notes: IRFs to one-standard-deviation country-specific uncertainty shock. Deviations in percent, except for country-specific component, inflation, and the policy rate, which are in ppts.

Figure B.6: IRFs to common uncertainty shock: country-level evidence (floaters sample)



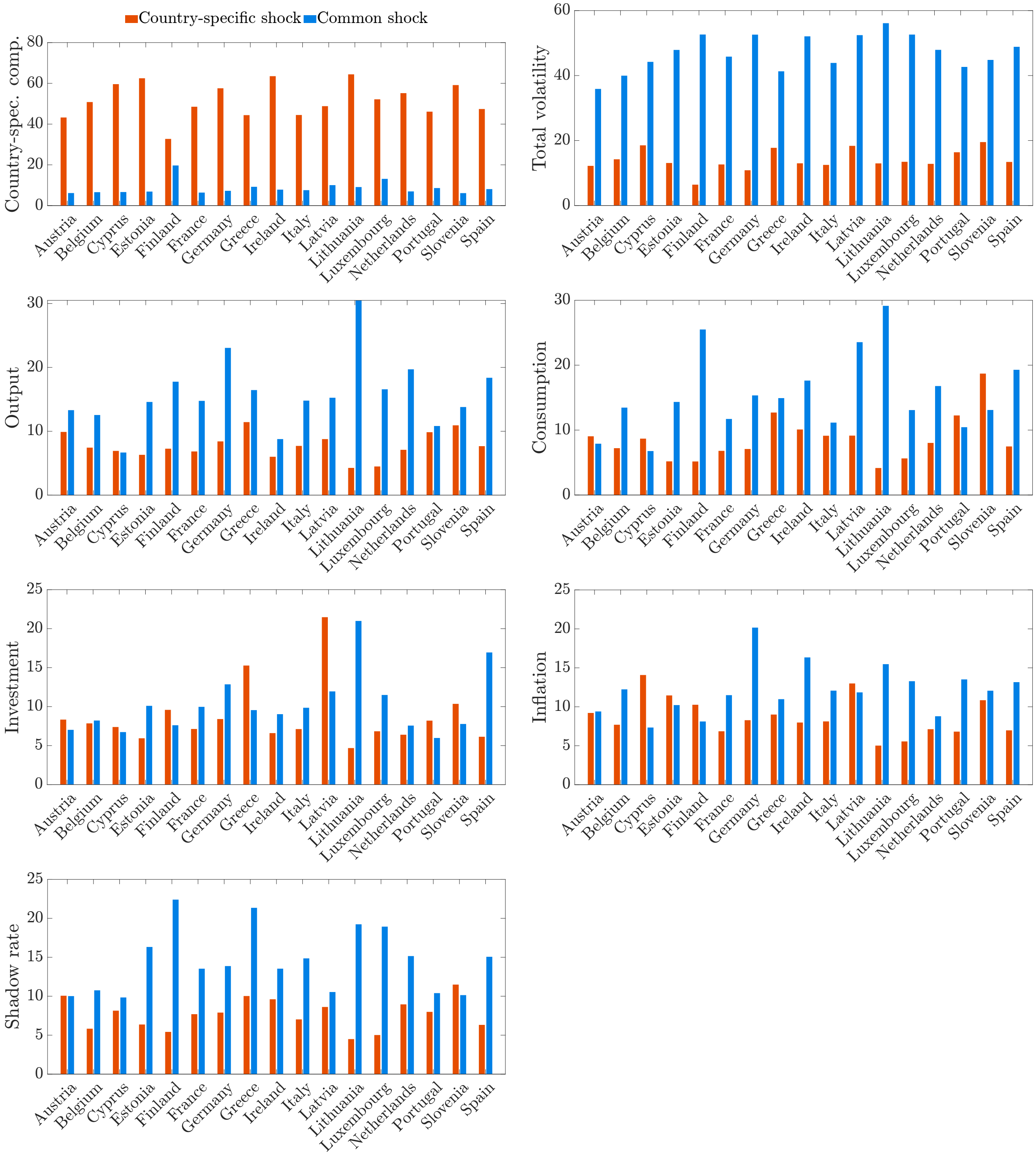
Notes: IRFs to one-standard-deviation common uncertainty shock. Deviations in percent, except for country-specific component, inflation, and the policy rate, which are in ppts.

Table B.1: Forecast error variance decomposition for median economy among global floaters

	Country-specific uncertainty shock	Common uncertainty shock
Country-specific component	74.16	7.68
	(69.64 , 78.53)	(5.7 , 10.28)
Total volatility	13.63	64.28
	(11.19 , 16.77)	(59.69 , 68.36)
Output	9.21	19.09
	(6.6 , 12.46)	(14.25 , 24.41)
Inflation	8.42	11.55
	(6.12 , 11.41)	(8.45 , 15.46)
Shadow rate	9.89	18.54
	(6.99 , 13.88)	(13.61 , 24.58)

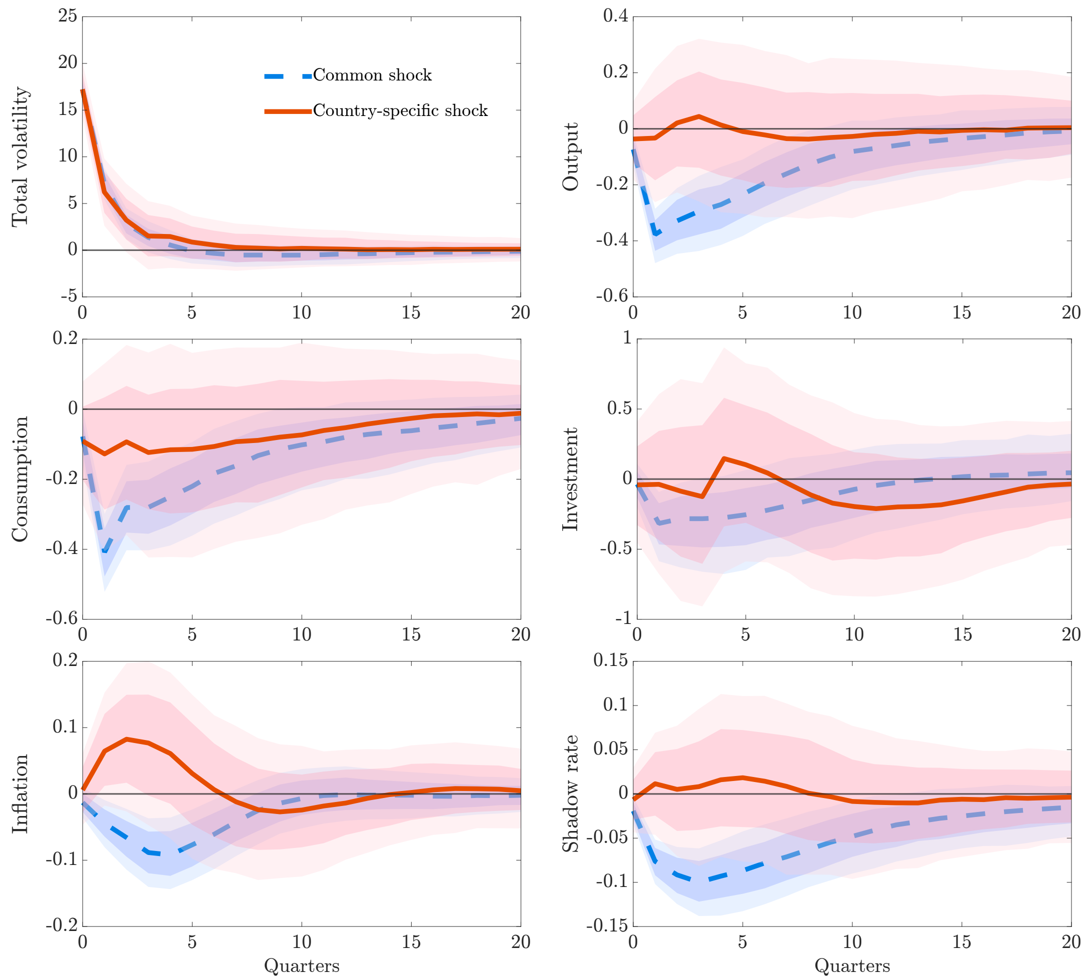
Notes: Contribution of country-specific (middle column) and common (right column) uncertainty shock to forecast error variance of each variable at horizon 20, in percent of total forecast error variance of that variable (with 68% HPDIs reported in parentheses).

Figure B.7: Forecast error variance decomposition at horizon 20



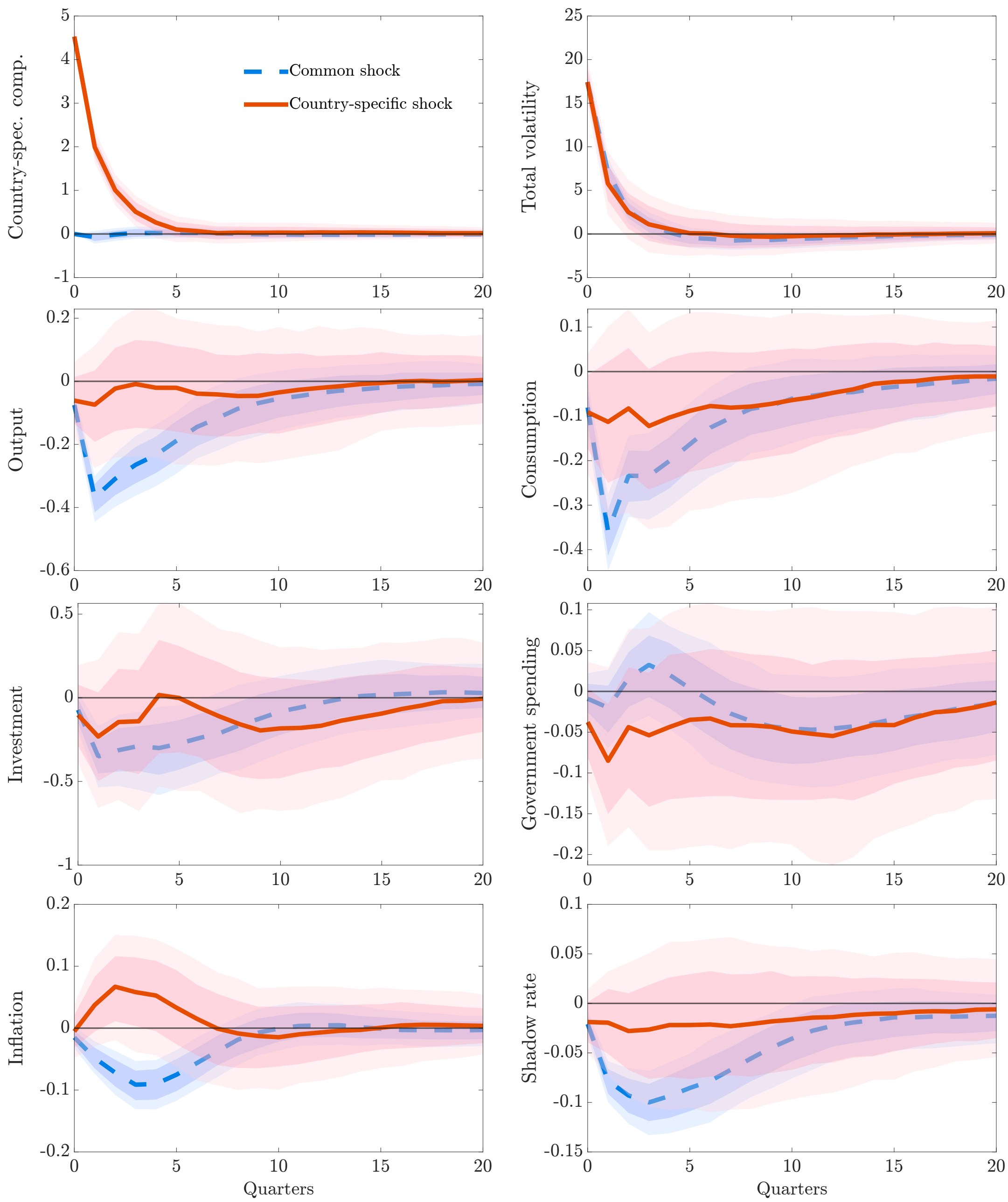
Notes: Contribution of country-specific (red) and common (blue) uncertainty shock to forecast error variance of each variable at horizon 20 as share of total forecast error variance of that variable. What we report here are median values for each country after computing the FEVD for 1000 random draws out of the posterior distribution of VAR coefficients.

Figure B.8: VAR robustness: small countries only



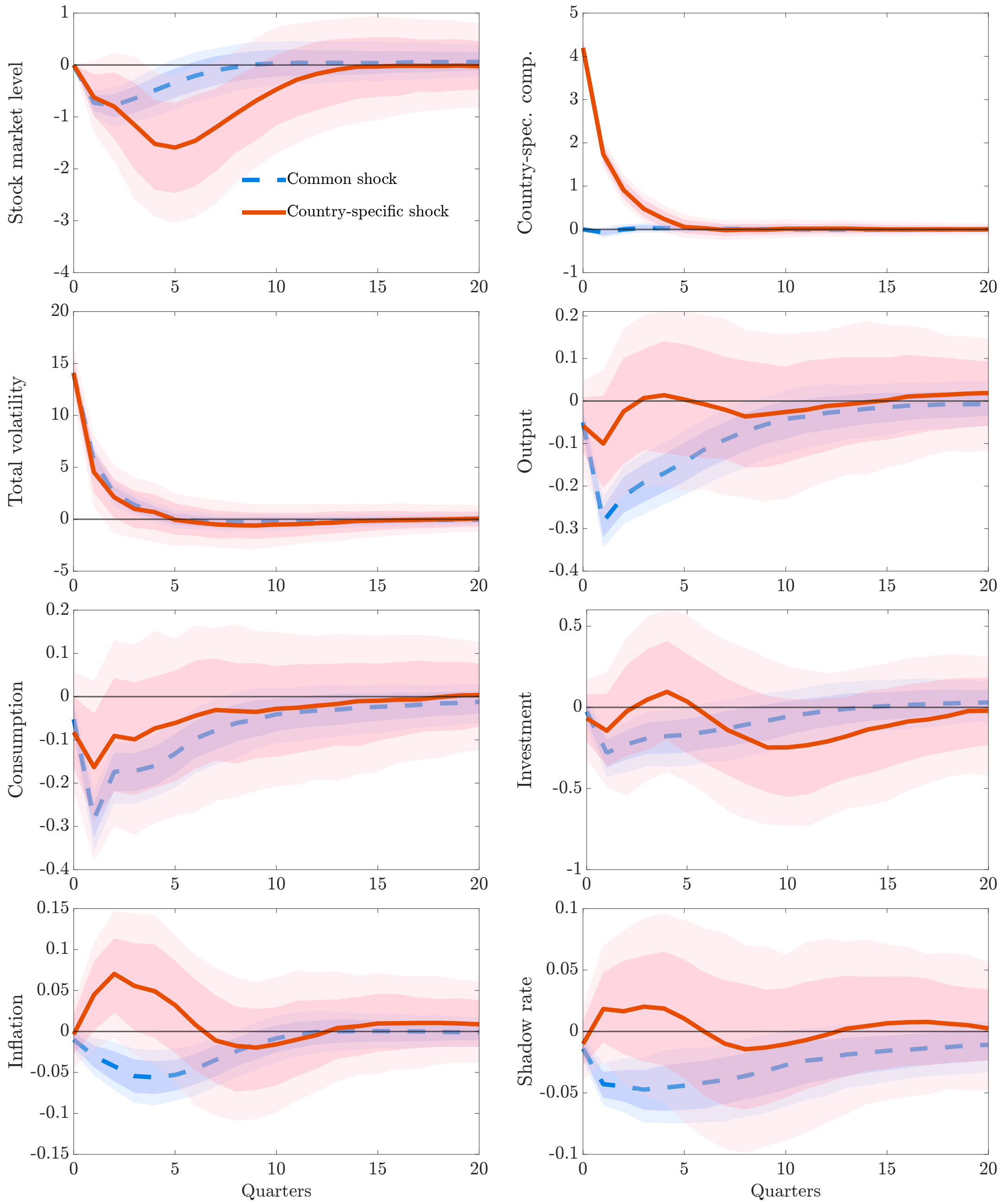
Notes: IRFs to one-standard-deviation country-specific (red solid lines) and equally-sized common (blue dashed lines) uncertainty shock. Shaded bands are pointwise 68% (dark) and 90% (light) HPDIs, respectively. Horizontal axis measures time in quarters, vertical axis measures deviations from pre-shock level in percent, except for inflation and the shadow rate (ppts). Country-specific volatility is included in the VAR, but not shown here. Sample only consists of countries that each make up less than 5 percent of aggregate output in the EA (excludes Germany, France, Italy, Spain, Netherlands).

Figure B.9: VAR robustness: include government spending



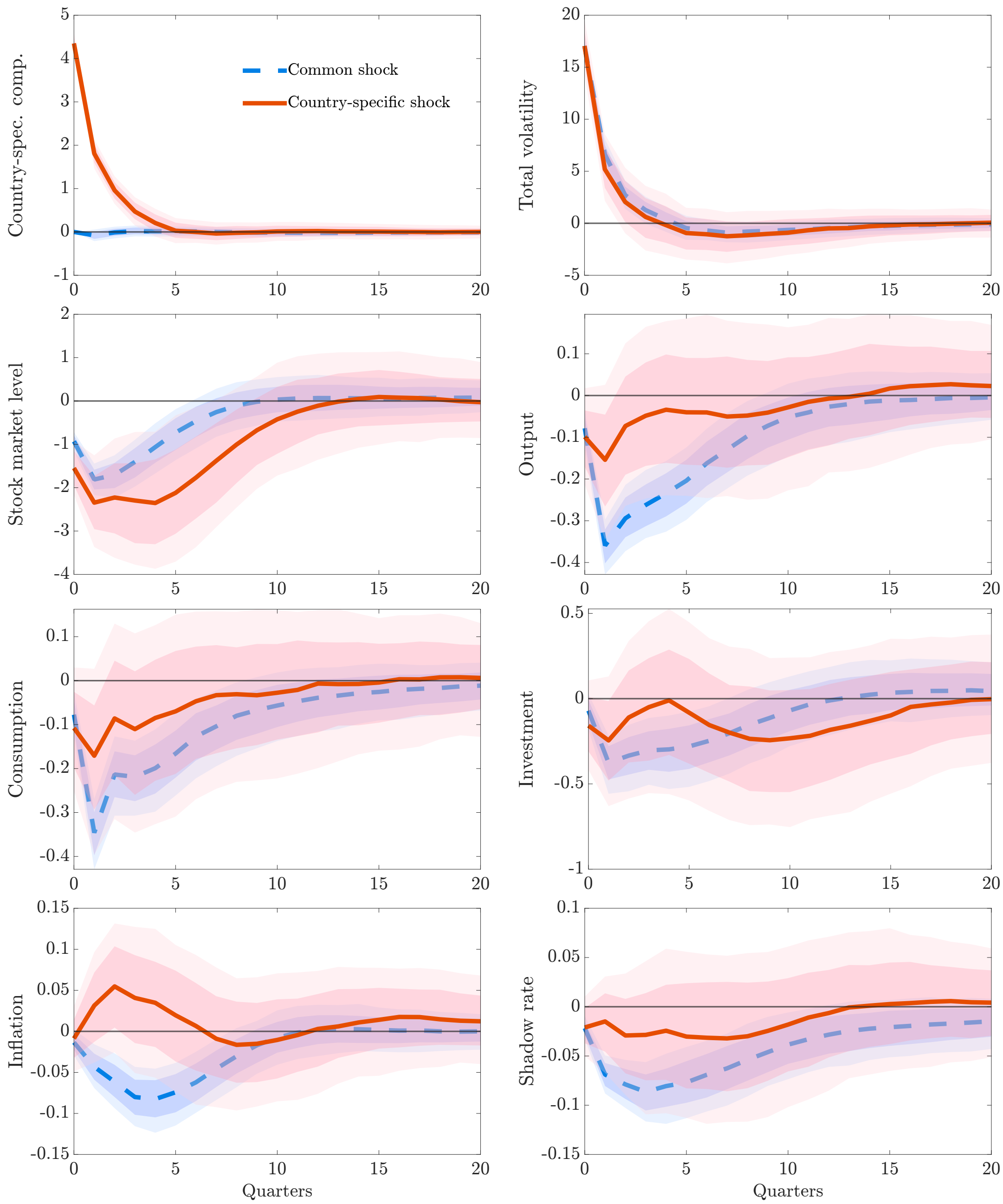
Notes: IRFs to one-standard-deviation country-specific (red solid lines) and equally-sized common (blue dashed lines) uncertainty shock. Shaded bands are pointwise 68% (dark) and 90% (light) HPDIs, respectively. Horizontal axis measures time in quarters, vertical axis measures deviations from pre-shock level in percent, except for country-specific component, inflation, and the shadow rate (ppts). VAR includes real per capita government consumption as additional variable.

Figure B.10: VAR robustness: include stock market level as first variable



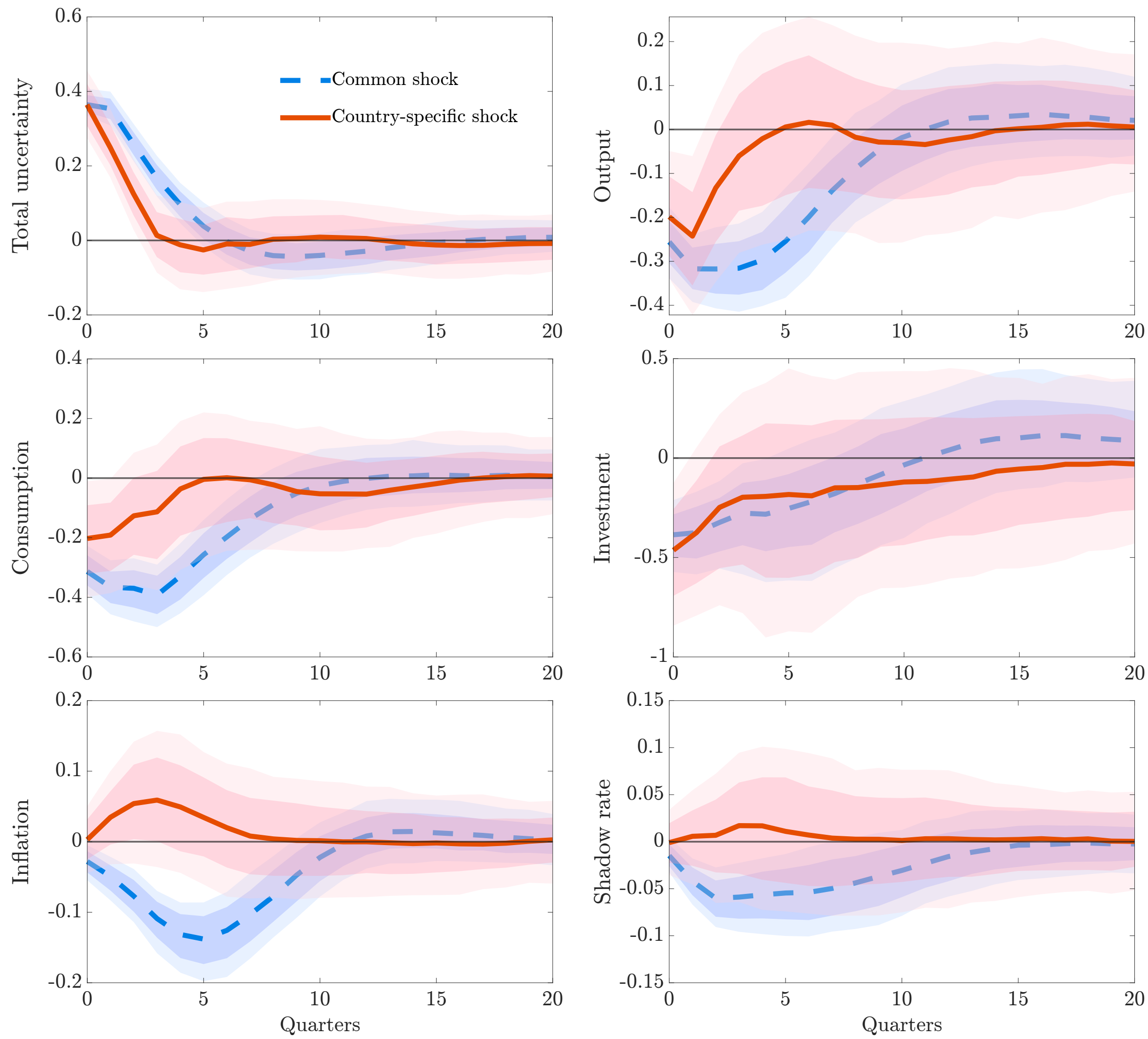
Notes: IRFs to one-standard-deviation country-specific (red solid lines) and equally-sized common (blue dashed lines) uncertainty shock. Shaded bands are pointwise 68% (dark) and 90% (light) HPDIs, respectively. Horizontal axis measures time in quarters, vertical axis measures deviations from pre-shock level in percent, except for country-specific component, inflation, and the shadow rate (ppts).

Figure B.11: VAR robustness: include stock market level third



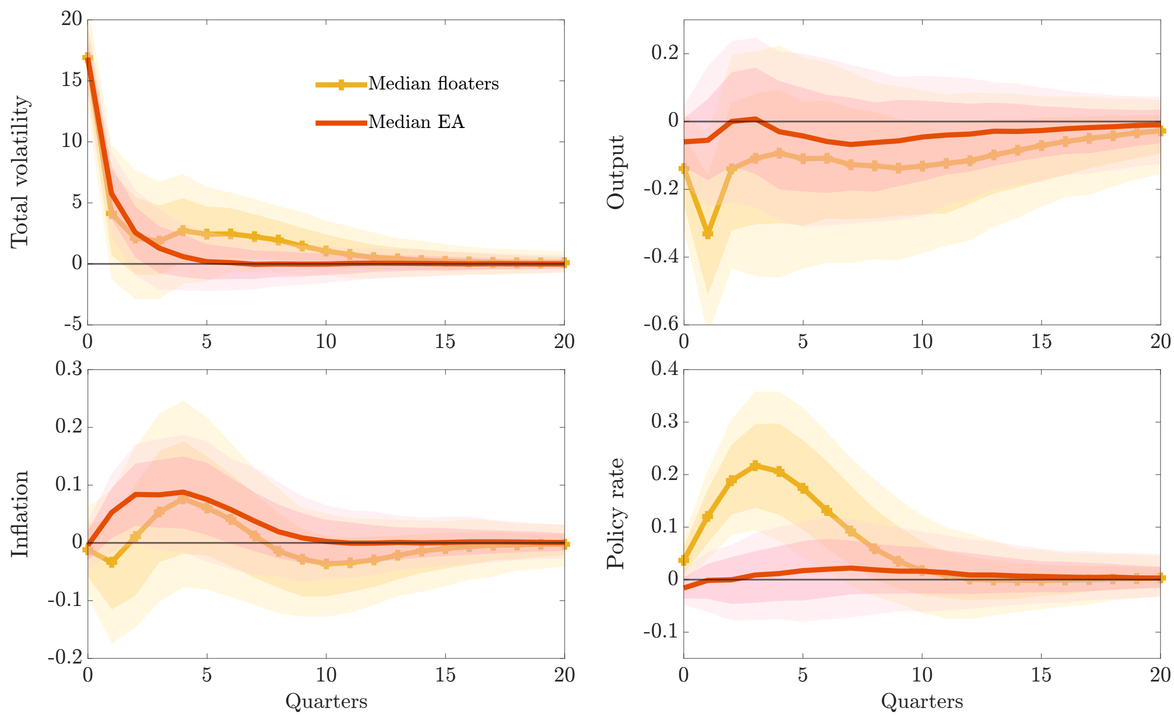
Notes: IRFs to one-standard-deviation country-specific (red solid lines) and equally-sized common (blue dashed lines) uncertainty shock. Shaded bands are pointwise 68% (dark) and 90% (light) HPDIs, respectively. Horizontal axis measures time in quarters, vertical axis measures deviations from pre-shock level in percent, except for country-specific component, inflation, and the shadow rate (ppts).

Figure B.12: VAR robustness: Jurado et al. (2015) uncertainty measure



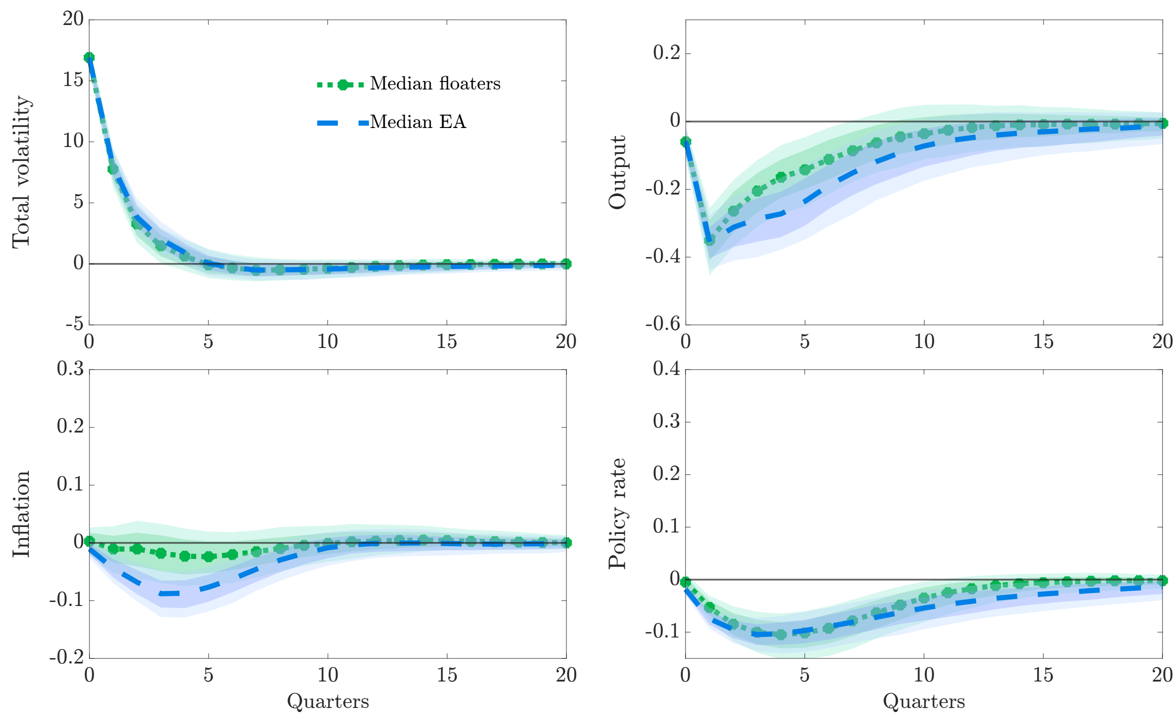
Notes: IRFs to one-standard-deviation country-specific (red solid lines) and equally-sized common (blue dashed lines) uncertainty shock. Shaded bands are pointwise 68% (dark) and 90% (light) HPDIs, respectively. Horizontal axis measures time in quarters, vertical axis measures deviations from pre-shock level in percent, except for inflation and the shadow rate (ppts). Stock market volatility replaced by forecast error-based macroeconomic uncertainty measure developed by Jurado et al. (2015) and provided by Comunale and Nguyen (2023) for all EA countries. Country-specific uncertainty is included in the VAR, but not shown here.

Figure B.13: VAR robustness: Exclude USA — country-specific uncertainty shock



Notes: Impulse responses to country-specific uncertainty shocks in the EA (red solid line) and among global floaters (yellow solid line with plus-shaped markers). Shock size rescaled so that the median impact on total volatility equals that of one-standard deviation country-specific uncertainty shock in the EA. Shaded areas indicate point-wise 68% (dark) and 90% (light) HPDIs, respectively. Horizontal axis measures time in quarters, vertical axis measures deviations from pre-shock level in percent, except for inflation and policy rate (ppts). Country-specific volatility is included in the VAR, but not shown here.

Figure B.14: VAR robustness: Exclude USA — common uncertainty shock



Notes: Impulse responses to common uncertainty shocks in the EA (blue dashed line) and among global floaters (green dotted line with octagonal markers). Shock sizes rescaled so that the median impact on total volatility equals that of one-standard deviation country-specific uncertainty shock in the EA. Shaded areas indicate point-wise 68% (dark) and 90% (light) HPDIs, respectively. Horizontal axis measures time in quarters, vertical axis measures deviations from pre-shock level in percent, except for inflation and policy rate (ppts). Country-specific volatility is included in the VAR, but not shown here.

C Model appendix

C.1 Definitions and derivations

Production

Competitive final good firms produce the final good $\mathcal{F}_t$ with price P_t,[22] using a CES aggregator:

$$\mathcal{F}_t = \left[(1-(1-n)v)^{\frac{1}{\eta}} \left(Y_{H,t}\right)^{\frac{\eta-1}{\eta}} + ((1-n)v)^{\frac{1}{\eta}} \left(Y_{F,t}\right)^{\frac{\eta-1}{\eta}}\right]^{\frac{\eta}{\eta-1}} \quad , \tag{3.1}$$

by bundling domestic goods $Y_{H,t}$ and foreign imported goods $Y_{F,t}$ to minimize their expenditure $P_t^H Y_t^H + P_t^F Y_t^F$ given demand and prices. In the following, we describe the setup and first-order conditions for Home goods, with equivalent considerations for Foreign goods. The first order condition for Home goods is

$$Y_{H,t} = (1-(1-n)v)\left(\frac{P_{H,t}}{P_t}\right)^{-\eta} \mathcal{F}_t \; . \tag{C.1}$$

The domestic good $Y_{H,t}$ is assembled from a continuum of differentiated intermediate inputs $Y_t(i)$, $i \in [0, n]$, using the constant returns to scale Dixit-Stiglitz-technology

$$Y_{H,t} = \left[\left(\frac{1}{n}\right)^{\frac{1}{\epsilon}} \int_0^n Y_t(i)^{\frac{\epsilon-1}{\epsilon}}\, di\right]^{\frac{\epsilon}{\epsilon-1}} \quad , \tag{C.2}$$

where $\epsilon > 1$ is the elasticity of substitution between intermediate goods. The optimal amount of inputs $Y_t(i)$, given their price $P_t(i)$, is determined by solving the following expenditure minimization problem:

$$\min_{Y_t(i)} \int_0^n P_t(i) Y_t(i) di + P_{H,t}\left[Y_{H,t} - \left[\left(\frac{1}{n}\right)^{\frac{1}{\epsilon}} \int_0^n Y_t(i)^{\frac{\epsilon-1}{\epsilon}} di\right]^{\frac{\epsilon}{\epsilon-1}}\right] \; . \tag{C.3}$$

Here, $P_{H,t}$ is the Lagrange multiplier, which has a natural interpretation as the price index for $Y_{H,t}$. The first order condition for each variety i is given by $Y_t(i) = \frac{1}{n}\left(\frac{P_t(i)}{P_{H,t}}\right)^{-\epsilon} Y_{H,t}$. Substituting for $Y_t(i)$ in (C.2), shows that $P_{H,t} =$

[22] We consider the final good as the numéraire and set its initial pre-shock value to 1.

$\left[\frac{1}{n}\int_0^n P_t(i)^{1-\epsilon}dj\right]^{\frac{1}{1-\epsilon}}$. Equivalent considerations apply to imported inputs $Y_t(j)$, produced by foreign intermediate goods firms $j \in (n,1]$. The Home demand for varieties produced in Home and Foreign is then given by

$$Y_t(i) = \frac{1}{n}\left(\frac{P_t(i)}{P_{H,t}}\right)^{-\epsilon} Y_{H,t}\ , \quad Y_t(j) = \frac{1}{1-n}\left(\frac{P_t(j)}{P_{F,t}}\right)^{-\epsilon} Y_{F,t}\ , \tag{C.4}$$

for $i \in [0,n]$ and $j \in (n,1]$, respectively, and where $P_{F,t} = \left[\frac{1}{1-n}\int_n^1 P_t(j)^{1-\epsilon}dj\right]^{\frac{1}{1-\epsilon}}$.

Substituting these expressions into (3.1) allows deriving the domestic CPI, equation (3.2). Substituting for $Y_{H,t}$ from (C.1) and its counterpart for $Y_{F,t}$ in the domestic demand function for varieties (C.4), we obtain the Home demand for Home and Foreign intermediates, respectively:

$$Y_t(i) = \frac{1-(1-n)v}{n}\left(\frac{P_t(i)}{P_{H,t}}\right)^{-\epsilon}\left(\frac{P_{H,t}}{P_t}\right)^{-\eta}\mathcal{F}_t\ , \tag{C.5}$$

$$\begin{aligned} Y_t(j) &= \frac{(1-n)v}{1-n}\left(\frac{P_t(j)}{P_{F,t}}\right)^{-\epsilon}\left(\frac{P_{F,t}}{P_t}\right)^{-\eta}\mathcal{F}_t \\ &= \left(\frac{P_t^*(j)}{P_{F,t}^*}\right)^{-\epsilon}\left(\frac{P_{F,t}^*}{P_t^*}\right)^{-\eta} v\mathcal{Q}_t^{-\eta}\mathcal{F}_t\ . \end{aligned} \tag{C.6}$$

The last equality makes use of the law of one price and the real exchange rate definition $\mathcal{Q}_t \equiv \mathcal{E}_t P_t^*/P_t$. Due to symmetry, foreign demand for Home and Foreign intermediates, respectively, is given by:

$$\begin{aligned} Y_t^*(i) &= \frac{nv}{n}\left(\frac{P_t^*(i)}{P_{H,t}^*}\right)^{-\epsilon}\left(\frac{P_{H,t}^*}{P_t^*}\right)^{-\eta}\mathcal{F}_t^* \\ &= \left(\frac{P_t(i)}{P_{H,t}}\right)^{-\epsilon}\left(\frac{P_{H,t}}{P_t}\right)^{-\eta} v\mathcal{Q}_t^{\eta}\mathcal{F}_t^*, \end{aligned} \tag{C.7}$$

$$Y_t^*(j) = \frac{1-nv}{1-n}\left(\frac{P_t^*(j)}{P_{F,t}^*}\right)^{-\epsilon}\left(\frac{P_{F,t}^*}{P_t^*}\right)^{-\eta}\mathcal{F}_t^*. \tag{C.8}$$

Global demand $Y_t^d(h)$ for a generic intermediate good $h \in [0,1]$ is the weighted average of domestic and foreign demand for this variety:

$$Y_t^d(h) = nY_t(h) + (1-n)Y_t^*(h)\,. \tag{C.9}$$

Summing up the respective Home and Foreign demand components, global demand for domestic and foreign varieties, respectively, is then given by:

$$Y_t^d(i) = \left(\frac{P_t(i)}{P_{H,t}}\right)^{-\epsilon} \left\{ \left(\frac{P_{H,t}}{P_t}\right)^{-\eta} \left[(1-(1-n)v)\mathcal{F}_t + (1-n)v\mathcal{Q}_t^{\eta}\mathcal{F}_t^*\right] \right\}, \quad \text{(C.10)}$$

$$Y_t^d(j) = \left(\frac{P_t^*(j)}{P_{F,t}^*}\right)^{-\epsilon} \left\{ \left(\frac{P_{F,t}^*}{P_t^*}\right)^{-\eta} \left[nv\mathcal{Q}_t^{-\eta}\mathcal{F}_t + (1-nv)\mathcal{F}_t^*\right] \right\}, \quad \text{(C.11)}$$

where $P_t^* = \left[nv\left(P_{H,t}^*\right)^{1-\eta} + (1-nv)\left(P_{F,t}^*\right)^{1-\eta}\right]^{\frac{1}{1-\eta}}$. Using (C.10) and (C.11), aggregate Home and Foreign output per capita is then given by

$$\begin{aligned} Y_t &= \left[\frac{1}{n}\int_0^n Y_t^d(i)^{\frac{\epsilon-1}{\epsilon}} di\right]^{\frac{\epsilon}{\epsilon-1}} && \text{(C.12)} \\ &= \left(\frac{P_{H,t}}{P_t}\right)^{-\eta} \left[(1-(1-n)v)\mathcal{F}_t + (1-n)v\mathcal{Q}_t^{\eta}\mathcal{F}_t^*\right] , \\ Y_t^* &= \left[\frac{1}{1-n}\int_n^1 Y_t^d(j)^{\frac{\epsilon-1}{\epsilon}} dj\right]^{\frac{\epsilon}{\epsilon-1}} = \left(\frac{P_{F,t}^*}{P_t^*}\right)^{-\eta} \left[nv\mathcal{Q}_t^{-\eta}\mathcal{F}_t + (1-nv)\mathcal{F}_t^*\right] . && \text{(C.13)} \end{aligned}$$

In the limiting case of $n \to 0$ (which implies $P_{F,t}^* = P_t^*$), we get

$$Y_t = \left(\frac{P_{H,t}}{P_t}\right)^{-\eta} \left[(1-v)\mathcal{F}_t + v\mathcal{Q}_t^{\eta}\mathcal{F}_t^*\right] \quad \text{and} \quad Y_t^* = \mathcal{F}_t^* . \quad \text{(C.14)}$$

Combining this expression with price adjustment costs and the Home terms of trade definition $\mathcal{S}_t = P_{F,t}/P_{H,t}$ yields the resource constraints (3.17) & (3.18) reported in the main text. The real exchange rate in the limiting case is linked to $\mathcal{S}_t$ via

$$\mathcal{Q}_t = \left[(1-v)\mathcal{S}_t^{\eta-1} + v\right]^{-\frac{1}{1-\eta}} . \quad \text{(C.15)}$$

Defining Consumer Price Index (CPI) inflation Π_t as P_t/P_{t-1} and Producer Price Index (PPI) inflation as $\Pi_{H,t} = P_{H,t}/P_{H,t-1}$, equation (3.2) implies that PPI and CPI are linked via

$$\Pi_t^{1-\eta} = (1-v)\left(\Pi_{H,t}S_{t-1}^{-1}Q_{t-1}\right)^{1-\eta} + v(Q_t\Pi_t)^{1-\eta} . \quad \text{(C.16)}$$

Households

We assume an endogenous discount factor that decreases in the consumption output ratio (see, e.g., Kollmann, 2016):

$$\beta_t = \bar{\beta}\left[1 - \phi_B\left(\frac{C_t}{Y_t} - \frac{C}{Y}\right)\right], \tag{C.17}$$

where $0 \leq \bar{\beta} \leq 1$ is the pure discount factor and ϕ_B measures the slope of the discount factor.

Similarly to equation (3.10) for Home, the household in Foreign has the value function

$$V_t^* = \max\left[(1-\beta_t^*)\left(\xi_{C,t}\left(C_t^*\right)^{\varphi}\left(1-N_t^*\right)^{1-\varphi}\right)^{\frac{1-\sigma}{\theta_V}} + \beta_t^*\left(\mathbb{E}_t\left[\left(V_{t+1}^*\right)^{1-\sigma}\right]\right)^{\frac{1}{\theta_V}}\right]^{\frac{\theta_V}{1-\sigma}}, \tag{C.18}$$

which is also subject to common but not country-specific fluctuations in the demand shifter; β_t^* is defined analogously to β_t in (C.17).

The Euler equations of the Home household for Home and Foreign bonds, respectively, are

$$1 = \mathbb{E}_t\left[M_{t,t+1}\frac{R_t}{\Pi_{t+1}}\right] \quad \text{and} \quad 1 = \mathbb{E}_t\left[M_{t,t+1}\frac{R_t^*\mathcal{E}_{t+1}}{\mathcal{E}_t\Pi_{t+1}}\right], \tag{C.19}$$

where $M_{t,t+1}$ is the stochastic discount factor derived below. The Foreign household's Euler equation for Foreign bonds is

$$1 = \mathbb{E}_t\left[M_{t,t+1}^*\frac{R_t^*}{\Pi_{t+1}^*}\right]. \tag{C.20}$$

C.2 Deriving the Stochastic Discount Factor

The stochastic discount factor is given by

$$M_{t,t+1} \equiv \frac{\partial V_t/\partial C_{t+1}}{\partial V_t/\partial C_t}, \tag{C.21}$$

where

$$\frac{\partial V}{\partial C_t} = V_t^{1-\frac{1-\sigma}{\theta_V}}\varphi(1-\beta_t)\frac{\left(\xi_{H,t}\xi_{C,t}C_t^{\varphi}(1-N_t)^{1-\varphi}\right)^{\frac{1-\sigma}{\theta_V}}}{C_t} \tag{C.22}$$

and, using the Benveniste-Scheinkman envelope theorem,

$$\begin{aligned}
\frac{\partial V_t}{\partial C_{t+1}} =& \frac{\theta_V}{1-\sigma}\left((1-\beta_t)\left(\xi_{H,t}\xi_{C,t}C_t^{\varphi}(1-N_t)^{1-\varphi}\right)^{\frac{1-\sigma}{\theta_V}} + \beta_t\left(\mathbb{E}_t V_{t+1}^{1-\sigma}\right)^{\frac{1}{\theta_V}}\right)^{\frac{\theta_V}{1-\sigma}-1} \\
&\times \beta_t \frac{1}{\theta_V}\left(\mathbb{E}_t V_{t+1}^{1-\sigma}\right)^{\frac{1}{\theta_V}-1}\mathbb{E}_t\left((1-\sigma)V_{t+1}^{-\sigma}\frac{\partial V_{t+1}}{\partial C_{t+1}}\right) \\
\overset{(C.22)}{=}& V_t^{1-\frac{1-\sigma}{\theta_V}}\beta_t\left(\mathbb{E}_t V_{t+1}^{1-\sigma}\right)^{\frac{1}{\theta_V}-1} \\
&\times \mathbb{E}_t\left(V_{t+1}^{-\sigma}V_{t+1}^{1-\frac{1-\sigma}{\theta_V}}\varphi(1-\beta_{t+1})\frac{\left(\xi_{H,t+1}\xi_{C,t+1}C_{t+1}^{\varphi}(1-N_{t+1})^{1-\varphi}\right)^{\frac{1-\sigma}{\theta_V}}}{C_{t+1}}\right).
\end{aligned} \tag{C.23}$$

Thus,

$$\begin{aligned}
M_{t,t+1} \equiv \frac{\frac{\partial V_t}{\partial C_{t+1}}}{\frac{\partial V}{\partial C_t}} =& \beta_t \mathbb{E}_t \frac{1-\beta_{t+1}}{1-\beta_t} \\
&\times \left(\frac{\xi_{H,t+1}\xi_{C,t+1}C_{t+1}^{\varphi}(1-N_{t+1})^{1-\varphi}}{\xi_{H,t}\xi_{C,t}C_t^{\varphi}(1-N_t)^{1-\varphi}}\right)^{\frac{1-\sigma}{\theta_V}}\frac{C_t}{C_{t+1}}\left(\frac{V_{t+1}^{1-\sigma}}{\mathbb{E}_t V_{t+1}^{1-\sigma}}\right)^{1-\frac{1}{\theta_V}}.
\end{aligned} \tag{C.24}$$

C.3 VXO

The VXO is given by 100 times the square root of the annualized conditional equity return variance under a risk-neutral measure:

$$\begin{aligned}
VXO &= 100\sqrt{4Var_t^{RN}(R_{t+1})} \\
&= 100\sqrt{4\left[\mathbb{E}_t^{RN}\left(\left[R_{t+1}^E\right]^2\right) - \left[\mathbb{E}_t^{RN}\left(R_{t+1}^E\right)\right]^2\right]},
\end{aligned} \tag{C.25}$$

where the superscript RN denotes the risk-neutral measure as opposed to the physical measure under which the regular expectations are computed. Equity returns are given by

$$R_t^E = \frac{\frac{P_t^E}{P_t} + \frac{D_t^E}{P_t}}{\frac{P_{t-1}^E}{P_{t-1}}}. \tag{C.26}$$

Under a risk-neutral measure, every asset returns the risk-free rate $R_t^{RF} =$

$1/\mathbb{E}_t M_{t+1}$ in expectations. Therefore, the following identities need to hold:

$$\mathbb{E}_t\left(M_{t+1}R^E_{t+1}\right) = \mathbb{E}_t\left(M_{t+1}\right)\mathbb{E}^{RN}_t\left(R^E_{t+1}\right)\ , \tag{C.27}$$

$$\mathbb{E}_t\left(M_{t+1}(R^E_{t+1})^2\right) = \mathbb{E}_t\left(M_{t+1}\right)\mathbb{E}^{RN}_t\left((R^E_{t+1})^2\right)\ . \tag{C.28}$$

This can be used to rewrite (C.25) as

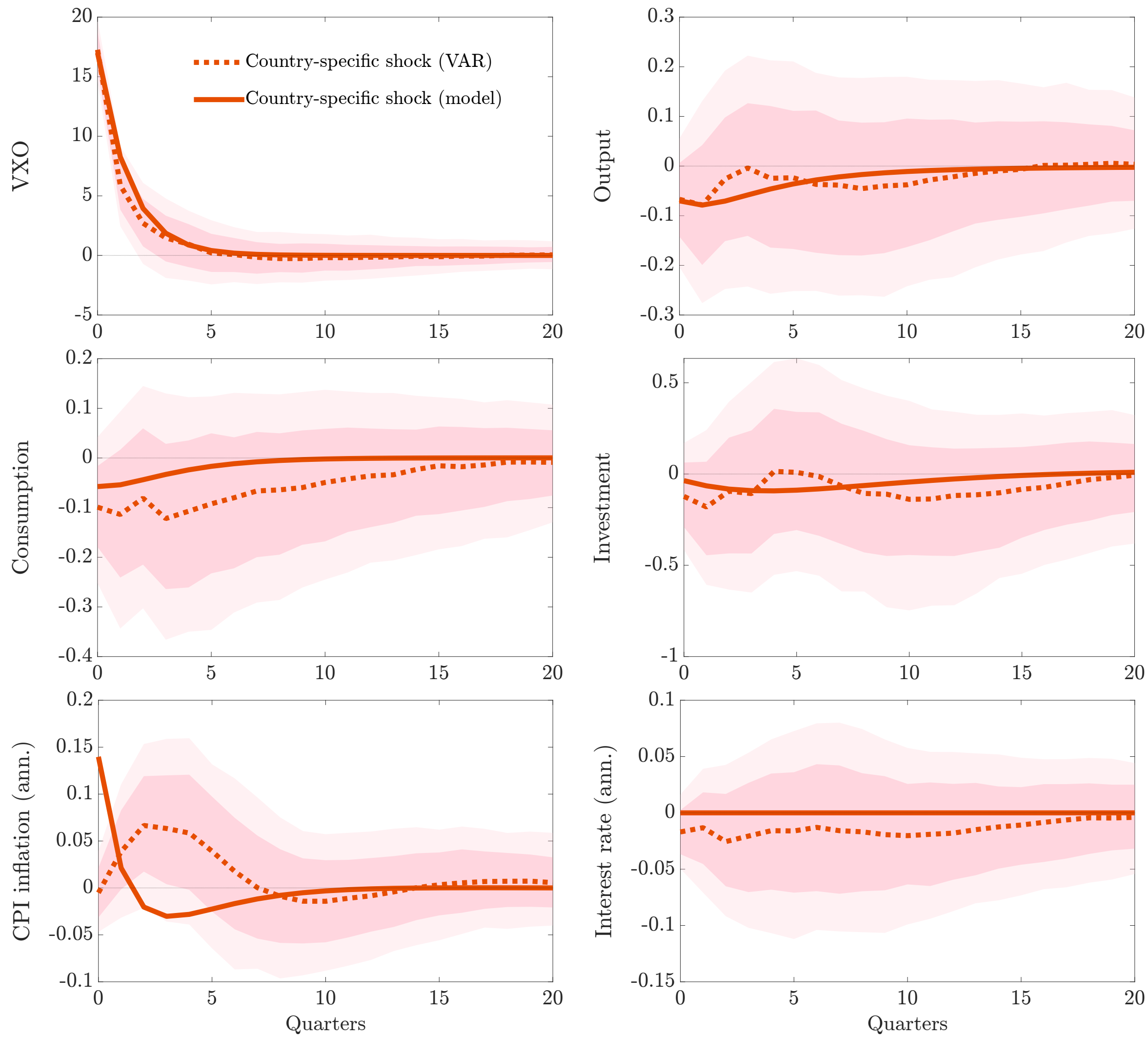
$$VXO = 100\sqrt{4\left[\frac{\mathbb{E}_t\left(M_{t+1}(R^E_{t+1})^2\right)}{\mathbb{E}_t\left(M_{t+1}\right)} - \left(\frac{\mathbb{E}_t\left(M_{t+1}R^E_{t+1}\right)}{\mathbb{E}_t\left(M_{t+1}\right)}\right)^2\right]}\ . \tag{C.29}$$

In contrast, the VXO under the physical measure is given by

$$VXO = 100\sqrt{4Var_t\left(R^E_{t+1}\right)} = 100\sqrt{4\left[\mathbb{E}_t\left[(R^E_{t+1})^2\right] - \left(\mathbb{E}_t\left(R^E_{t+1}\right)\right)^2\right]}\ . \tag{C.30}$$

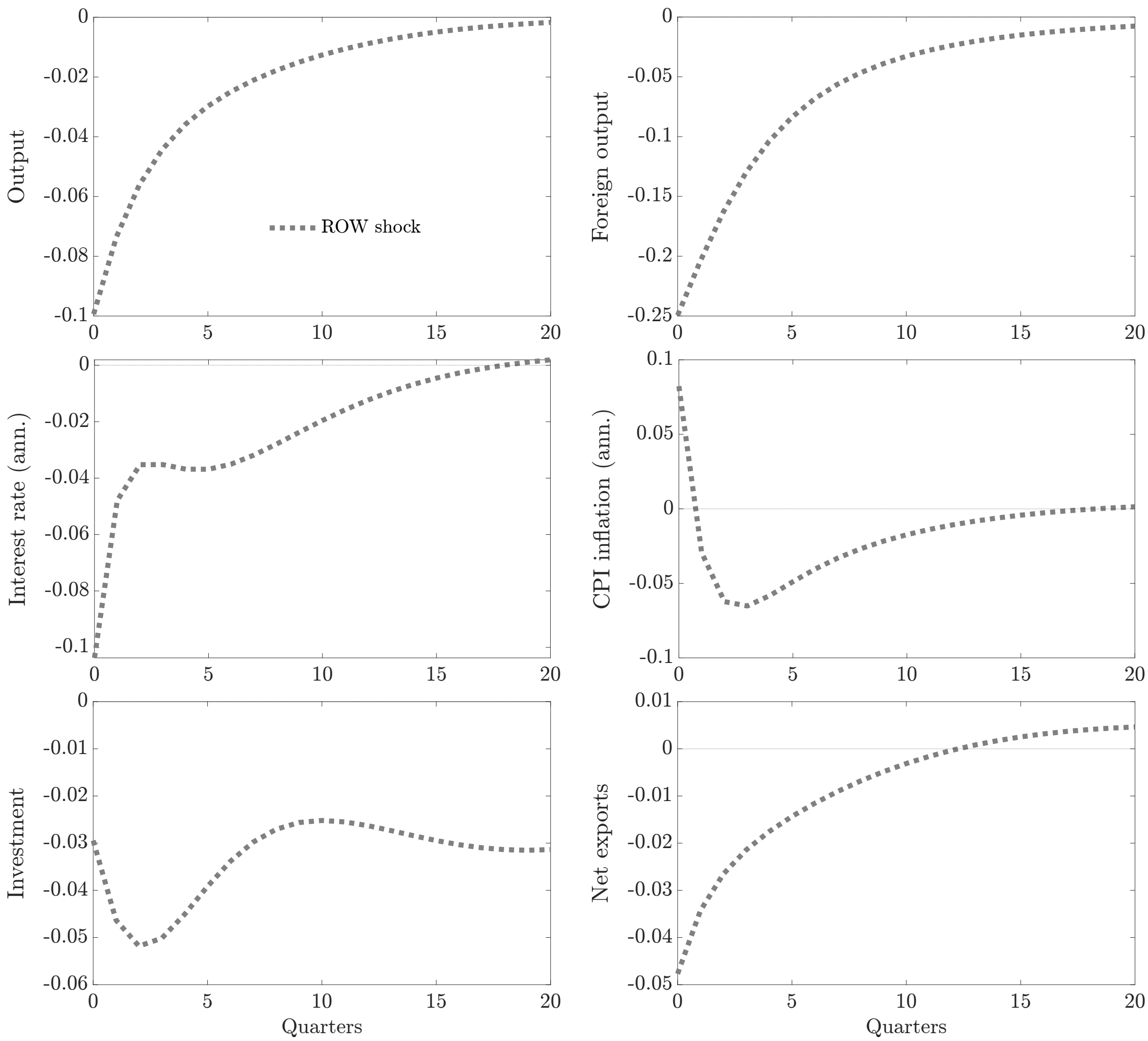
There is no difference between the physical and risk-neutral VXO at third order.

Figure C.1: Model fit—external validation



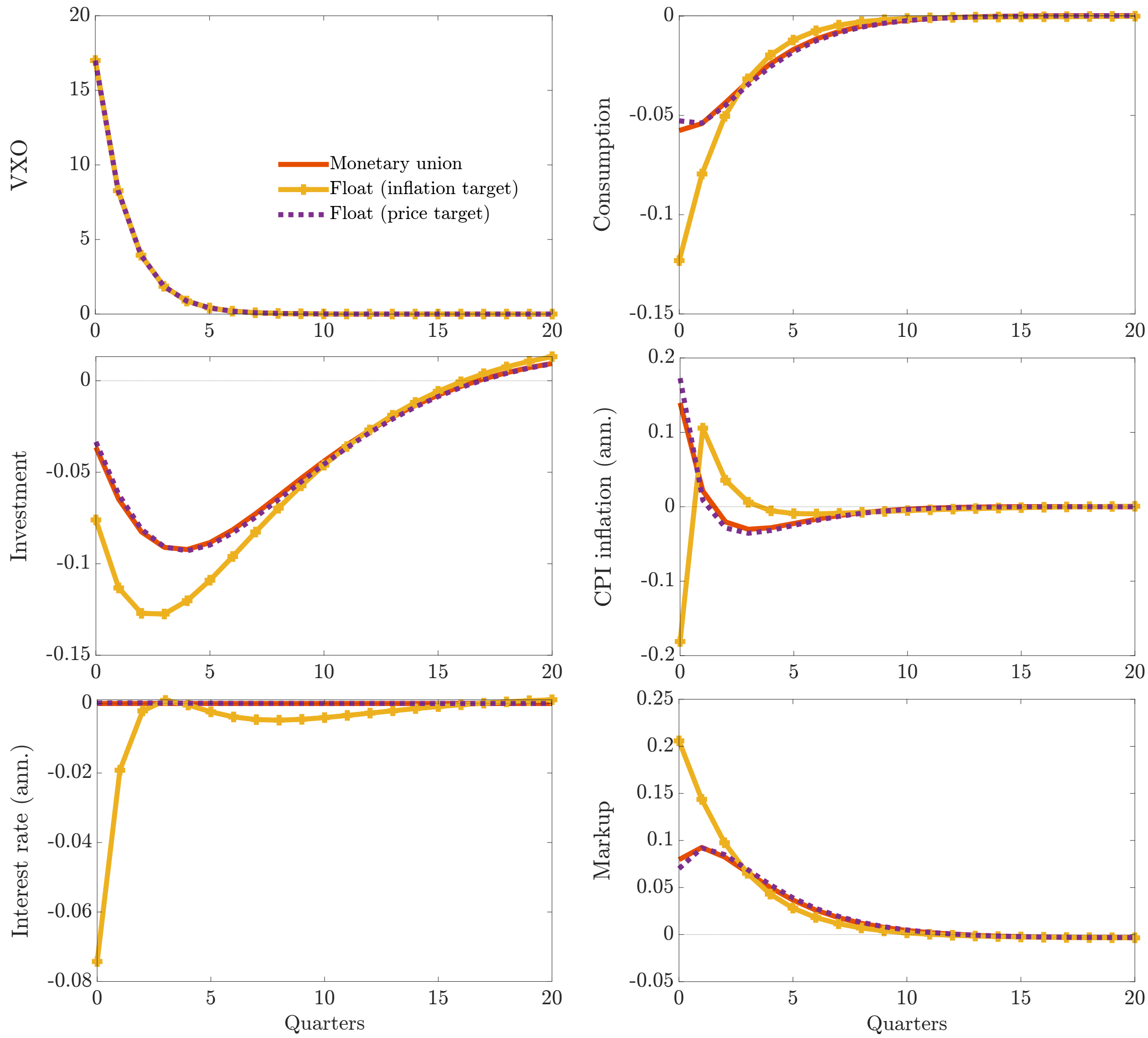
Notes: Model (red solid line) and untargeted VAR (red dotted line) impulse response functions (IRFs) to a one-standard-deviation country-specific preference uncertainty shock in a monetary union. Quarterly responses are in percentage deviations from the stochastic steady state, except for CPI Inflation, which is in ppts. Bands are pointwise 68% (dark) and 90% (light) HPDIs.

Figure C.2: Effects of shock that only hits Foreign



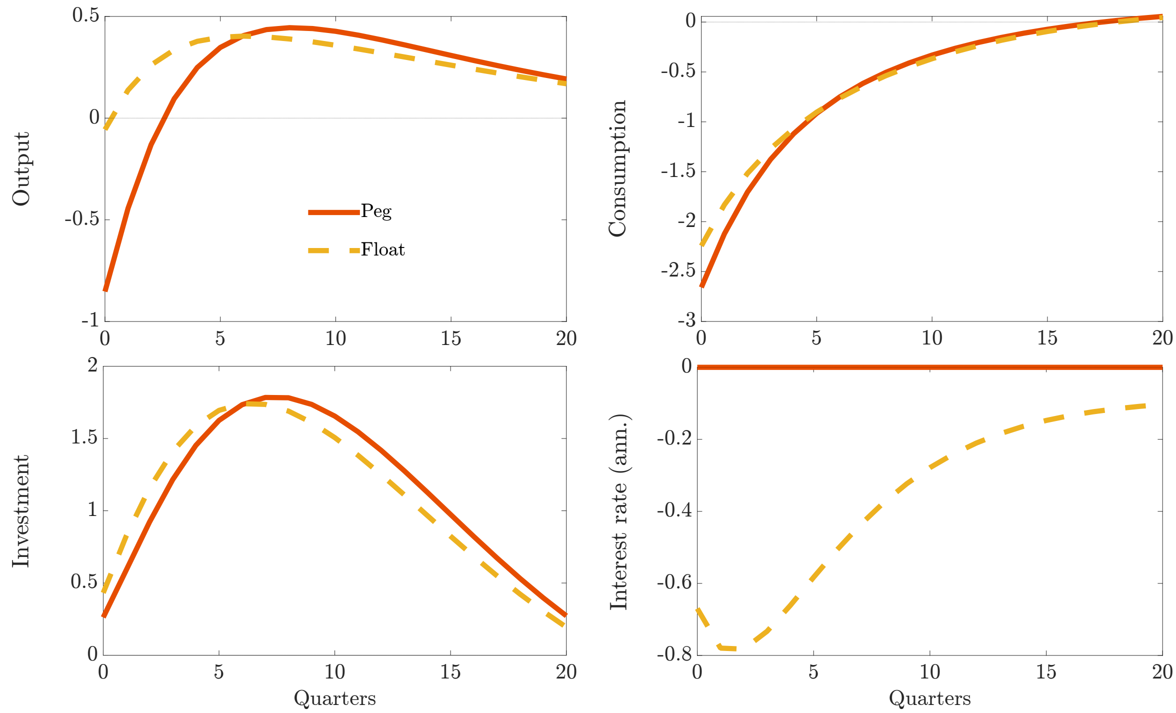
Notes: Model IRFs to a one-standard-deviation Foreign-specific (grey dotted line) preference uncertainty shock in a monetary union. Quarterly responses are in percentage deviations from the stochastic steady state, except for CPI Inflation and the interest rate, which are in ppts, and net exports, which are in percent of output at the stochastic steady state.

Figure C.3: Effects of country-specific uncertainty shocks w/ alternative policies



Model IRFs to a one-standard-deviation country-specific preference uncertainty shock in a monetary union (red solid line), a float with inflation targeting (yellow solid line with plus-shaped markers), and a float with price targeting (purple dotted line). Quarterly responses are in percentage deviations from the stochastic steady state, except for CPI Inflation and the interest rate, which are in ppts.

Figure C.4: Effects of country-specific level shock



Notes: Model IRFs to a one-standard-deviation country-specific (red solid line) and common (yellow dashed line) preference shock in a monetary union. Quarterly responses are in percentage deviations from the stochastic steady state, except for the interest rate, which is in ppts.