Foreign economic policy uncertainty shocks and real activity in the Euro area

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Abstract

This paper estimates a Bayesian VAR model on Euro area data and quantifies the reaction of real activity to economic policy uncertainty shocks that originate abroad. Our findings show that US and Chinese uncertainty explains larger shares of fluctuations than European uncertainty. In an extended set-up, we perform a counterfactual simulation and verify the presence of a foreign economic policy uncertainty spillovers channel that magnifies the real effects of US and Chinese uncertainty shocks. The simulation also documents a non-negligible role played by bilateral trading activities in the transmission mechanism of Chinese shocks. In an application with Dutch data, we highlight that structural domestic factors shape region and country-specific uncertainty in the propagation of foreign economic policy uncertainty shocks onto the economy.

Keywords: Uncertainty shocks; Euro area spillovers; real activity; US; China; Bayesian VAR.

JEL Classification: C32, E30, Q54.

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1 Introduction

Are foreign economic policy uncertainty (EPU) shocks important for the real business cycle? Some recent contributions find that variations in the degree of uncertainty in other countries and regions' policy decisions can explain a non-negligible share of domestic real activity fluctuations (Mumtaz and Theodoridis, 2017; Mumtaz and Musso, 2021). Notably, this conclusion is drawn by assuming uncertainty at regional and global level, i.e. supranational, with respect to the country of interest. Therefore, despite the growing literature on this topic, few empirical applications have identified the domestic real effects of EPU shocks that originate externally,[1] see for example Caggiano et al. (2020a). Furthermore, the evidence is retrieved as based on a single source of foreign EPU shock so far. Our aim is to extend such analysis and elaborate on a comparison of external EPU shocks.

This paper studies the effects of economic policy uncertainty shocks that originate in the US and China on real activity indicators in the Euro area. Our empirical framework relies on a monthly Bayesian VAR (BVAR) model. As proxies for economic policy uncertainty, we use the EPU indices constructed by Baker et al. (2016). We control for either foreign and domestic uncertainty in order to allow (but not require) foreign EPU shocks to affect the domestic economy via uncertainty spillovers. As for real activity indicator, we consider the industrial production index as benchmark. The choice relies on the fact that the industrial production index mainly aggregates information from sectors that are highly tradable. This allows to focus on a measure of real activity that is plausibly connected with external developments. The model contains a number of additional Euro area macro variables, including inflation, the policy rate, the bilateral real exchange rate, and, in an extended set-up, real net exports. The identification strategy assumes exogeneity of macroeconomic uncertainty with respect to changes in the business cycle. Furthermore, we impose no contemporaneous reaction of external EPU shocks to domestic EPU shocks. The reliability of this exogeneity assumption is supported by three factors. First, it is consistent with the analysis of spillover effects. Second, it reduces the distortion of the outcomes by minimizing potential second-round effects. Third, it is in line with similar specifications that Cholesky-order external uncertainty indicators before the other variables (Fontaine

[1]Throughout the paper, the terms 'external' and 'foreign' are used interchangeably.

et al., 2017; Caggiano et al., 2020a). We estimate a BVAR model for each of the foreign shocks considered, i.e. two BVAR models are estimated. Considering two distinguished models allows (i) to avoid biased assumptions on the Cholesky-identified EPU shocks and (ii) to deepen the analysis BY loading specific bilateral indicators that precisely describe the mutual relationship between the Euro area and the other country.

The results point to statistically and economically relevant effects of foreign EPU shocks for the real business cycle in the Euro area. Our findings report a deterioration of industrial production after a foreign EPU shock. The real fluctuations show a one-year persistence and a peak response months after US and Chinese EPU shocks. The values for the one-year ahead forecast error variance decomposition are also comparable, i.e. 15% and 10% after a US and Chinese EPU shock, respectively. We investigate the transmission mechanism of foreign EPU shocks via a counterfactual analysis. It consists of two simulations that detect the response of the real activity indicators in the case of (i) the domestic EPU index is not allowed to react to systematic movements in foreign EPU shocks and (ii) net exports do not react to foreign EPU shocks. This exercise documents the existence of a foreign EPU spillovers channel that magnifies the real effects of US and Chinese EPU shocks. The foreign uncertainty multiplier quantifies a one-year cumulative amplification effect of around 80% and 40% in the case of a jump in the US and Chinese EPU index, respectively. We replicate the analysis using the Netherlands as a representative case study. The application stresses the importance of properly disentangling across the different sources of foreign EPU shocks. The results with Dutch data indeed highlight that US and Chinese EPU shocks can propagate into the domestic economy in multiple ways, e.g. via region or country-specific uncertainty channels.

Our paper enriches the literature that uses time series models to study the economic effects of uncertainty shocks. This strand of the literature has significantly widened since the contribution of Bloom (2009). This work mainly relates to the applications by Colombo (2013), Fontaine et al. (2017), Netšunajev and Glass (2017), Caggiano et al. (2020a), and Nilavongse et al. (2020). Colombo (2013) estimates a two-block VAR model on US and Euro area pre-GFC data and shows that US EPU shocks are more relevant than hikes in domestic uncertainty for industrial production, prices, and the interest rate in the Euro area. Fontaine et al. (2017)

document that the US economy significantly reacts to Chinese EPU shocks, especially during economic busts. Netšunajev and Glass (2017) include uncertainty indices and unemployment rates for the US and the Euro area together with global real activity in a VAR model. They find that US uncertainty shocks are relevant either for the internal and the Euro area labour market. The outcomes also reveal a quicker reaction of the labour market in the US to absorb uncertainty shocks with respect to the Euro area. Caggiano et al. (2020a) investigate the effects of US EPU shocks on the Canadian business cycle. To do so, they estimate a nonlinear VAR model and find that US EPU shocks exert asymmetric effects on the unemployment rate in Canada and the impact is larger during periods of recession. Nilavongse et al. (2020) document that US EPU shocks have a greater impact on the UK real economy, UK EPU shocks are more important for the volatility of the British pound. Other contributions focus on the identification of uncertainty factors in economic fluctuations. Mumtaz and Theodoridis (2017) use a factor model with stochastic volatility to decompose the time-varying variance of macroeconomic and financial variables into contributions from country-specific uncertainty and uncertainty common to 11 OECD countries. They report that the common component drives most of the time-varying volatility of nominal and financial variables. Carriero et al. (2020) exploit a large VAR framework to measure international macroeconomic uncertainty and its effects on selected economies. They provide evidence of significant commonality in macroeconomic volatility, with one common factor driving strong co-movements across economies and variables. Their evidence suggests that unexpected increases in uncertainty reduce output and stock prices, adversely affect labour market conditions, and in some economies lead to an easing of monetary policy Mumtaz and Musso (2021) estimate a dynamic factor model with time-varying parameters and retrieve measures of global, regional, and country-specific uncertainty by decomposing the variance of each variable from a panel of 22 countries. They show that global uncertainty plays a primary role in explaining the volatility of inflation, interest rates, and stock prices in most of the countries considered, despite time variation of the effects is also present. A more 'autharchic' application is developed by Mumtaz and Theodoridis (2018) that work with a VAR and DSGE framework to study the time-varying effects of uncertainty shocks on US real and financial indicators. They find that the transmission mechanism has weakened over time as a result of an increase in the monetary authorities' anti-inflation stance and a 'flattening' of the Phillips curve. A number of other papers monitor the effects of country-related uncertainty indices on the domestic economy. Caggiano et al. (2014, 2017a, 2021b) find that real variables respond asymmetrically to uncertainty shocks across different phases of the business cycle. Aastveit et al. (2017) and Castelnuovo and Pellegrino (2018) address the effectiveness of monetary policy across tranquil and uncertain times in the US and Euro area, respectively. The results point to negligible real effects of monetary shocks during periods of high uncertainty. Pellegrino (2018) replicates the application with Euro area data. In terms of monetary policy regime, Caggiano et al. (2017b) explore the real effects of uncertainty shocks at the zero lower bound (ZLB) state. They document that the contractionary effects of uncertainty shocks are statistically larger when the ZLB is binding. Their findings lend support to theoretical contributions on the interaction between uncertainty shocks and the stance of monetary policy. Few other works replicate the nature and the size of uncertainty shocks connected to specific events and quantify the impact of episode-specific jumps in uncertainty. Caggiano et al. (2020b) identify the Covid-19-induced uncertainty shock and evaluate its impact on global output. Caggiano et al. (2021a) quantify the finance-uncertainty multiplier via a BVAR model identified with sign and ratio restrictions. Moreover, by using the financial narrative associated with the Black Monday (October 1987) and the GFC (September and October 2008), they sharpen the identification of financial uncertainty and credit supply shocks for the US. Similarly, Caggiano et al. (2021b) emphasize the role of nonlinearities when uncertainty-unemployment links are analysed along a sample period that includes the GFC. The literature also applies time series models to focus on the role played by financial indicators in transmission mechanism of uncertainty shocks. Nodari (2014) shows that the widening of the US credit spread after a surge of financial regulation policy uncertainty leads to a deterioration of the real business cycle. Caldara et al. (2016) use a penalty function approach within a VAR framework to examine the interaction between financial conditions and economic uncertainty. Their results depict the Great Recession as an acute manifestation of the toxic interaction between the two types of shocks. In a BVAR model identified with sign and ratio restrictions, Furlanetto et al. (2019) find that the relevance of uncertainty shocks for output fluctuations is limited when other financial factors, e.g. credit and housing, are taken into account.

The rest of the paper is organized as follows. Section 2 describes the econometric framework, the data used, and the identification strategy. Section 3 presents the baseline results. Section 4 discusses the results from a counterfactual exercise based on an extended set-up. Section 5 performs a country-based analysis using Dutch data. Section 6 concludes.

2 The empirical framework

On the construction of the EPU indices. As stressed in the Introduction, we control for EPU by using the EPU indices developed by Baker et al. (2016). The indices are based on newspaper coverage frequency. We here describe largerly the procedure to construct the EPU index for the US. The constructions of the other EPU indices included follows the same approach. For the US, Baker et al. (2016) use two overlapping sets of newspapers. The first spans the 1900-85 period and comprises The Wall Street Journal, The New York Times, The Washington Post, The Chicago Tribune, The Los Angeles Times, and The Boston Globe. Since 1985, USA Today, The Miami Herald, The Dallas Morning Tribune, and The San Francisco Chronicle have been added to the set. The authors perform within-month searches of all articles, starting in January 1900, for terms related to economic and policy uncertainty. In particular, they search for articles containing the term 'uncertainty' or 'uncertain', the terms 'economic', 'economy', 'business', 'commerce', 'industry', and 'industrial', and the terms: 'congress', 'legislation', 'white house', 'regulation', 'federal reserve, deficit', 'tariff', or 'war'. The article is included in the count if it features terms in all three categories pertaining to uncertainty, the economy and policy. To deal with changing volumes of news articles for a given newspaper over time, (Baker et al., 2016) divide the raw counts of policy uncertainty articles by the total number of news articles containing terms regarding the economy or business. They then normalize each newspaper's series to unit standard deviation prior to December 2009 and then sum up all series. The same procedure is applied to construct the EPU index for China and Europe. The differences are in the choice of words and newspapers, since for either the Chinese and European EPU index, the authors select country/region specific words and newspapers. Additional details can be found on the 'Economic Policy Uncertainty' website. It is worth specifying that, in line with Fontaine et al. (2017), our choice for the Chinese EPU index falls on the one constrcuted by using info from the South China Morning Post, which is the leading English newspaper in Hong Kong. In fact, as indicated by Baker et al. (2016), the construction of the EPU index for China represents a special challenge given censorship and government pressure to the media. Baker et al. (2016) minimize this issue by relying on Hong Kong English-language newspaper.[2]

The BVAR model. The econometric set-up relies on a monthly VAR model estimated with a Bayesian technique. The reduced-form of the model reads as follows:

$$\mathbf{X}_t = \mathbf{c} + \sum_{i=1}^{P} \mathbf{B}_i \mathbf{X}_{t-i} + \varepsilon_t \tag{1}$$

where $\mathbf{X}_t$ represents a $(N \times 1)$ vector of N endogenous variables, $\mathbf{c}$ is a $(N \times 1)$ vector of constants, $\mathbf{B}_i$ are $(N \times N)$ parameter matrices, where $i = 1, ..., P$, and P represents the number of lags. In line with other empirical contributions that use monthly VAR models (Nodari, 2014; Caggiano et al., 2021a), in our baseline specification we set $P = 6$. This lag order allows to control for serial correlation of the residuals. Our findings are robust to alternative lag orders. The vector ε_t contains $(N \times 1)$ i.i.d. error terms where $\varepsilon_t \sim N(0, \Sigma)$. We estimate the model in (1) via a Bayesian technique as in Arias et al. (2018). For the reduced-form parameters in (1), we assume a conjugate prior distribution, i.e. a normal-inverse-Wishart distribution. This leads to a posterior normal-inverse-Wishart distribution over the reduced-form parameters.

Modelled vector and identification. The endogenous variables enter the vector $\mathbf{X}_t$ in the following order: $\mathbf{X}_t = [EPU_t^{US(China)}, EPU_t, \Delta y_t, \pi_t, R_t, \Delta\epsilon_t]'$, where $EPU_t^{US(China)}$ is the US (China) EPU index and EPU_t is the European EPU index. The rest of the variables are indicators for the Euro area economy. In particular, Δy_t is the year-on-year change of the industrial production index, π_t is the annual inflation rate, R_t is the policy rate, and $\Delta\epsilon_t$ is the year-on-year change of the bilateral real exchange rate.[3] The choice of observables corresponds with the selected variables in Caggiano et al. (2020a). The first observation is 2000M1 in the model with the US EPU index. This choice is dictated by the fact that we construct our

[2] Baker et al. (2016) face a comparable challenge when it comes to the construction of the Russian EPU index. Because of that, they rely on *Kommersant*, a Russian newspaper that focuses on financial topics and is proven fairly free from pressures exercised by government bodies.

[3] The change in bilateral real exchange rate is computed as the annual growth of change of the nominal excahnge rate adjusted by the inflation rates in the two countries.

dataset starting from the birth of the Euro area, i.e. from 1999M1. Some variables are then transformed to annual growth rates shifting the first observation in the model to 2000M1. For the model with the Chinese EPU index, the first observation corresponds to 2005M1, given the availability of the data on the Euro/Yuan exchange rate. In both specifications, the final observation is 2019M12. We perform a recursive identification of shocks.[4] The ordering of the variables that entails positioning the uncertainty indicator(s) first is standard in the literature on uncertainty shocks (Nodari, 2014; Caggiano et al., 2014; Baker et al., 2016; Leduc and Liu, 2016; Caggiano et al., 2017a). We borrow the exogeneity assumption for the foreign-domestic EPU indices from Fontaine et al. (2017) and Caggiano et al. (2020a).

3 The impact of foreign EPU shocks

We present here the impulse-response functions (IRFs) that describe the dynamic reaction of the Euro area variables to US and Chinese EPU shocks.

Baseline results. Figure 1 contains the IRFs to one-standard deviation US and Chinese EPU shock of the European EPU index and the Euro area economic indicators. We first observe a similar upward movement of the European EPU index following an exogenous jump in US and Chinese EPU. A comparable reaction is also detected with regard to the IPI growth. In particular, it temporarily contracts, with a median peak response of -0.6% and -0.4% six months after a US and Chinese EPU shock, respectively, before gradually going back to trend after about one year (considering the associated 68% credible set). This evidence of spillover effects are in line with Caggiano et al. (2020a) and their application on the impact of US EPU shocks on the Canadian business cycle. The pattern of the real activity indicator also matches with the one documented by Bloom (2009). Bloom (2009) highlights the short to medium term peak-rebound trajectory of industrial production. Inflation and, to a greater extent, the policy rate respond to foreign EPU shocks. Nonetheless, the effects on inflation is poorly significant, especially after US EPU shocks, and short-lived (see Castelnuovo, 2022 for an overview of the inflationary effects of uncertainty shocks), while the reaction of policy denotes a substantial reaction of monetary policy to counteract uncertainty-driven economic downturns. The depreciation of

[4] The Appendix contains details on the structural representation of the model.

Euro vis-a-vis US dollar testifies an increase of perceived global risk that entails a flight to safety and dollar value growth.

Sensitivity checks. We verify the validity of our baseline results by performing a number of robustness checks. The results from these exercises are reported in terms of median IRFs in Figure 2.

The first check concerns the specification and identification strategy. With regard to the former, we estimate a single model that includes either the US EU index and the Chinese EPU index. With regard to the latter, the enlarged model requires a different staregy to identify the foreign EPU shocks. Short-run restrictions might indeed be misleading and generate biased resuls. To overcome the problem, we identify foreign and domestic EPU shocks via a combination of sign, ratio, and narrative restrictions as in Caggiano et al. (2021a). The estimation of the model follows the approach by Antolín-Díaz and Rubio-Ramírez (2018). A specific description of the identification approach is reported in the Appendix. Moreover, we modify the model with the respective to the baseline specification by substituting the bilateral real exchange rate with the Euro real effective exchange rate. Because of the availability of the Euro real effective exchange rate observations, we estimate the model over the sample period 2000M1-2019M12. The second test involves the Cholesky ordering. By loading the uncertainty proxies before the macroeconomic variable, we impose no contemporaneous effects of business cycle shocks on the EPU indices. We check this exogeneity assumption by moving the uncertainty block after the bilateral real exchange rate and re-estimating the model. The last two robustness checks entails the variation of the baseline lag order. In particular, we (i) increase to nine and (ii) reduce to three the lag order and re-estimate our BVAR. The outcomes from these tests corroborate the robustness of our baseline results.

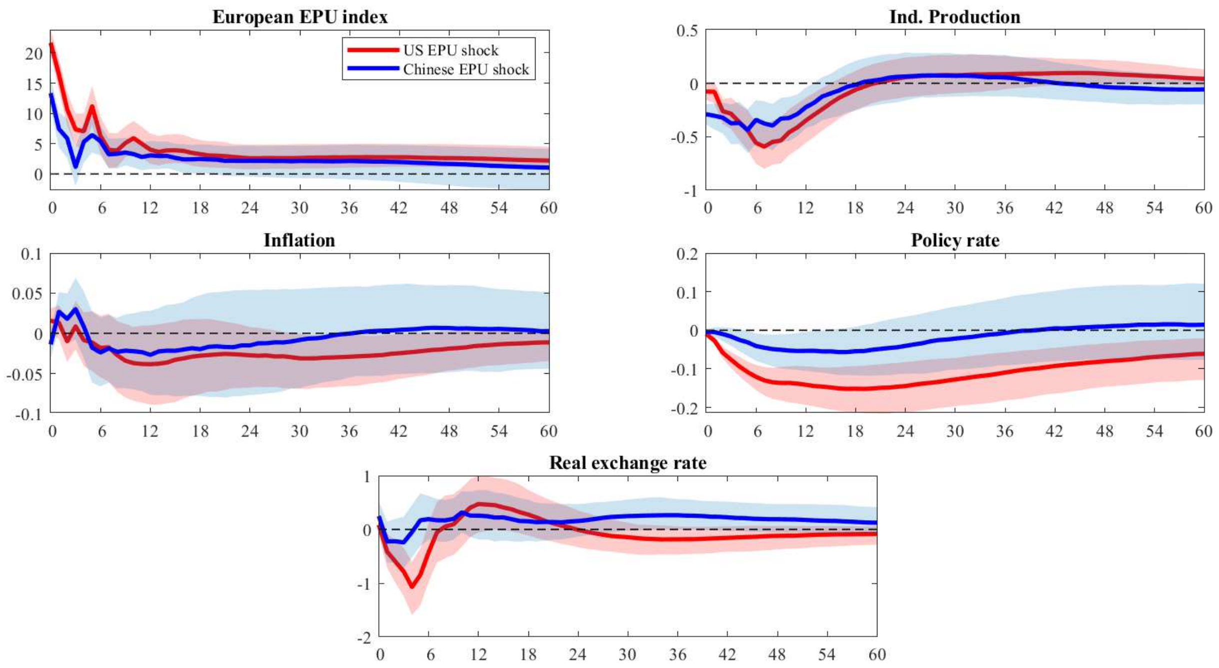

Figure 1: Impulse-response functions of the European EPU index and Euro area macroeconomic variables to a one-standard deviation foreign EPU shock.

Notes: Red plots represent IRFs to a US EPU shock. Blue plots represent IRFs to a Chinese EPU shock. The solid line represents the posterior median at each horizon and the shaded area indicates the 68^{th} posterior probability region of the estimated impulse responses.

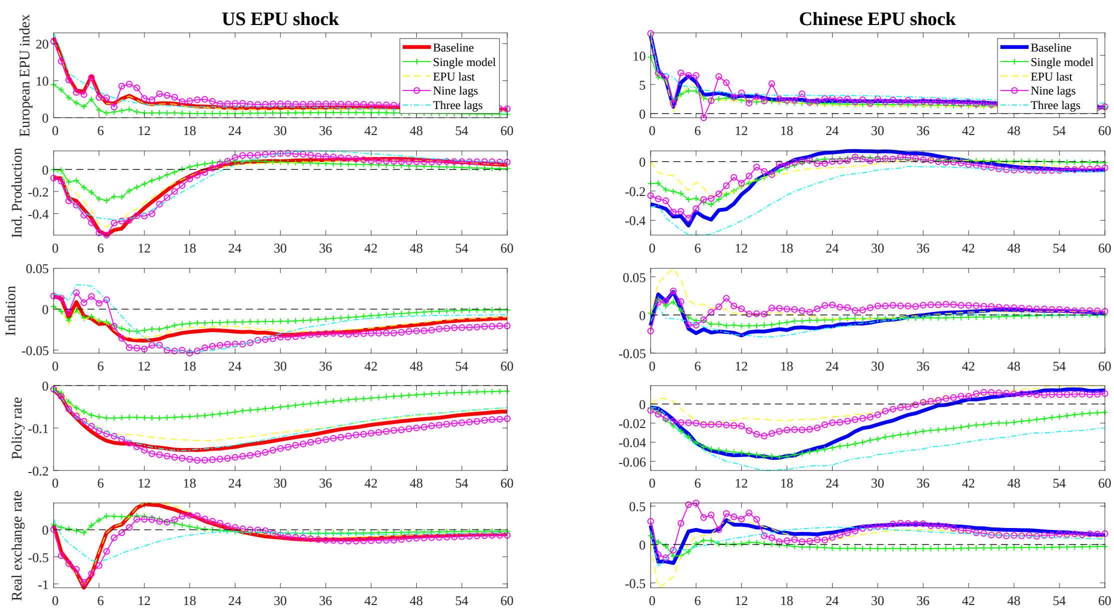

Figure 2: Rbostness. Median impulse-response functions of the European EPU index and Euro area macroeconomic variables to a one-standard deviation foreign EPU shock.

4 The contribution and transmission channel of foreign EPU shocks

FEVD. What is the role played by the EPU shocks in explaining Euro area indicators fluctuations? Table 1 compares the 12-month ahead FEVD for each variable of interest to different EPU shocks. The two foreign EPU shocks explain a significant share of the volatility of the European EPU index, despite the importance of the EPU shocks originating in the US is greater than Chinese EPU shocks. The readings for IPI growth suggest a comparable contribution of EPU shocks from the US and China (15% vs 10%). It is also worth noting that the contribution of domestic EPU shocks is lower with respect to foreign EPU in both specifications. The US EPU shocks also contribute substantially to policy rate fluctuations (23%) while Chinese EPU shocks account only marginally (3%). The role played by foreign EPU shocks in explaining inflation and exchange rate disturbances is negligible, as testified by values equal to 2%-3% for inflation and 5%-3% for the real exchange rate in the model with the US and Chinese EPU index, respectively.

Through which channel foreign EPU shocks transmit onto the Euro area real economy? To address this question, we check two potential alternatives. First, it might be that uncertainty shocks abroad can affect domestic uncertainty and propagate into the economy via an international EPU spillovers channel. Second, given the trading relationships between the Euro area and the other two countries, we could also conclude about the existence of a trade channel that works via the uncertainty-driven contraction of US and Chinese demand for imports and depresses the Euro area economy as a whole.

Variable	$\tilde{\varepsilon}_t^{USEPU}$	$\tilde{\varepsilon}_t^{ChinaEPU}$	$\tilde{\varepsilon}_t^{EPU(US)}$	$\tilde{\varepsilon}_t^{EPU(China)}$
EPU_t	0.41 [0.10 0.23]	0.16 [0.32 0.50]	0.45 [0.37 0.54]	0.62 [0.54 0.71]
Δy_t	0.15 [0.07 0.25]	0.10 [0.04 0.20]	0.07 [0.05 0.15]	0.09 [0.03 0.19]
π_t	0.02 [0.01 0.07]	0.03 [0.01 0.06]	0.15 [0.05 0.42]	0.02 [0.01 0.06]
R_t	0.25 [0.13 0.34]	0.03 [0.01 0.09]	0.07 [0.02 0.14]	0.18 [0.09 0.30]
$\Delta\epsilon_t$	0.05 [0.03 0.10]	0.03 [0.01 0.06]	0.02 [0.01 0.06]	0.05 [0.02 0.10]

Table 1: One-year ahead median forecast error variance decomposition.

Notes: The individual contributions are reported as fractions of the total forecast error variance decomposition. Brackets contain the 68$^{\text{th}}$ posterior probability region of the forecast error variance decomposition distribution. The second and third column contains the results related to US and Chinese EPU shocks while the fourth and fifth column contain the results related to European EPU shocks.

Counterfactual simulation. We empirically assess the transmission mechanism of foreign EPU shocks via the two channels by performing a counterfactual simulation following the approach by Kilian and Lewis (2011). In particular, we first augment the baseline vector with bilateral net exports to control for trade-related activities.

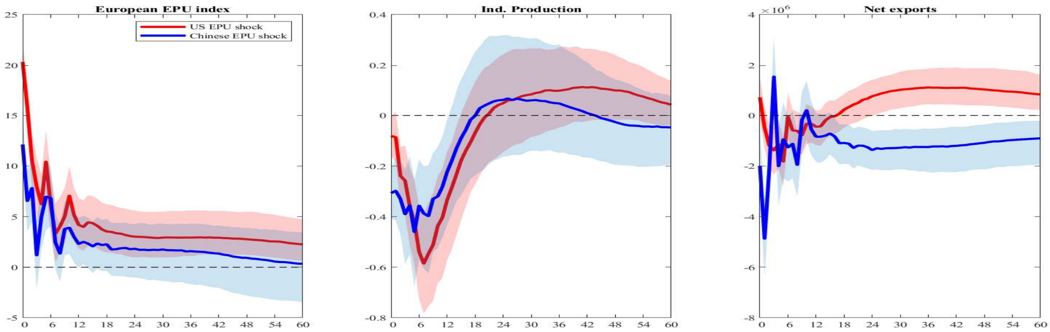

Figure 3: Impulse-response functions of the European EPU index and Euro area industrial production and bilateral net exports to a one-standard deviation foreign EPU shock.

Notes: Red plots represent impulse-response functions to a US EPU shock. Blue plots represent IRFs to a Chinese EPU shock. The solid line represents the posterior median at each horizon and the shaded area indicates the 68^{th} posterior probability region of the estimated impulse responses.

After that, we run the new specified model twice and simulate the response of our endogenous variables to a foreign EPU shock while keeping the variable of interest at its pre-shock level. The first exercise generates fictitious shocks to European EPU that perfectly offset the impact of the foreign EPU shock on the European EPU index as well as the dynamic feedback effects going from the other endogenous variables of the system to European EPU . We replicate the exercise by generating fictitious shocks to net exports. Figure 4 reports the results in terms of difference IRFs of IPI growth computed in the baseline and counterfactual scenario. The results point to statistically different effects for IPI growth in the case of a European uncertainty reaction to a US EPU shock. This implies that when US EPU uncertainty spikes, the response of domestic uncertainty plays a role in the transmission of US EPU shocks to the real economy. In particular, we detect an amplification of the real effects of US EPU shocks due to US-European uncertainty spillovers. The role of domestic uncertainty is also important when a Chinese EPU shock takes place. Despite to a lower extent, movements of the European EPU index testify the presence of a Chinese-European uncertainty spillover mechanism that augments the deterioration of IPI growth. Turning to the exercise in which net exports are not allowed to react to foreign EPU shocks, we can observe different outcomes. The difference IRFs in the case of US EPU shocks are not statistically significant, while in the case of a Chinese EPU shock a small short-run reaction is retrieved.

Foreign uncertainty multiplier. We complement the counterfactual simulation analysis with the computation of the foregin uncertainty multiplier at a one-year horizon. To do this, we follow the procedure by Caggiano et al. (2021a) and compute the cumulative multiplier as follows:

$$\frac{\sum_{h=0}^{12}\left[\frac{\partial \Delta y_{t+h}}{\partial \tilde{\varepsilon}_t^{USEPU}}\right]}{\sum_{h=0}^{12}\left[\frac{\partial \Delta y_{t+h}}{\partial \tilde{\varepsilon}_t^{USEPU}}\middle|\frac{\partial EPU_{t+h}^{EU}}{\partial \tilde{\varepsilon}_t^{USEPU}}=0\right]} \tag{2}$$

In (2), the numerator is the sum of the values of the impulse response of industrial production to a foreign uncertainty shock in the baseline scenario, and the denominator is the sum of such values when the response of the European EPU index is kept fixed. This way of computing multipliers naturally accounts for the overall real impact of an uncertainty shock in the two scenarios. However, for realizations of real activity that cross the zero line over the considered horizon, this way of computing the multiplier may generate an excessive dispersion of realizations of the cumulative multiplier (when the sum of the values of the counterfactual response of industrial production over the considered horizon gets close to zero). Hence, although our benchmark is the cumulative multiplier, we also compute the peak multiplier to check the solidity of the results. the peak multiplier is computed via the following formula:

$$\frac{\min_{\forall h\in[0,12]}\left[\frac{\partial \Delta y_{t+h}}{\partial \tilde{\varepsilon}_t^{USEPU}}\right]}{\min_{\forall h\in[0,12]}\left[\frac{\partial \Delta y_{t+h}}{\partial \tilde{\varepsilon}_t^{USEPU}}\middle|\frac{\partial EPU_{t+h}^{EU}}{\partial \tilde{\varepsilon}_t^{USEPU}}=0\right]} \tag{3}$$

where the numerator in (3) is the minimum value (i.e., the largest negative realization) of industrial production to a foreign EPU shock when the response of the European EPU index is unconstrained (baseline scenario), the denominator is the minimum value of such response when the response of the European EPU index is kept fixed The median readings, together with the 16^{th} and 84^{th} percentiles of the multipliers distributions, are reported in Table 2. The values for the cumulative multiplier state that the domestic uncertainty channel magnifies the real effects of US and Chinese EPU shocks by more than 80% and 40%, respectively. With regard to the trade channel, we find no role after a jump in the US EPU index, while non-negligible amplifications are found after changes in the Chinese EPU index. The results for the peak multiplier are qualitatively comparable.

Employment as real activity indicator. We substitute our baseline indicator of real activity with employment. We repeat the analysis and report the multipliers as computed in the previous subsection. The new outcomes in Table 3 corroborate the presence of a foreign EPU spillovers channel, while the magnitude of the multipliers referred to the trade channel are close to one either in the case of a US and Chinese EPU shock. This might relate to the presence of labour market frictions that can affect employment and do not allow the labour market to fluctuate substantially (see Leduc and Liu, 2016; Cacciatore and Ravenna, 2021; Caggiano et al., 2021a).

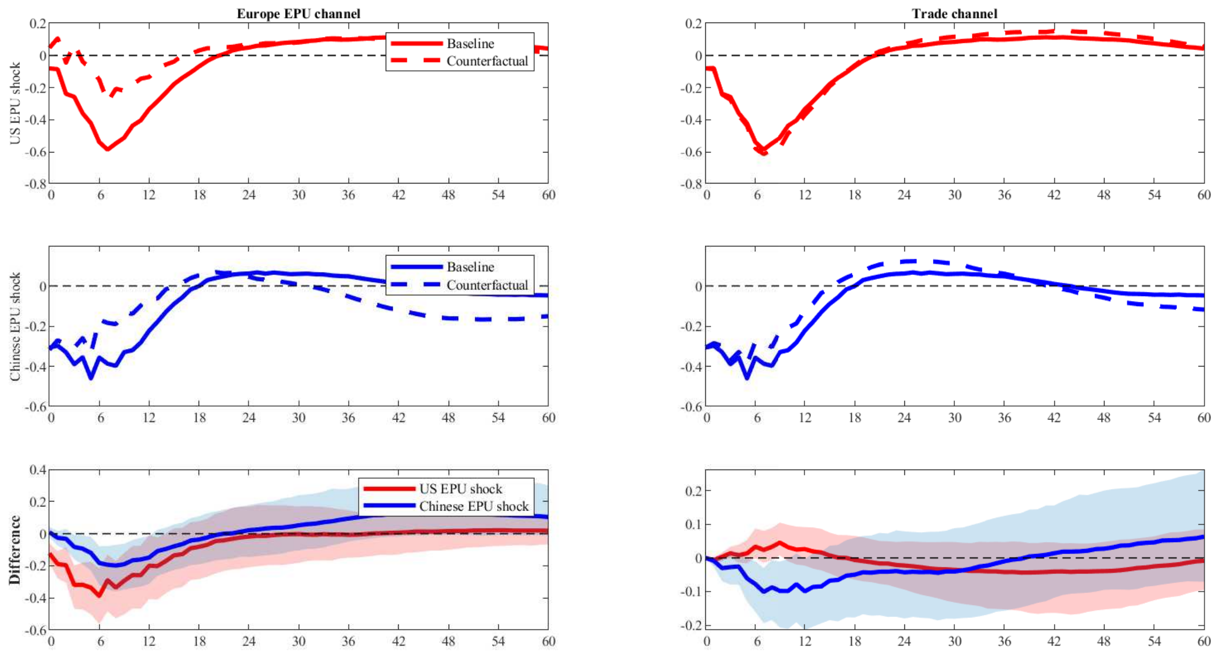

Figure 4: Impulse-response functions of the Euro area IPI growth to a one-standard deviation foreign EPU shock. The last row contains the difference between IRFs in the baseline and counterfactual scenario.

Notes: Red plots represent impulse-response functions to a US EPU shock. Blue plots represent IRFs to a Chinese EPU shock. The solid line represents the posterior median at each horizon and the shaded area indicates the 68^{th} posterior probability region of the estimated impulse responses.

Model	*EPU channel*	*Trade channel*
US EPU	$1.87\ (1.75)$ $[-3.03\ (1.16)\quad 4.76\ (3.58)]$	$0.95\ (0.96)$ $[0.86\ (0.87)\quad 1.00\ (1.02)]$
Chinese EPU	$1.46\ (1.18)$ $[1.06\ (0.98)\quad 2.41\ (1.50)]$	$1.20\ (1.08)$ $[1.01\ (1.00)\quad 1.58\ (1.27)]$

Table 2: Cumulative (peak) foreign uncertainty multiplier of industrial production at a one-year horizon.

Notes: Brackets contain the 68% credible interval of the multiplier distribution.

Model	*EPU channel*	*Trade channel*
US EPU	$1.22\ (1.23)$ $[0.79\ (0.86)\quad 2.19\ (1.89)]$	$0.94\ (0.96)$ $[0.84\ (0.87)\quad 1.00\ (1.01)]$
Chinese EPU	$1.28\ (1.20)$ $[0.69\ (0.88)\quad 2.35\ (1.94)]$	$0.95\ (0.95)$ $[0.78\ (0.77)\quad 1.10\ (1.09)]$

Table 3: Cumulative (peak) foreign uncertainty multiplier of employment at a one-year horizon.

Notes: Brackets contain the 68% credible interval of the multiplier distribution.

5 Foreign EPU shocks: evidence from a country-level analysis

Mumtaz and Musso (2021) highlight the different profiles of the transmission mechanism of uncertainty shocks from a geographical point of view. The propagation of uncertainty shocks onto domestic business cycles can vary whether the shock originates at country, regional, or world level. We extend further our analysis and apply our framework to Dutch data. The choice of the Netherlands relies on three factors. First, its economy represents a key portion for the total Euro area business cycle. It is worth mentioning that the Dutch GDP is the largest among the non-big 4 Euro area countries. Second, the Dutch EPU index is not a building elements of the European-related uncertainty and constitutes a clean measure of national uncertainty. The EPU indices related to the big-4 member states are subcomponents of the European EPU index, providing biased information of the nature of national uncertainty. This feature is a fundamental reason for the choice of the Netherlands since we modify the model specification to control, among the other factors, for country-specific uncertainty.[5] Third, Netherlands's trade openness is substantially large with compared to the EU average.[6] This profile can quantify the role of country characteristics in the transmission process of foreign uncertainty shocks, e.g. via trade. Our new specification for the Netherlands replicates the one from the models with real next exports while augmenting the vector of observable with the Dutch EPU index. The Dutch EPU index is placed between the European EPU index and the (Dutch) industrial production growth. We run an extended version of the counterfactual analysis run with the Euro area data. Figure 5 reports the estimated IRFs from the application with Dutch data.

[5]The Dutch EPU index is available from March 2003. Because of this, the model that investigates the effects of US EPU shocks on the domestic business cycle is estimated on the sample March 2003 - December 2019.

[6]Over the sample period 2000-2019, the average trade openness is 156% for the Netherlands and 81% for the overall EU according to https://ourworldindata.org/.

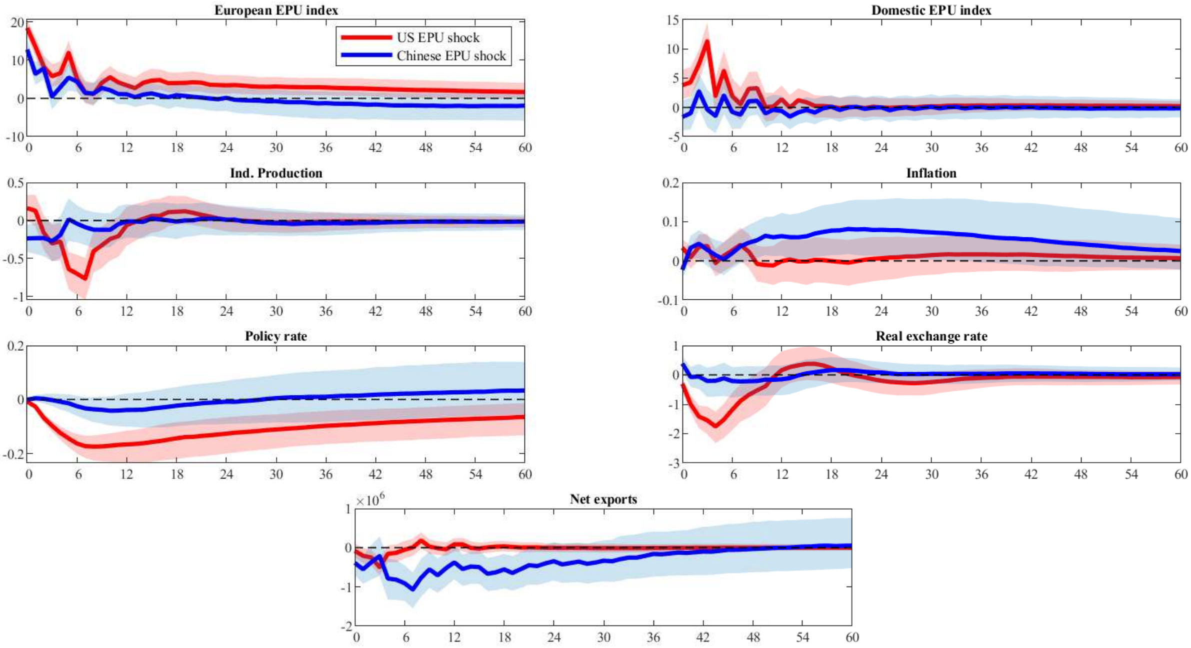

Figure 5: Impulse-response functions of the Europe and Dutch EPU index and Dutch macroeconomic variables to a one-standard deviation foreign EPU shock.

Notes: Red plots represent IRFs to a US EPU shock. Blue plots represent IRFs to a Chinese EPU shock. The solid line represents the posterior median at each horizon and the shaded area indicates the 68^{th} posterior probability region of the estimated impulse responses.

Specifically, next to the European EPU and trade channel, we also control for the domestic EPU channel. To sum up, we perform three counterfactual analysis by not allowing the European EPU index, Dutch EPU index, and net exports to react to foreign EPU shocks. Table 4 contains the results for the multipliers for this set of applications. The outcomes suggest a significant importance of the trade channel for the transmission mechanism of external EPU shocks in the Netherlands. This is consistent with the high share of trading activities over total GDP (openness) in the Netherlands. Such degree of openness might be responsible in the identification of the trade channel as a significant player in the transmitting foreign EPU shocks to the domestic real business cycle.

Model	European EPU channel	Domestic EPU channel	Trade channel
US EPU	0.92 (0.93) [0.54 (0.73) 1.55 (1.23)]	1.18 (1.13) [0.81 (0.93) 1.89 (1.43)]	1.11 (1.05) [0.99 (0.97) 1.42 (1.17)]
Chinese EPU	1.04 (0.94) [0.18 (0.75) 1.93 (1.23)]	1.00 (1.00) [0.64 (0.89) 1.33 (1.13)]	1.09 (1.02) [0.35 (0.98) 1.80 (1.18)]

Table 4: Cumulative (peak) foreign uncertainty multiplier of industrial production at a one-year horizon for the Netherlands.

Notes: Brackets contain the 68% credible interval of the multiplier distribution.

6 Conclusions

This paper estimates a Bayesian VAR on Euro area data and explores the contribution of EPU shocks originated in the US ad China to real business cycle fluctuations. Our results show that foreign EPU shocks are more important than European EPU shocks in terms of the total contraction of real activity. We investigate the propagation mechanism of foreign EPU shocks onto the Euro area economy and testify the existence of a domestic EPU channel through which external EPU shocks amplify. We also document that Chinese EPU shocks propagate via a trade channel onto the real business cycle in the Euro area. We replicate the analysis at a country-level by using the Netherlands as case study. We find that, given the high degree of trade openness, the trade channel plays a key role, especially in the propagation of Chinese EPU shocks.

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Appendix of the paper ”Foreign economic policy uncertainty shocks and real activity in the Euro area” by Filippo Arigoni and Črt Lenarčič

A1 From the reduced-form to the structural BVAR

This Section reports the derivation of the structural model of the econometric approach by closely following Arias et al. (2018).

We rewrite the model in (1) in its structural form as follows

$$\mathbf{X}_t'\mathbf{A}_0 = \alpha + \sum_{k=1}^{P} \mathbf{X}_{t-k}'\mathbf{A}_k + \upsilon_t' \tag{A1}$$

where $\mathbf{X}_t$ is an $(N \times 1)$ vector of variables, υ_t is an $(N \times 1)$ vector of structural shocks, $\mathbf{A}_k$ is an $(N \times N)$ matrix of parameters for $0 \leq k \leq P$ with $\mathbf{A}_0$ invertible, α is a $(1 \times N)$ vector of parameters, P is the lag length, and $1 \leq t \leq T$, with T be the sample size. The vector υ_t, conditional on past information and the initial conditions $\mathbf{X}_0, \ldots, \mathbf{X}_{1-P}$ is Gaussian with mean zero and covariance matrix I_N, the $(N \times N)$ identity matrix. The model described in (A1) can be rewritten as

$$\mathbf{X}_t'\mathbf{A}_0 = \mathbf{Z}_t'\mathbf{A}_+ + \upsilon_t' \tag{A2}$$

where $\mathbf{A}_+' = [\mathbf{A}_1' \cdots \mathbf{A}_P' \, \alpha']$ and $\mathbf{Z}_t' = [\mathbf{X}_{t-1}', \ldots, \mathbf{X}_{t-P}', 1]$. In (A2), $\mathbf{A}_+$ has dimension $M \times N$, while $\mathbf{Z}_t$ has dimension $M \times 1$, with $M = NP + 1$. If we rewrite the reduced-form model in (1) as $\mathbf{X}_t' = \mathbf{Z}_t'\mathbf{B} + \varepsilon_t'$, their elemens are such that $\mathbf{B} = \mathbf{A}_+\mathbf{A}_0^{-1}$, $\varepsilon_t' = \upsilon_t'\mathbf{A}_0^{-1}$, and $\mathbb{E}[\varepsilon_t\varepsilon_t'] = \mathbf{\Sigma} = (\mathbf{A}_0\mathbf{A}_0')^{-1}$. The elements $\mathbf{B}$ and $\mathbf{\Sigma}$ are the reduced-form parameters, whereas $\mathbf{A}_0$ and $\mathbf{A}_+$ are the structural ones. At the same time, $\varepsilon_t^{'}$ are the reduced-form innovations and $\upsilon_t^{'}$ are the structural shocks. Further details on the estimation strategy can be found in Arias et al. (2018).

A2 Sign, ratio, and narrative-restricted model

We document here the details of the robustness check that entails the identification of shocks via a combination of sign, ratio and narrative restrictions. The procedure follows the one implemented by (Caggiano et al., 2021) that enrich the sign and narrative restrictions approach with ratios to constrain IRFs as in Furlanetto et al. (2019). Table A2 contains the sign and ratio restrictions to identify key structural shocks for our analysis.

Variable	US EPU shock	Chinese EPU shock	EU EPU shock
US EPU index	+	+	+
Chinese EPU index	+	+	+
EU EPU index	+	+	+
US EPU index/Chinese EPU index	+	−	
US EPU index/EU EPU index	+		−
Chinese EPU index/EU EPU index		+	−

Table A2: Sign and ratio restrictions imposed on the response of variables (rows) to shocks (columns). The restrictions are imposed on impact. No sign denotes that the response of the variable is left unrestricted.

We complement the identification by imposing two narrative restrictions (Antolín-Díaz and Rubio-Ramírez, 2018), one related to US EPU shocks, the other linked to Chinese

EPU shocks. The former entails to assume an US EPU shock as the main contributor to the US EPU index volatility on September 2001. The latter imposes a Chinese EPU shock as the main factor contributing to the Chinese EPU index deviation on January 2017.

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