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Impact of United States Economic Policy Uncertainty on Central American Economic Activity: A Bayesian Global VAR Approach

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Abstract

Rising economic policy uncertainty (EPU) in the United States has generated significant global spillovers, posing particular challenges for small open economies in Central America given their strong trade dependence and relatively shallow financial markets. In a context of tightening financial conditions, weakening global growth, and elevated policy uncertainty, understanding the transmission of US uncertainty shocks is crucial for macroeconomic stabilization. This paper estimates the impact of US EPU on Central American economic activity using a Bayesian Global Vector Autoregression (BGVAR) model covering twenty-one economies linked through a trade-weighted network. The framework incorporates Minnesota and stochastic search variable selection (SSVS) priors, along with stochastic volatility, to address parameter dimensionality and time-varying uncertainty. The model includes measures of real activity, inflation, policy rates, and real effective exchange rates for Central American countries and their major trading partners. Generalized impulse response functions indicate that a standardized increase in US EPU significantly reduces US industrial activity, with the contraction in US output serving as the primary transmission channel to the region. Spillover magnitudes and persistence vary across countries and prior specifications, with SSVS generally widening posterior uncertainty. The findings underscore the importance of macro-financial buffers and trade diversification to mitigate vulnerability to external uncertainty shocks.

1. Introduction

The growing uncertainty surrounding economic policy decisions in the United States has generated global repercussions, affecting key variables in emerging economies. Central America, given its high degree of trade integration with the United States, requires a rigorous analysis that captures both direct effects and regional spillovers. In the current macroeconomic environment characterized by rising trade barriers, tightening financial conditions, weakening economic growth and consumer confidence, and increasing economic policy uncertainty (Organisation for Economic Co-operation and Development (OECD), 2025), the Central Banks of Central American countries needs to have tools that allow them to anticipate the impacts of shocks stemming from economic uncertainty abroad on key variables such as economic growth and inflation.

Economic policy uncertainty in the United States can propagate quickly through trade and financial linkages and thereby affect macroeconomic outcomes in small open economies. This is especially relevant for Central American countries that exhibit high bilateral trade shares with the United States and relatively shallow domestic financial markets. Measuring how a shock to US economic policy uncertainty transmits to local activity is therefore of central interest for monetary and fiscal authorities tasked with stabilizing output and inflation. This study provides an empirically grounded assessment of that transmission using a Bayesian Global Vector approach that combines the advantages of large-scale multivariate time series models with Bayesian shrinkage and stochastic volatility.The empirical object of interest is the response of domestic economic activity to a standardized US economic policy uncertainty shock. The EPU index is a widely used measure of policy-related uncertainty and has been shown to correlate with reductions in real activity and investment in advanced economies and in emerging markets. The original construction and validation of the EPU index is well documented in the literature and it is a natural candidate to capture policy-related uncertainty episodes. Recent contributions show that uncertainty episodes can affect both real activity and financial conditions and that these channels interact across countries through trade and financial exposures (Baker et al., 2016; Cuaresma et al., 2016).

A Bayesian GVAR is particularly suitable for this task for three main reasons. First, it formalizes cross-country spillovers through a trade-weighted network so that each country model conditions explicitly on foreign aggregates derived from bilateral trade flows. This helps to capture indirect transmission channels where shocks to a large economy propagate first to major trading partners and then to the rest of the network (Cuaresma et al., 2016; D´ees et al., 2007). Second, the Bayesian treatment with shrinkage priors mitigates overparameterization and improves out-of-sample performance when the cross-sectional dimension is large relative to sample size (Ban´bura et al., 2010; Giannone et al., 2015). Third, allowing for stochastic volatility improves inference on impulse responses during episodes of changing volatility, such as the global financial crisis and the COVID period (Primiceri, 2005). The BGVAR estimation framework used in this paper implements these features in a computationally efficient manner and is described in the package documentation and methodological papers (Cuaresma et al., 2016; Feldkircher & Tondl, 2020).

The choice of variables and countries is motivated by economic mechanisms and data availability. For each country, the model includes year-on-year activity measures and headline inflation together with a short-run policy rate proxy and a real effective exchange rate. For a subset of countries where standard monthly activity measures are unavailable, the industrial production index is used as a consistent alternative. Interest rates are included to capture monetary policy transmission, and where canonical policy rate series are not available, a short-run deposit or interbank rate proxy is used as a passive rate so that the small country specification remains comparable across units. The cross-sectional set includes the Central American economies of interest and a set of global partners that matter for trade and financial linkages. Including major non-regional economies is important to properly capture third-party amplification channels and to avoid omitted variable bias in the construction of foreign aggregates (D´ees et al., 2007).

Methodologically, the Minnesota prior shrinks country-specific autoregressive dynamics towards a parsimonious random walk type benchmark so that spurious large responses due to overfitting are reduced (Banbura et al., 2010; Doan et al., 1984). The stochastic search variable selection prior is used as a complementary specification that is more data-driven and allows the posterior to select relevant coefficients that the overall shrinkage prior might otherwise attenuate (George & McCulloch, 1993; George et al., 2008). Comparing both priors clarifies how inferences depend on prior information versus data evidence. Stochastic volatility is estimated jointly to account for changes in innovation variances that are known to affect identification and the size of impulse responses during crisis episodes (Primiceri, 2005).

Empirical results provide three key findings. First, a one-unit standardized increase in US economic policy uncertainty leads to a statistically significant decline in US industrial activity across specifications. Second, the indirect channel operating via US activity contraction is typically a more robust and larger driver of regional spillovers than the direct contemporaneous effect of the EPU index itself. Third, the magnitude and persistence of spillovers differ across countries and across prior specifications. The SSVS specification tends to widen posterior uncertainty and, in some instances, yields larger median responses than the Minnesota baseline, which is consistent with the fact that SSVS permits less blanket shrinkage and therefore can reveal stronger historical relationships when the data support them (Ban´bura et al., 2010; George & McCulloch, 1993; George et al., 2008; Giannone et al., 2015).

These results are consistent with findings in the GVAR and uncertainty literatures. Previous GVAR studies document that the bulk of short-run variance is typically domestic yet spillovers from major economies can become more important at medium horizons (Cuaresma et al., 2016; Dées et al., 2007). Studies on economic policy uncertainty document that policy-related uncertainty depresses investment and industrial activity and can amplify through trade and financial linkages (Baker et al., 2016; Bloom, 2014).

Our contribution is to combine a scalable Bayesian GVAR implementation with a careful prior comparison and stochastic volatility treatment and to apply this framework to the Central American context, where trade dependence and shallow domestic markets increase vulnerability to external uncertainty. The rest of the paper proceeds as follows. Section two provides a targeted literature review that situates this work relative to global VAR and uncertainty literatures.

Section three describes data construction and model specification, including details on prior choice and the construction of the trade weighting matrix. Section four reports estimation results and robustness checks and section five offers concluding remarks and policy implications.

2. Literature Review

2.1 Economic Policy Uncertainty (EPU)

Periods of heightened international uncertainty regarding global economic growth and inflation inevitably alter the decision–making of economic agents, and such alterations tend to intensify when a greater number of uncertainty triggers converge within the same time frame.

The study of these phenomena is not entirely new, as the analysis of uncertainty effects can be traced back to research from the previous century, such as Bernanke (1983) and Hassler (1996), who emphasized the importance of risk fluctuations in explaining fluctuations in economic activity and identified empirical evidence indicating a negative relationship between uncertainty and overall economic performance.In recent decades, however, uncertainty, its measurement, and its impact on economic activity have gained particular relevance, given the considerable magnitude of the effects that various economic and non–economic events during this period have had on the performance of the world's largest economies. This led to the development of the Economic Policy.

Uncertainty (EPU) index proposed by Baker et al. (2016), has become the standard for quantifying economic policy uncertainty at the global level. According to these authors, the EPU—constructed based on the frequency of coverage in major newspapers—allowed them to identify for the U.S. economy that "at the firm level, uncertainty is associated with greater stock price volatility and reduced investment and employment in policy–sensitive sectors, and at the macroeconomic level, variations in policy uncertainty foreshadow declines in investment, output, and employment" (Baker et al., 2016). These conclusions are consistent with findings in other studies, such as Bloom (2009), who explains that an uncertainty shock leads to declines in output and employment because firms respond to such events by temporarily suspending investment and hiring. This is reaffirmed by the author in a subsequent study (Bloom, 2014). Similar evidence is reported by de Mexico (2019), in an article on the effects of uncertainty on U.S. business investment, which notes that under conditions of high volatility, U.S. firms tend to postpone hiring and investment decisions, resulting in a weakening of fixed business investment.

Other authors, such as Koop et al. (1996) and Jackson et al. (2018), have also investigated the influence of uncertainty on the U.S. economy, finding, in line with the studies mentioned above, that when uncertainty reaches very high levels it not only has an immediate effect on investment but also generates amplified nonlinear macroeconomic impacts that alter the way shocks propagate through the U.S. economy, which may subsequently spill over to other economies.

2.2 Effects of Uncertainty Across Economies and Its Measurement Through Global VAR Models

The current world economy is characterized by a high degree of trade and financial integration. Firms diversify production and investment across domestic and foreign locations, and capital flows face no restrictions. As a result, business cycles are increasingly correlated, and countries are more exposed than ever to global economic disturbances (Boeck et al., 2022). In this regard, authors such as Feldkircher et al. (2015) and Morla (2023) explain that the evident globalization of trade and cross–border financial flows implies that countries are now more exposed to economic shocks and to uncertainty–driven volatility from abroad. This has heightened the interest of policymakers in understanding how international dynamics affect major macroeconomic variables.However, to fully capture these interrelations and, in particular, the exposure of countries to external economic disturbances, conventional methods should not be used. As noted by Cuaresma et al. (2016), standard macroeconomic tools often treat countries as isolated from the rest of the world and may overlook important information for forecasting and counter-factual analysis.

For this reason, they recommend the use of Global Vector Autoregressive (GVAR) models, which systematically and transparently account for the indirect effects of the world economy. Several studies indicate that the empirical literature on GVAR models has been largely influenced by the work of M. H. Pesaran et al. (2004) and that most applications of these models focus on the quantitative assessment of how macroeconomic shocks propagate using historical data (Boeck et al., 2022). For example, Dees et al. (2005) estimate a GVAR model for 26 countries over the period 1979–

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2003, incorporating exogenous components that collectively capture the interdependence between domestic and international macroeconomic factors within the euro area and other countries. For the specific case of applying this methodology to measure the transmission of uncertainty from the United States to other countries, which is the central focus of this study, reference must be made to the works of Bussi`ere et al. (2009) and Kempa and Khan (2019). The former uses a GVAR model applied to a panel of 21 emerging and advanced economies to estimate the effects of different scenarios on trade flows, including a U.S. output shock, a shock to the U.S. real effective exchange rate, and shocks to foreign variables.

The latter use the same model to analyze the international contagion effects of trade restrictions imposed by the United States, modeled as a reduction in its imports, as well as the effects of a trade war in which the rest of the world responds symmetrically to the U.S. measures, examining impacts on variables such as economic growth and the trade balance. For a more specific approach to the analysis of uncertainty and its transmission from the United States to other economies, it is necessary to refer to the study by Feldkircher et al. (2015), who, through a GVAR model, show that a monetary policy path anticipated by Federal Reserve officials with restrictive stances generates cumulative losses in global output, while prices respond only marginally.

Similarly, Tam (2018) quantifies, also using a GVAR, the impacts of economic policy uncertainty (EPU) on global trade flows, providing evidence of the importance of policy uncertainty in China and the United States for influencing exports and imports of individual countries. Finally, although the aforementioned studies provide clear evidence of U.S. uncertainty spillovers to key macroeconomic variables in advanced and emerging economies, there is little evidence specifically focused on countries in the Central American region. Within this limited body of evidence we find only the works of Cuaresma et al. (2016) and Morla (2023).

Building on the findings of Huber and Feldkircher (2017) regarding the improved capacity of Bayesian GVAR models to deliver more precise parameter estimates and impulse–response functions, these authors apply a Bayesian variant of the traditional GVAR model, incorporating priors, to explore the transmission mechanisms of international macroeconomic shocks.

The work of Morla (2023) constitutes the closest application of this methodology to the Central American countries and the Dominican Republic, and therefore stands as one of the foundational references for the present study.

The author employs a Bayesian Global VAR model to examine the transmission of international shocks to the countries under study, finding results that reveal substantial geographic interlinkages in the propagation of inflationary pressures and the vulnerability of the Dominican Republic and Central American economies to such phenomena.

3. Methodology

3.1 Econometric Framework

3.1.1 Model Estimation

The methodological framework builds on the Bayesian GVAR approach developed by Boeck et al. (2022) and its recent empirical application by Morla (2023). The procedure consists of two clearly differentiated stages.

First, individual VAR models are estimated for each country, and second, these models are integrated to generate a global structure. ***Let*** $i = 1, \ldots, N$ denote the cross-sectional units (countries). For each unit, we define the vector of endogenous variables y_{it} of dimension k_i which evolves according to a VAR of order ***P,*** complemented with contemporaneous and lagged weakly exogenous variables of dimension k_i^*, denoted by y_i^{*t}.

These exogenous regressors are constructed as weighted averages of the other units, using predefined weights w_{ij} (with $\sum_i w_{ij} = 1$ and $w_{ii} = 0$)

such that: $y_i^{*t} = \sum_{j=1}^{N} w_{ij} y_{it}$

While here we simplify by applying the same weights to all components, more complex weighting schemes can be adopted depending on the type of variable.

The country-specific VAR model is specified as:

$$y_{it} = a_{i0} + a_{i1}t + \sum_{j=1}^{P} \Phi_{ij} y_{i,t-j} + \sum_{q=0}^{Q} \Lambda_{iq} y_i^{*t-q} + \varepsilon_{it} \quad ,\varepsilon_{it} \sim N(0,\Sigma_{it})$$

where $a_{i0}, a_{i1} \in R^{k_i}$ are deterministic parameters (intercept and trend), Φ_{ij} *is a* $k_i \times k_i$ matrix associated with own lags, Λ_{iq} *is a* $k_i \times k_i^*$, matrix linked to weakly exogenous variables, and ε_{it} is the error term with variance Σ_{it}. Errors are assumed independent across units, which implies that the global variance Σ_t is a block-diagonal matrix $diag(\Sigma_{1t}, \cdots, \Sigma_{Nt})$.

Furthermore, Σ_{it} is factorized as $\Sigma_{it} = V_i D_{it} V_i^{\top}$, where V_i is lower triangular with ones on the diagonal, and $D_{it} = diag(e_{di1,t}, \cdots, e_{diki,t})$ with entries following AR(1) processes. The second step combines all submodels algebraically to derive a global GVAR representation. In the simplified case with $P = Q = 1$, each unit satisfies:

$$y_{it} = a_{i0} + a_{i1}t + \Phi_{i1} y_{i,t-1} + \Lambda_{i0} y_i^{*t} + \Lambda_{i1} y_i^{*t-1} + \varepsilon_{it}$$

To express this in terms of the global vector $y_t = (y_{1t}^{\top}, \ldots y_{Nt}^{\top})^2 \in R^k$ we define connection matrices G_i and selection matrices S_i (both of dimension $k_i \times k$). Applying these matrices yields:

$$S_i y_t = a_{i0} + a_{i1}t + \Phi_{i1} S_i y_{t-1} + \Lambda_{i0} G_i y_t + \Lambda_{i1} G_i y_{t-1} + \varepsilon_{it}.$$

Rearranging terms gives:

$$(S_i - \Lambda_{i0} G_i) y_t = a_{i0} + a_{i1}t + (\Phi_{i1} S_i + \Lambda_{i1} G_i) y_{t-1} + \varepsilon_{it}.$$

Stacking for $i = 1, \ldots, N$ results in:

$$G y_t = a_0 + a_1 t + H_1 y_{t-1} + \varepsilon_t,$$

with G, a_0, a_1, H_1 *and* ε_t concatenating the blocks of each unit. By inverting ***G***, the global GVAR is obtained as:

$$y_t = G^{-1}a_0 + G^{-1}a_1 t + G^{-1}H_1 y_{t-1} + G^{-1}\varepsilon_t$$
$$= b_0 + b_1 t + F_1 y_{t-1} + e_t,$$

where $e_t \sim N(0, \Sigma_{e_t})$ with

$\Sigma_{e_t} = G^{-1}\Sigma_t (G^{-1})^{\top}$. In the general case with ***P*=max (P,Q)***, the model takes the form:

$$y_t = b_0 + b_1 t + \sum_{j=1}^{P^*} F_j y_{t-j} + e_t, \quad e_t \sim N(0, \Sigma_{et}), \text{(1)}$$

Which resembles a standard VAR of dimension ***k.***

3.1.2 Prior Specification

Given the high dimensionality of the GVAR, we implement a Bayesian estimation strategy. Instead of the SSVS prior employed by Boeck et al. (2022) and Morla (2023), we adopt the Minnesota-type prior available in the BGVAR package for R. This prior applies differently across parameter blocks. For own lags, the variance is λ_1^2/k^2; for lags of other endogenous variables, it is: $(\sigma_j^2/\sigma_s^2)\lambda_1^2\lambda_2^2/k^2$

where σ_j *and* σ_s are the standard deviations of OLS residuals from the corresponding univariate AR processes; and for weakly exogenous variables, it is $\left(\sigma_j^2/\sigma_s^2\right)\lambda_1^2\lambda_3^2/(k+1)^2$.

The deterministic block is assigned variance λ_4, sufficiently large to remain uninformative. To enhance flexibility, the hyperparameters $\lambda_1, \lambda_2, \lambda_3$, follow weak Gamma priors, which allows their joint estimation with the model coefficients, consistent with the approach of Giannone et al. (2015).

We also adopt a Bayesian approach based on Stochastic Search Variable Selection (SSVS) to impose regularization and ensure model parsimony and test for robustness. Specifically, each coefficient ψ_{ij} from the parameter set Ψ is assigned a mixture of two normal priors:

$$\psi_{ij} \mid \delta_{ij} \sim (1-\delta_{ij})N\left(0,\tau_0^2\right) + \delta_{ij}N\left(0,\tau_1^2\right), \delta \sim \textit{Bernoulli (p)},$$

Where $\tau_0^2 \ll \tau_1^2$ determines the shrinkage magnitude toward zero when $\delta_{ij} = 0$, and p reflects the prior inclusion probability of the variable in the model. By incorporating these mixture priors, the SSVS procedure ensures that only coefficients with sufficient empirical support escape the strong shrinkage toward zero, while the remaining parameters are pushed toward negligible values. Inference proceeds through a Gibbs sampling scheme that alternates between updating the binary indicators δ_{ij} and the coefficients ψ_{ij}.

This facilitates the automatic identification of the most relevant relationships without requiring manual pre-selection of regressors. The SSVS prior has proven effective in high-dimensional settings, providing numerical stability and robust inference in large-scale GVAR models.

3.1.3 Structural Inference

One of the main utilities of the GVAR framework lies in the analysis of impulse responses (IRFs), both generalized (GIRFs as in H. Pesaran and Shin (1998), orthogonalized (via Cholesky decomposition), or signrestricted. Local identification is achieved by applying the Cholesky decomposition only to the block corresponding to the unit subject to the shock. For example, for unit 1 we define the factor Q_1^{-1} of the variance Σ_{1t} and transform the model as follows:

$$Q_1^{-1} y_{1t} = Q_1^{-1} a_{10} + Q_1^{-1} a_{11} t + \sum_{j=1}^{P} \Phi_{1j} y_{1,t-j} + \sum_{q=0}^{Q} \Lambda_{1q} y_1^{*t-q} + Q_1^{-1} \varepsilon_{1t}$$

We then define a block-diagonal matrix

$$Q = diag(Q_1, I_{k_2}, \ldots, I_{k_N}),$$

so that the global structural errors are given by: $u_t = Q^{-1} \varepsilon_t = (v_{1t}, \varepsilon_{2t}, \ldots, \varepsilon_{Nt})$. This ensures that the shock in unit 1 is orthogonalized, while the effects on the remaining units are interpreted as GIRFs. This approach thus allows the analysis of both domestic and international responses to specific shocks, combining different identification strategies depending on the analytical context.

3.2 Data

The model includes the principal trading partners of the Central American region; the full list of countries is shown in Table 1. Bilateral trade shares are used to build the weighting matrix that defines foreign aggregates for each domestic equation. These trade-weighted foreign variables constitute the weakly exogenous block that summarizes each economy's exposure to external demand, price and financial conditions, consistent with standard practice in the GVAR literature (Chudik & Pesaran, 2016; M. H. Pesaran et al., 2004). For each country we include a small set of monthly macroeconomic indicators that are commonly used in the GVAR literature (Trung, 2019) following the original development of the model from M. H. Pesaran et al. (2004) and the main repository of the model version for MATLAB from Mohaddes and Raissi (2020). The price variable is the year-on-year change in the Consumer Price Index, π_{it}.

The measure of economic activity, y_{it}, is the year-onyear percent change of the Monthly Index of Economic Activity where available; when such a series is not published the Industrial Production Index is used instead. We follow the commonly used database and activity definition as in the GVAR literature (Abdel-Latif & Popescu, 2025; Feldkircher & Tondl, 2020; Trung, 2019) and in the EPU spillovers literature (Baker et al., 2016; Coronado et al., 2020; Giraldo et al., 2025; Jurado et al., 2015). Monthly activity indicators provide timely information about short-run dynamics and avoid the large real-time uncertainty attached to output-gap estimates (Orphanides & van Norden, 2002). For robustness we estimate alternative specifications that use output-gap measures. The monetary stance enters the model through a short-term interest rate series, i_{it}. When the official central bank policy rate is reported and unambiguous it is used as the preferred measure of the policy instrument.

If an official policy rate is not available for a country the short-run passive market rate (for example a short-term deposit rate) is adopted as a proxy. Using a domestic short-term bank rate as a proxy for the policy stance is a common practical choice in cross-country empirical work because pass-through from the policy instrument to market rates is typically

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Table 1: Bilateral Trade Weighting Matrix for the Period 2003-2022

Country	CR	DR	GT	HN	NI	PA	SL	AR	BR	CA	CH	CN	CO	EC	IN	JP	KO	ME	UK	US	EZ
CR	0	0.00830	0.03196	0.01537	0.01792	0.01604	0.01610	0.00424	0.01929	0.01911	0.00830	0.10711	0.01195	0.00222	0.00597	0.02866	0.01298	0.07423	0.04663	0.42580	0.12783
DR	0.01265	0	0.00823	0.00384	0.00110	0.00151	0.00436	0.00897	0.02225	0.03057	0.00402	0.07959	0.02444	0.00321	0.01692	0.00854	0.01396	0.04254	0.01461	0.59043	0.10824
GT	0.03697	0.00777	0	0.04553	0.01947	0.00820	0.07405	0.00597	0.01344	0.02348	0.01060	0.08014	0.01591	0.00631	0.00979	0.01915	0.02190	0.09600	0.00582	0.42596	0.07356
HN	0.02603	0.00628	0.07050	0	0.01404	0.00256	0.06968	0.00275	0.00748	0.01585	0.00282	0.04331	0.00754	0.00366	0.00809	0.00851	0.01096	0.05445	0.00764	0.55987	0.07797
NI	0.06371	0.00517	0.06628	0.03231	0	0.00409	0.06224	0.00458	0.00958	0.02170	0.00351	0.05193	0.00359	0.00653	0.00660	0.01133	0.01858	0.11566	0.00564	0.45577	0.05123
PA	0.01807	0.00385	0.01940	0.00660	0.00139	0	0.00452	0.00325	0.01987	0.00393	0.00604	0.21405	0.05126	0.05634	0.00870	0.21629	0.07742	0.02553	0.00778	0.17571	0.08000
SL	0.03774	0.00780	0.15610	0.06002	0.03889	0.00845	0	0.00318	0.01236	0.01009	0.00623	0.05579	0.00992	0.00828	0.00459	0.01195	0.01185	0.06827	0.00390	0.42178	0.06280
AR	0.00118	0.00231	0.00168	0.00064	0.00061	0.00090	0.00073	0	0.30408	0.01611	0.05856	0.14333	0.01401	0.00692	0.02575	0.01878	0.01796	0.02819	0.01573	0.14774	0.19479
BR	0.00132	0.00225	0.00106	0.00048	0.00037	0.00067	0.00073	0.08595	0	0.01890	0.02925	0.25735	0.01210	0.00313	0.02399	0.04132	0.03386	0.03149	0.02249	0.21263	0.22066
CA	0.00030	0.00068	0.00037	0.00012	0.00018	0.00013	0.00015	0.00178	0.00686	0	0.00267	0.06131	0.00168	0.00050	0.00617	0.02637	0.01232	0.02542	0.02504	0.76359	0.06436
CH	0.00236	0.00072	0.00204	0.00051	0.00032	0.00069	0.00087	0.04166	0.07459	0.02038	0	0.27226	0.01495	0.01493	0.01950	0.07806	0.05331	0.03137	0.01414	0.20352	0.15382
CN	0.00101	0.00121	0.00092	0.00044	0.00031	0.00054	0.00066	0.00749	0.03729	0.03540	0.01662	0	0.00557	0.00223	0.03431	0.15423	0.11084	0.03614	0.03939	0.27872	0.23668
CO	0.00555	0.00802	0.00636	0.00242	0.00067	0.00456	0.00199	0.01548	0.05098	0.01931	0.02713	0.13114	0	0.03567	0.02683	0.02368	0.02016	0.06235	0.01808	0.38679	0.15282
EC	0.00194	0.00282	0.00448	0.00354	0.00227	0.00156	0.00455	0.01848	0.02839	0.01532	0.05392	0.12981	0.07570	0	0.01699	0.03788	0.02358	0.02357	0.00981	0.40124	0.14415
IN	0.00051	0.00145	0.00089	0.00041	0.00033	0.00039	0.00034	0.00879	0.02702	0.02020	0.00803	0.24845	0.00639	0.00149	0	0.05075	0.06073	0.02136	0.05390	0.23678	0.25177
JP	0.00079	0.00054	0.00054	0.00021	0.00018	0.00042	0.00030	0.00216	0.01372	0.02885	0.01092	0.35828	0.00202	0.00107	0.01734	0	0.10466	0.02484	0.02588	0.25694	0.15035
KO	0.00049	0.00045	0.00088	0.00019	0.00015	0.00052	0.00034	0.00266	0.01623	0.01825	0.01057	0.40828	0.00217	0.00134	0.02669	0.14311	0	0.02925	0.01907	0.18869	0.13066
ME	0.00187	0.00140	0.00340	0.00098	0.00086	0.00071	0.00136	0.00383	0.01247	0.04312	0.00526	0.05884	0.00694	0.00117	0.00708	0.02133	0.01841	0	0.00570	0.73113	0.07413
UK	0.00040	0.00038	0.00014	0.00012	0.00007	0.00010	0.00007	0.00160	0.00798	0.02843	0.00183	0.08430	0.00139	0.00039	0.01776	0.02738	0.01412	0.00573	0	0.14575	0.66205
US	0.00396	0.00473	0.00370	0.00174	0.00076	0.00107	0.00214	0.00408	0.02037	0.20932	0.00778	0.17291	0.00932	0.00442	0.02256	0.07820	0.04084	0.17659	0.04255	0	0.19296
EZ	0.00152	0.00120	0.00091	0.00065	0.00022	0.00053	0.00040	0.00829	0.03349	0.02908	0.00874	0.21409	0.00537	0.00223	0.03581	0.06486	0.04042	0.02670	0.25152	0.27399	0

Notes: Entries report average bilateral trade weights for the period 2003–2022. Rows represent the reporting country, while columns represent the trading partner country. Each row is normalized so that the weights sum to one after excluding domestic trade. Accordingly, diagonal elements are set to zero, as own-country trade is not included in the construction of foreign aggregates. These weights are used to construct the trade-weighted foreign variables (y_i^*) employed in the Bayesian Global Vector Autoregressive (BGVAR) model. Country codes are defined as follows: CR = Costa Rica, DR = Dominican Republic, GT = Guatemala, HN = Honduras, NI = Nicaragua, PA = Panama, SL = El Salvador, AR = Argentina, BR = Brazil, CA = Canada, CH = Chile, CN = China, CO = Colombia, EC = Ecuador, IN = India, JP = Japan, KO = South Korea, ME = Mexico, UK = United Kingdom, US = United States, and EZ = Euro Area.

Source: Authors' estimates based on IMF data.

incomplete and heterogeneous but sufficiently informative for capturing monetary conditions in small open economies (Beyer et al., 2024; Illes & Lombardi, 2019) and also is commonly used in the GVAR literature (Abdel-Latif & Popescu, 2025; Boeck et al., 2022; Feldkircher & Tondl, 2020; Trung, 2019).

Finally, the vector of global control variables, dt, contains internationally relevant price indicators such as the natural logarithm of WTI crude oil, poil,t, and the natural logarithm of the global food price index, pa,t. Although these factors operate transversally, their specific impact on each economy depends on structural characteristics such as the degree of energy import dependence and the composition of the consumption basket. The bilateral trade weighting matrices in Tables 1 and 2 make explicit the exposure basis used to construct the foreign aggregates for the period 2003–2022.

Data for the Central American economies were obtained from the Executive Secretariat of the Central American Monetary Council (SECMCA). For the remaining countries, information was collected from their respective Central Banks and from international

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organizations that regularly publish updated economic reports and statistics, particularly the Organisation for Economic Co-operation and Development (OECD) and the International Monetary Fund (IMF) through its International Financial Statistics. The selected frequency of the data is monthly, covering the period from January 2003 to December 2024. The country codes used are presented in Table 3.

3.3 Convergence diagnostics and implications for model selection

A significant Geweke statistic for a parameter provides evidence that the early and late parts of the chain have different means which is inconsistent with draws emanating from a single stationary posterior distribution. A high share of flagged parameters signals incomplete mixing or insufficient run length for that parameter under the current sampler settings and identification choices. Because Geweke uses only one chain its power to detect complex mixing problems is limited relative to diagnostics that exploit multiple chains but it remains useful as a fast single chain check (Boeck et al., 2022).

The empirical convergence rates reported in Table 4 can therefore be read as direct indicators of relative sampler performance across prior choices and cross-sectional sizes. These numbers show two clear patterns. First short runs of 1,000 draws produce large fractions of flagged parameters across all

Table 2: Bilateral Trade Weighting Matrix for 12-Country Model, 2003–2022

	CR	DR	GT	HN	NI	PA	SL	CN	ME	UK	US	EZ
CR	0	0.00935	0.03602	0.01732	0.02019	0.01808	0.01815	0.12072	0.08366	0.05255	0.47989	0.14407
DR	0.01459	0	0.00949	0.00443	0.00127	0.00174	0.00503	0.09179	0.04906	0.01685	0.68092	0.12483
GT	0.04232	0.0089	0	0.05213	0.02229	0.00939	0.08478	0.09175	0.10991	0.00666	0.48767	0.08422
HN	0.02792	0.00674	0.07562	0	0.01505	0.00275	0.07474	0.04645	0.0584	0.0082	0.60051	0.08363
NI	0.0697	0.00566	0.07251	0.03535	0	0.00447	0.06809	0.05682	0.12654	0.00617	0.49865	0.05604
PA	0.03246	0.00691	0.03483	0.01185	0.0025	0	0.00812	0.38436	0.04584	0.01397	0.31552	0.14365
SL	0.04095	0.00847	0.16938	0.06513	0.0422	0.00917	0	0.06054	0.07408	0.00423	0.45769	0.06814
CN	0.0017	0.00203	0.00154	0.00074	0.00052	0.0009	0.00111	0	0.06063	0.06609	0.46763	0.39709
ME	0.00212	0.00159	0.00386	0.00111	0.00098	0.0008	0.00155	0.06684	0	0.00647	0.83048	0.0842
UK	0.00045	0.00042	0.00015	0.00014	0.00008	0.00011	0.00008	0.09376	0.00638	0	0.1621	0.73634
US	0.00656	0.00785	0.00613	0.00289	0.00127	0.00177	0.00355	0.28669	0.29281	0.07055	0	0.31994
EZ	0.00197	0.00155	0.00118	0.00084	0.00028	0.00069	0.00052	0.27742	0.0346	0.32592	0.35504	0

Notes: The matrix reports average bilateral trade weights for the 12 countries included in the Bayesian Global Vector Autoregressive (BGVAR) model over the period 2003–2022. Rows represent the reporting country, whereas columns represent the trading partner country. Each entry indicates the share of a country's foreign trade associated with the corresponding trading partner. Diagonal elements are set to zero because domestic trade is excluded from the construction of foreign aggregates, and each row is normalized to sum to one after excluding own-country trade. These trade weights are used to construct the trade-weighted foreign variables (y_i^*) employed in the BGVAR framework. Country codes are defined as follows: CR = Costa Rica, DR = Dominican Republic, GT = Guatemala, HN = Honduras, NI = Nicaragua, PA = Panama, SL = El Salvador, CN = China, ME = Mexico, UK = United Kingdom, US = United States, and EZ = Euro Area.

Source: Authors' estimates based on International Monetary Fund (IMF) data.

specifications. Second increasing the number of draws generally reduces the flagged share and the largest improvement occurs when runs are extended to 5,000 and again when the 21 country runs reach 20,000 draws. The smallest flagged share in the table is 9.11 percent for the Minnesota prior with 21 countries at 20,000 draws. This outcome supports treating that specification as the best available candidate from the perspective of numerical stability of the MCMC output.

Important caveats apply before declaring a specification as objectively the best for substantive inference. Geweke measures chain stationarity but it does not evaluate model fit predictive performance or identification robustness. A specification with excellent Geweke behavior can still be misspecified or yield poor out of sample forecasts. Conversely a more flexible prior such as SSVS require substantially longer runs to achieve comparable Geweke metrics because it places less shrinkage on coefficients and therefore explores a higher dimensional posterior more slowly and it is computationally expensive. Thus, higher flagged shares for SSVS in the table indicate slower mixing rather than a priori inferiority of the prior family.

We follow the best specification model after convergence of 21 countries and Minnesota prior with 20,000 draws and burn-in.The selection of the 21-country model seeks to balance global representativeness and econometric tractability.

Table 3: Country Codes International Code

International Code	*Country*
*CR	Costa Rica
*DR	Dominican Republic
*GT	Guatemala
*HN	Honduras
*NI	Nicaragua
*PA	Panama
*SL	El Salvador
AR	Argentina
BR	Brazil
CA	Canada
CH	Chile
*CN	China
CO	Colombia
EC	Ecuador
IN	India
JP	Japan
KO	Korea
*ME	Mexico
*UK	United Kingdom
*US	United States
*EZ	Eurozone

Note: Countries mark with * are those included in the 12 country BGVAR model.

Table 4: Share of parameters flagged by the Geweke diagnostic across draw schemes

Draws and burn ins	Minnesota 12	Minnesota 21	SSVS 12	SSVS 21
1,000	20.06%	24.07%	22.77%	25.99%
5,000	11.69%	14.22%	16.12%	16.03%
8,000	**9.60%**	14.58%	**11.54%**	16.49%
10,000	11.93%	15.09%	15.73%	17.91%
20,000	10.75%	**9.11%**	12.30%	**13.65%**

Note: Each cell reports the percentage of model parameters flagged by the Geweke convergence diagnostic under the corresponding draw scheme.

Lower values indicate better convergence performance. Bold entries denote the lowest percentage flagged within each model specification across the reported numbers of draws. Minnesota refers to the Minnesota prior, whereas SSVS refers to stochastic search variable selection. The numbers 12 and 21 denote the two model configurations considered in the estimation, based on the number of countries included.

Source:Authors' estimates.

The aim is to capture both the direct transmission of a U.S. Economic Policy Uncertainty (EPU) shock to Central American economies and the broader, internationally mediated spillovers that operate through trade and financial linkages. Including the eight Central American and Caribbean economies of interest together with their main global partners ensures that the foreign aggregates entering each country equation reflect the actual external exposure of small open economies to external demand, price, and financial shocks. Constructing these aggregates using bilateral trade weights, as reported in Table 1, follows the standard approach in the GVAR literature (D´ees et al., 2007; M. H. Pesaran et al., 2004).

At the same time, identification of an external uncertainty shock requires that the model include the main global sources of macroeconomic variation, most notably the United States, the Euro area, China, and other large trading partners such as Mexico, Canada, Japan, and Korea. Excluding these economies would risk attributing to regional transmission mechanisms what are, in fact, globally propagated effects, thereby biasing the impulse responses for small open economies (Boeck et al., 2022; Feldkircher et al., 2015).

Nevertheless, there is a trade-off between comprehensiveness and overparameterization. A very large country set increases dimensionality and can compromise convergence and precision. The Bayesian GVAR framework mitigates this issue through shrinkage priors, allowing inclusion of a broader but still economically meaningful set of countries while preserving stability and interpretability of the posterior draws (Boeck et al., 2022; Cuaresma et al., 2016).

The resulting 21-country specification is therefore a balanced configuration, it is large enough to represent the key channels of international transmission documented in the literature, but parsimonious enough to ensure efficient estimation and clear structural inference.

Finally, a smaller 12-country version, comprising the Central American block and its most influential partners, is also estimated as a robustness check. This nested specification validates that the qualitative dynamics of the generalized impulse responses to U.S. EPU shocks are not artifacts of the extended model size.

The bilateral trade weighting matrices in Tables 1 and 2 document the economic structure underpinning the foreign variable construction and make explicit the hannels through which external uncertainty is transmitted across countries.

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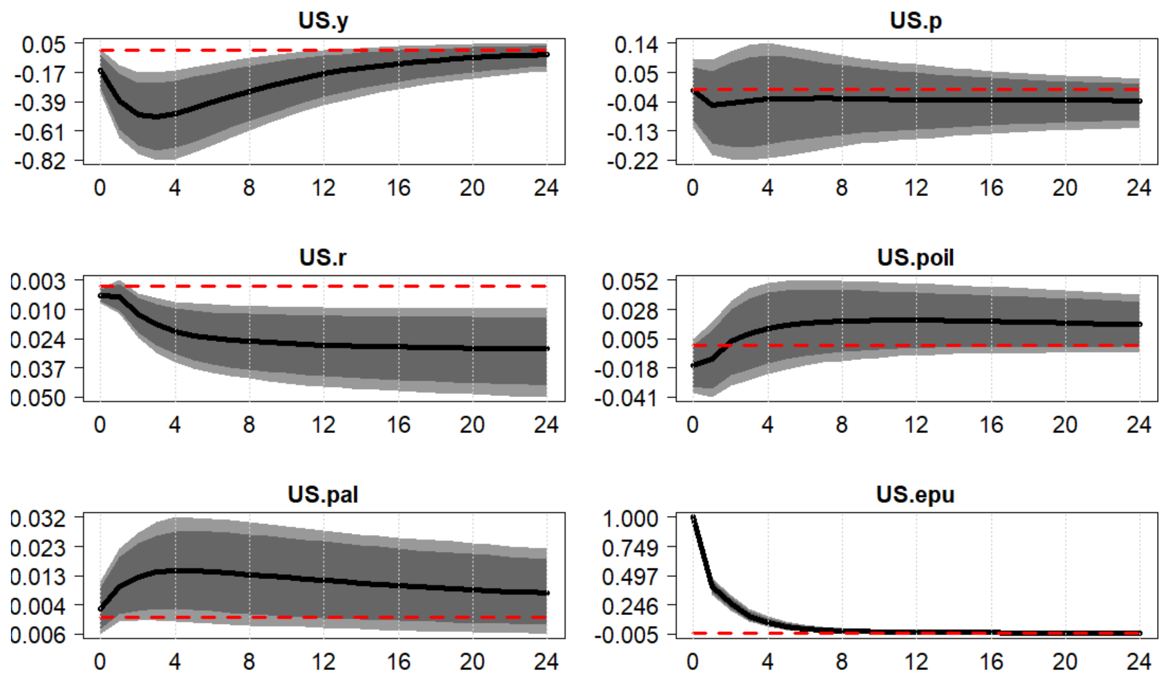
Figure 1: Generalized Impulse Response Functions of a U.S. EPU Shock on U.S. Macroeco-nomic Variables



Note: The figure reports generalized impulse response functions to a one-standard-deviation shock in U.S. economic policy uncertainty (EPU). Responses correspond to U.S. macroeconomic variables included in the BGVAR model and are computed using the Minnesota prior under the 21-country specification with 20,000 posterior draws. The horizontal axis represents the forecast horizon in periods, while the vertical axis shows the magnitude of the response. Generalized impulse responses are invariant to variable ordering and capture the dynamic effects of the shock on each variable over time.

Source: Authors' estimates.

We performed CUSUM and Nyblom stability tests for each country equation in the BGVAR, results are shown in Table 5. The CUSUM test rejects parameter stability in 4.1% of equations while the Nyblom test rejects in 42.3% of equations. The contrast indicates few abrupt breaks but substantial evidence of gradual parameter variation or changes in error volatility. The CUSUM test rarely rejects stability in many cases. This is consistent with the pattern described by D´ees et al. (2007) where mass rejections in sequential tests often reflect changes in error variance rather than systematic changes in short-run coefficients.

4. Results

In this section, we examine the spillover effects of the U.S. EPU shock on the economic activity of other economies. Before proceeding, we first focus on confirming the robustness of our BGVAR model by analyzing the effects on the U.S. economy. The estimated response of real output is negative for a period exceeding 12 months. However, this effect is small, reaching its lowest point approximately three months after the shock. This finding can be attributed to the ”wait-and-see” channel, which has been widely documented for the U.S.economy (Baker et al., 2016; Bloom, 2009).

Figure 1 presents the generalized impulse response functions for the United States, revealing a clear and homogeneous reaction pattern to the economic policy uncertainty shock. First, output (US.y) exhibits an immediate decline of approximately –0.38 percentage points in the first month, then gradually recovers to the trend level by month twelve. Inflation (US.p) experiences a slight decrease of about -0.05 percentage points during the initial quarter, before fading almost linearly over the following 18 months; however, this change is not statistically significant. The deflation process is attributed to the deceleration of the economic activity.

The nominal interest rate (US.r) responds with a very moderate downward adjustment of approximately -0.004 points, and rapidly decreasing from the second month until twelve months after the shock, suggesting an accommodative bias in monetary policy. The economic policy uncertainty index (US.epu) follows the typical rapid transition profile, with an initial peak defined in +1.0 unit in month zero and a near-complete decay by month four, indicating that the worsening of uncertainty itself is transitory.

Meanwhile, the oil price (US.poil) appreciates slightly by up to +0.02 percentage points during the first four months before stabilizing, indicating some transmission of global uncertainty to energy markets, related to a lower aggregated demand. The food price index (US.pal) shows an increase of 0.015 percentage points, peaking in month four and progressively returning to its initial level. Overall, the response of the United States is characterized by a rapid initial peak, moderate magnitude, and full recovery in the medium term.

The analysis of the generalized impulse response functions (GIRFs, see Fig. 2) under the BGVAR model with a Minnesota prior, endogenous and weakly exogenous lag of order two, and stochastic volatility reveals heterogeneous reactions of the year-on-year growth rate of economic activity to a positive 1sd shock in the economic policy uncertainty index from US. The point estimate (black line) for most economies shows an initial contraction followed by a gradual recovery toward the equilibrium value, while the 68% and 90% credibility bands (dark and light gray areas) indicate high uncertainty during the first months that diminishes over the forecast horizon.

In the short term (0–3 months), Costa Rica (CR), the Dominican Republic (DR), and Honduras (HN) exhibit pronounced declines of approximately –0.35, –0.30, and –0.30 percentage points, respectively, with credibility bands that do not include zero in the initial months, indicating robust statistical significance. Nicaragua (NI) also shows a severe contraction near –0.12, with bands excluding zero in the short term, confirming statistical significance in this sample. El Salvador (SL), although experiencing a smaller initial impact (–0.6 points), has bands that include zero in the first periods, suggesting a non-significant response.

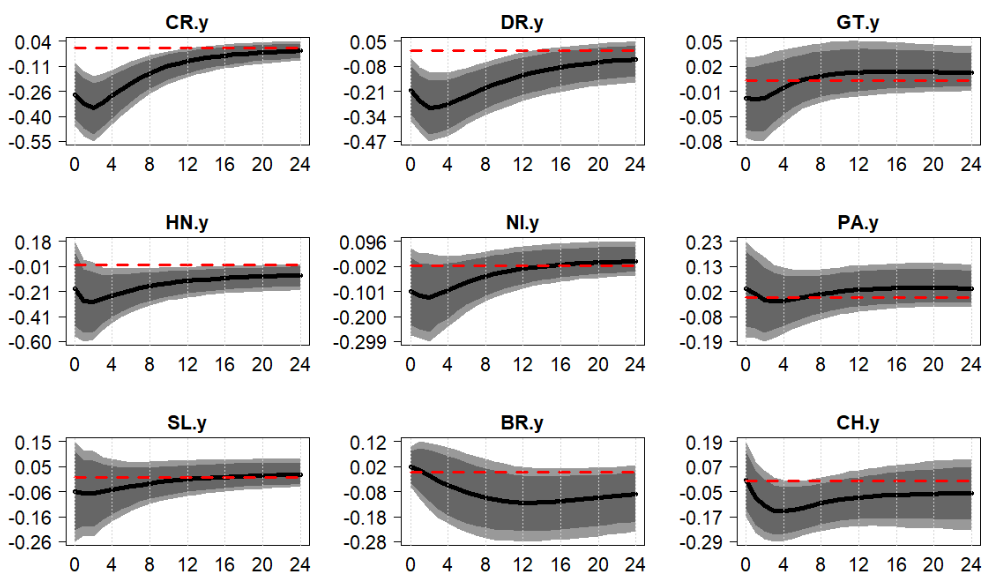
Table 5: Number of rejections of the null of parameter constancy per variable across the country-specific models at the 5% level - CUSUM y Nyblom test

Variable	N eq.	CUSUM rech.	CUSUM pct.	Nyblom rech.	Nyblom pct.
Activity (y)	21	0	0.0	5	23.8
Inflation (p)	21	0	0.0	8	38.1
Interest rate (r)	21	1	4.8	15	71.4
Real exchange rate (e)	20	0	0.0	8	40.0
EPU	12	3	25.0	5	41.7
Oil price	1	0	0.0	0	0.0
Food price	1	0	0.0	0	0.0
Total	97	4	4.1	41	42.3

Note: The table reports the number and percentage of rejections of the null hypothesis of parameter constancy across country-specific equations in the BGVAR model. The column N eq. indicates the number of estimated equations for each variable. CUSUM rech. and Nyblom rech. report the number of equations for which the null hypothesis of parameter stability is rejected at the 5% significance level according to the CUSUM and Nyblom tests, respectively. The corresponding percentages are presented in CUSUM pct. and Nyblom pct.. Higher values indicate stronger evidence of parameter instability. The Total row summarizes the results across all estimated equations.

Source: Authors' estimates.

Figure 2: Generalized Impulse Response Functions of a U.S. EPU Shock on Central American Economic Activity



Note: The figure presents generalized impulse response functions (GIRFs) of real economic activity in Central American countries following a one-standard-deviation shock to U.S. economic policy uncertainty (EPU). The responses are derived from the BGVAR model estimated using the Minnesota prior under the 21-country specification and based on 20,000 posterior draws. The horizontal axis represents the forecast horizon (in periods), while the vertical axis measures the magnitude of the response for each country. Generalized impulse response functions are invariant to variable ordering and capture the dynamic transmission of uncertainty shocks from the United States to Central American economies over time.

Source: Authors' estimates.

In contrast, Panama (PA) and Guatemala (GT) shows no statistically significant responses. Their credibility intervals include zero throughout the entire horizon, despite median estimates indicating mild contraction. This lack of significance suggests that the effects of economic policy uncertainty on activity in these economies are weaker or dominated by other compensating factors specific to local conditions. Chile (CH), on the other hand, exhibits a moderate and marginally significant response in the early periods (drop of up to –0.15), but the bands quickly widen to include zero, reducing the robustness of this result. In the medium term (6–12 months), most economies gradually return to their trend trajectory. By month 12, responses stabilize around zero, although slight asymmetries persist. The dispersion of the bands decreases progressively, reflecting reduced uncertainty following the initial shock.

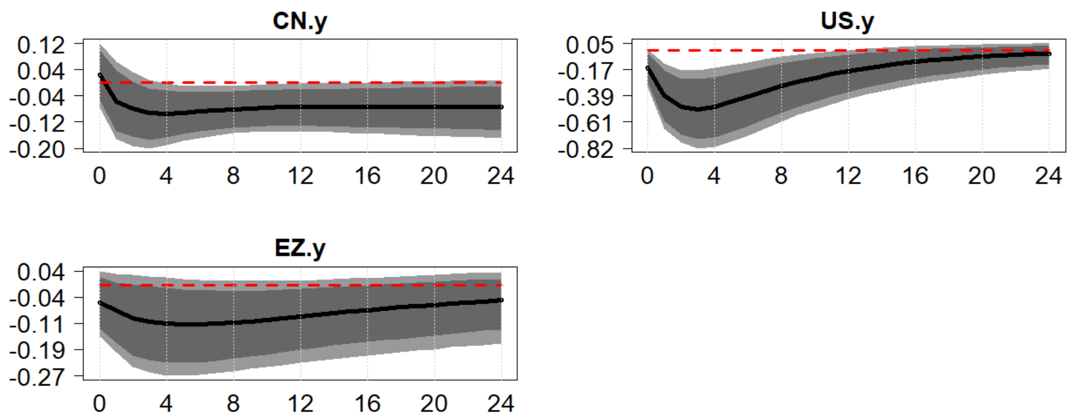
Extra regional economies, including China (CN), the United States (US), and the Eurozone (EZ), also display distinct patterns (see Fig. 3). China exhibits a moderate negative impact, with an initial contraction of approximately –0.10 points and credibility bands remaining below zero, evidencing a significant and persistent effect of US uncertainty on its industrial activity. The United States experiences a deeper decline of around –0.50 points in the first three months, followed by a gradual recovery, reflecting rapid absorption of the shock due to automatic stabilizers and strong domestic demand. The Eurozone experiences an initial decline of about –0.12 points, with slightly wider bands; while statistically significant at the 68% level during the first four months, the 90% intervals include zero from month three onward, indicating greater uncertainty in shock transmission within a heterogeneous bloc.

Overall, the results emphasized the diversity in the transmission of political uncertainty shocks across the analyzed economies. While Costa Rica, the Dominican Republic, El Salvador, Nicaragua, and Honduras show short-term negative and statistically significant effects, Panama, Guatemala, Brazil, and, to a lesser extent, Chile, do not exhibit significant responses, which may reflect differences in institutional structures, expectation anchoring, and the degree of economic integration in each country.

Additionally, Figure 4 presents the generalized impulse response functions of a U.S. EPU shock on inflation in Central American economies. The results indicate no evidence that an uncertainty shock generates inflationary pressures. In fact, the direction of the responses is predominantly negative, reflecting that uncertainty tends to contract economic activity and thus produces deflationary pressures. However, with the exception of El Salvador and Costa Rica, there is no statistical evidence that the shock is significant. These findings are consistent with previous studies such as Largaespada (2021) and Giraldo et al. (2025).

The results show that a U.S. economic policy uncertainty shock transmits heterogeneously across countries and that not all effects are significant. Previously, the direct impact of

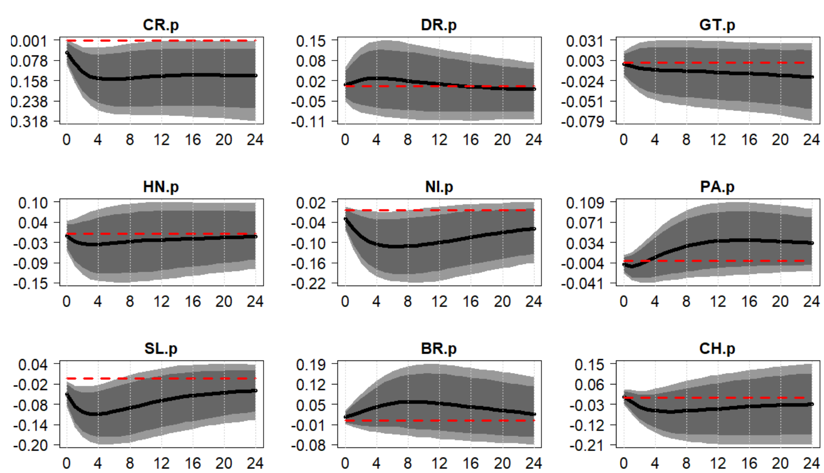
Figure 3: Generalized Impulse Response Functions of a U.S. EPU Shock on Economic Ac-tivity in China, the United States, and the Eurozone



Note: The figure presents generalized impulse response functions (GIRFs) of real economic activity in China, the United States, and the Euro Area following a one-standard-deviation shock to U.S. economic policy uncertainty (EPU). The responses are derived from the BGVAR model estimated using the Minnesota prior under the 21-country specification and based on 20,000 posterior draws. The horizontal axis represents the forecast horizon (in periods), while the vertical axis measures the magnitude of the response. Generalized impulse response functions are invariant to variable ordering and capture the dynamic transmission of uncertainty shocks across major global economies over time.

Source: Authors' estimates.

Figure 4: Generalized Impulse Response Functions of a U.S. EPU Shock on Inflation in Central America



Note: The figure presents generalized impulse response functions (GIRFs) of inflation in Central American countries following a one-standard-deviation shock to U.S. economic policy uncertainty (EPU). The responses are derived from the BGVAR model estimated using the Minnesota prior under the 21-country specification and based on 20,000 posterior draws. The horizontal axis represents the forecast horizon (in periods), while the vertical axis measures the magnitude of the inflation response. Generalized impulse response functions are invariant to variable ordering and capture the dynamic transmission of external uncertainty shocks to domestic inflation dynamics over time.

Source: Authors' estimates.

U.S. uncertainty on domestic activity was analyzed, revealing that economic policy uncertainty tends to reduce industrial activity in the United States. This effect constitutes a clear transmission channel to Central American economies. Consequently, we next quantify the impact of a decline in U.S. economic activity on the economic activity of Central America.

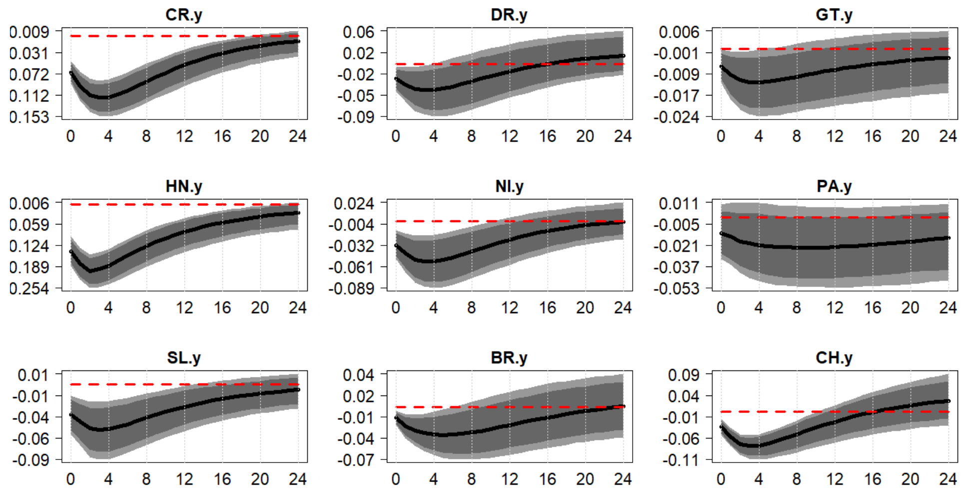
We now analyze how a contraction in U.S. economic activity, previously identified as a global policy uncertainty transmission channel, affects the year-on-year growth rate of industrial production in each Central American economy, as well as in Brazil and Chile. The generalized impulse response functions (GIRFs, see Fig. 5) were estimated using a BGVAR with 21 units, a Minnesota prior with 20,000 draws and burn-in with a Geweke rate of convergence of 9.11%, an endogenous and weakly exogenous lag of order two, and stochastic volatility. In each panel, the black line represents the posterior median response, while the shaded bands (dark gray at 68% and light gray at 90%) allow assessment of statistical significance over a 24-month horizon.

In Costa Rica (CR.y) and Honduras (HN.y), the initial decline in activity following a –1 p.p. shock in the United States reaches approximately –0.10 and –0.18 p.p., respectively, in the first month. These responses are statistically significant (the credibility bands exclude zero). After this period, recovery progresses monotonically, and the bands overlap with the zero line, indicating dissipation of the effect in the medium term. The size of the GIRF from Costa Rica is smaller than Nicaragua (NI.y), El Salvador and Dominican Republic (DR.y) in contrast, experiences a less pronounced and prolonged displacement, the decline reaches almost –0.05 p.p. in month 4, and the bands remain below zero until month 12, reflecting a strong and sustained transmission channel consistent with its high exposure to the U.S. market.

Panama (PA.y), and Guatemala (GT.y) respond with smaller variations ranging from –0.02 to –0.01 p.p. respectively, with credibility intervals including zero from month 1 onward, preventing statistical confirmation of a definitive effect from the U.S. contraction in these countries. This lack of significance may be due to less export-oriented productive structures or internal macroeconomic stabilization mechanisms.

Brazil (BR.y) reacts with a maximum decline of –0.04 p.p. in month 4; the bands exclude zero until month 8, after which the impact gradually dissipates, aligning with the median behavior of large emerging economies. Finally, Chile (CH.y) shows a decline of –0.08 p.p. in month 3, and full recovery around month 18. The observed heterogeneity confirms that the degree of trade openness, sectoral diversification, and macroeconomic institutions determine the intensity and duration of the transmission channel of U.S. shocks to the region. Regarding other regions, the same shock generates a significant contraction in economic activity in the Eurozone and China (see Fig. 6)

Figure 5: Generalized Impulse Response Functions of a U.S. Economic Activity Shock on Central American Economic Activity



Note: The figure presents generalized impulse response functions (GIRFs) of real economic activity in Central American countries following a one-standard-deviation shock to U.S. real economic activity. The responses are derived from the BGVAR model estimated using the Minnesota prior under the 21-country specification and based on 20,000 posterior draws. The horizontal axis represents the forecast horizon (in periods), while the vertical axis measures the magnitude of the response. Generalized impulse response functions are invariant to variable ordering and capture the dynamic transmission of real activity shocks from the United States to Central American economies over time.

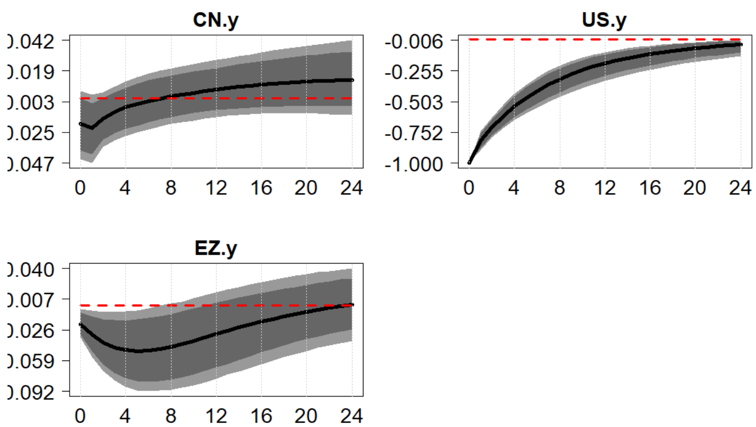
Source: Authors' estimates.

The joint analysis of the GIRFs for the economic policy uncertainty index (EPU) and U.S. industrial activity reveals a dual transmission channel to the economies of Central America, Brazil, and Chile. First, the U.S. EPU acts as a trigger, although its direct effects on some Central American countries are imprecise (their credibility intervals include zero), this uncertainty does generate a statistically significant contraction of U.S. industrial production. Second, this decline in U.S. activity itself propagates more consistently and robustly to the region, as shown by the analyzed GIRFs, countries with closer trade ties and value chain linkages, such as Costa Rica, the Dominican Republic, Honduras, El Salvador, Nicaragua, and Chile, exhibit initial declines of up to –0.20 p.p. in their annual growth rates, with statistical significance for several months.

This compound mechanism implies that, even if the EPU does not directly impact all countries with the same strength, the deterioration of U.S. industrial demand stemming from the EPU constitutes an indirect transmission channel that exacerbates the regional slowdown.

In particular, export-oriented economies with a limited macroeconomic cushioning will suffer not only the effect of global uncertainty but also the reduction in income from external demand. In contrast, Panama, whose responses to U.S. activity shocks are not statistically significant, may have developed productive

Figure 6: Generalized Impulse Response Functions of a U.S. Economic Activity Shock on Economic Activity in China, the United States, and the Eurozone



Note: The figure presents generalized impulse response functions (GIRFs) of real economic activity in China, the United States, and the Euro Area following a one-standard-deviation shock to U.S. real economic activity. The responses are derived from the BGVAR model estimated using the Minnesota prior under the 21-country specification and based on 20,000 posterior draws. The horizontal axis represents the forecast horizon (in periods), while the vertical axis measures the magnitude of the response. Generalized impulse response functions are invariant to variable ordering and capture the dynamic propagation of real activity shocks across major economies over time.

Source: Authors' estimates.

diversification mechanisms or countercyclical policies that mitigate the transmission of external disturbances.

From a policy perspective, these findings suggest that containing the risk of economic policy uncertainty in the United States should be complemented by internal stabilization strategies and market diversification for the most exposed economies. Measures such as promoting regional value chains, strengthening contingency financial networks, and designing targeted fiscal buffers could reduce activity volatility in the face of remote shocks. Moreover, understanding that the indirect channel of the U.S. activity decline tends to be more consistent than the direct EPU shock guides the formulation of macroprudential policies and export promotion strategies toward sectors less sensitive to external demand fluctuations.

The results regarding economic activity and price dynamics are consistent with those found by Largaespada (2021) for Nicaragua and by Giraldo et al. (2025) for the rest of Central America and Latin America. The BGVAR results indicate that the United States can exert a stronger impact on partner economies than what bilateral trade alone would imply, because third-market linkages and deep financial integration amplify the cross-border transmission of business cycles (International Monetary Fund, 2019).

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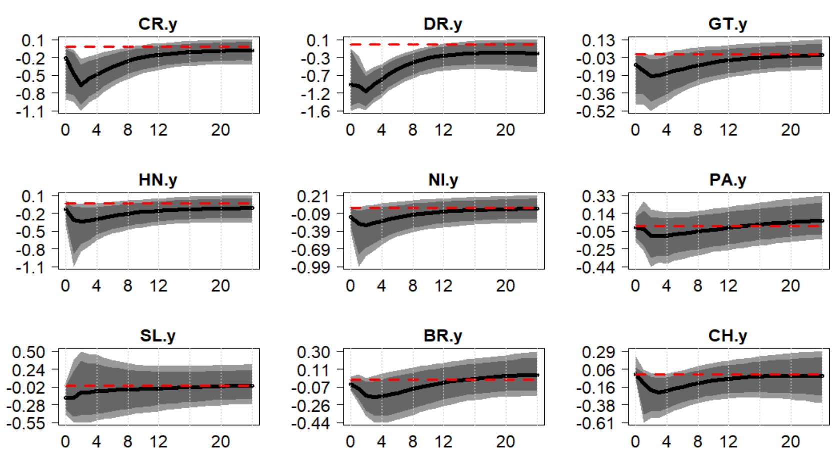
4.1 Robustness via SSVS: Interpretation and Implications

To evaluate the robustness of the inferences obtained under the Minnesota prior, we estimated a version of the BGVAR that incorporates a stochastic search variable selection prior (SSVS). SSVS operates by allowing many coefficients to be effectively zero in the posterior via a narrow–wide mixture; in practice, this favors parsimonious solutions and facilitates the identification of relevant cross-unit coefficients. In contrast to the Minnesota prior, which imposes systematic shrinkage toward low-complexity and highly persistent structures, SSVS relaxes the transnational dependence structure and lets the data determine which relationships survive the selection process.

The results shown in Figure 7 indicate that, overall, the shape and direction of the median responses are consistent with the estimates under Minnesota, suggesting that the identified directional effects are relatively robust. However, the width of the credibility bands changes noticeably for some countries, with Nicaragua and El Salvador being the most notable case, the posterior bands under SSVS are considerably wider than under Minnesota.

This expansion of posterior uncertainty can be interpreted as less information in the data for estimating spillovers to that country, or as the presence of alternative models with similar plausibility that differ in the significance of cross-unit coefficients. Statistically, SSVS allowsfor a greater posterior mass near zero

Figure 7: Generalized Impulse Response Functions of a U.S. Economic Uncertainty Shock on Central American Economic Activity under SSVS prior



Note: The figure presents generalized impulse response functions (GIRFs) of real economic activity in Central American countries following a one-standard-deviation shock to U.S. economic policy uncertainty (EPU). The responses are derived from the BGVAR model estimated using the stochastic search variable selection (SSVS) prior under the 21-country specification and based on 20,000 posterior draws. The horizontal axis represents the forecast horizon (in periods), while the vertical axis measures the magnitude of the response. Generalized impulse response functions are invariant to variable ordering and capture the dynamic transmission of uncertainty shocks under an alternative prior specification.

Source: Authors' estimates.

for certain coefficients, which increases the posteriorvariance of the responses driven by uncertainties about the contagion structure.

Economically, the higher posterior uncertainty for Nicaragua implies caution in attributing a stable transmission channel from the U.S. contraction. While the Minnesota prior may have produced relatively precise inferences due to its shrinkage toward own and persistent effects, SSVS reveals that this precision may depend on assumptions about the density of the coefficient matrix. Therefore, the conclusion that Nicaragua is highly sensitive to a U.S. Economic Policy Uncertainty shock loses statistical robustness when greater parsimony in the interdependence matrix is allowed. For countries whose significance is maintained under both priors, such as Costa Rica and the Dominican Republic in the sample, the conclusions about external impact are more reliable and deserve greater attention in policy design.

The findings indicate that the difference in MN and SSVS prior responses is driven by the way each prior treats cross-equation coefficients. George and McCulloch (1993) show that SSVS uses a mixture prior that lets coefficients that are strongly supported by the likelihood remain sizable while shrinking weak coefficients toward zero, which permits coherent sets of cross-country spillovers to survive prior shrinkage. By contrast, the Minnesota style prior implements systematic global shrinkage toward simple autoregressive benchmarks, which reduces the prior mass on cross effects and tends to dampen dynamic multipliers as documented in the Bayesian VAR literature (Banbura et al., 2010; Doan et al., 1984). Methodologically, this outcome reflects the familiar bias-variance trade-off where SSVS is more data adaptive and can therefore amplify historically concentrated transmission channels while Minnesota yields more conservative and stable impulse responses (Carriero et al., 2013; De Mol et al., 2008). For interpretation of the figures we therefore read larger SSVS GIRF magnitudes as indicative of stronger historical evidence for particular transmission paths rather than as proof that the true structural effect is uniformly larger (D´ees et al., 2007).

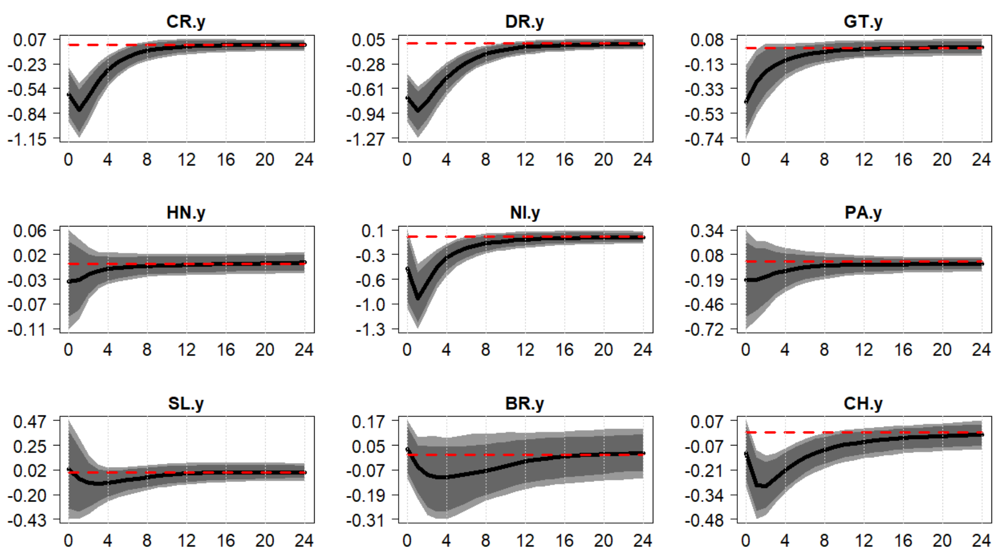
4.2 Response to a U.S. Monetary and Fiscal Policy Uncertainty Shock

The evaluation using the U.S. monetary policy-specific uncertainty index (see Fig. 8) confirms the general direction of the negative effect observed with the aggregate EPU but moderates it in relevant aspects for regional transmission. In the United States, the monetary shock generates an immediate and statistically significant decline in industrial activity, accompanied by more pronounced responses in financial variables and interest rates, highlighting a stronger role of the financial channel relative to the pure demand channel. In Central America, the GIRFs reveal heterogeneity, median responses are in some cases larger than those obtained with the aggregate EPU, and the credibility bands exclude zero over short-and medium-term horizons for economies with stronger financial integration and exposure to capital flows. By contrast, countries whose transmission depends mainly on trade exhibit smaller responses with intervals that often include zero. Overall, the evidence suggests that uncertainty regarding monetary policy does not replace the effect of general political uncertainty but complements it, amplifying spillovers through financial channels and increasing the vulnerability of financially integrated economies. From a policy perspective, these results emphasize the need for macrofinancial instruments and resilience measures against external monetary uncertainty shocks, since such shocks can be transmitted rapidly through asset prices and credit conditions. Evidence of the credit channel as a transmitter of U.S. eco-nomic policy uncertainty to economic activity can be found in Largaespada (2021) for the case of Nicaragua.

The estimated impulse response functions for the U.S. fiscal policy uncertainty index (see Fig. 9) show that the direct effect on the year-on-year activity of most economies is imprecise and statistically indistinguishable from zero. Only the Dominican Republic exhibits a clearly negative and statistically significant median response, persistent in the months following the shock. Guatemala also shows a negative median but with much wider credibility intervals and faster dissipation of the effect, indicating weak evidence of a stable transmission channel.

For the rest of the Central American countries, as well as China and the United States, the reported quantiles (Q5–Q95) cover the zero line across most horizons, preventing a conclusion

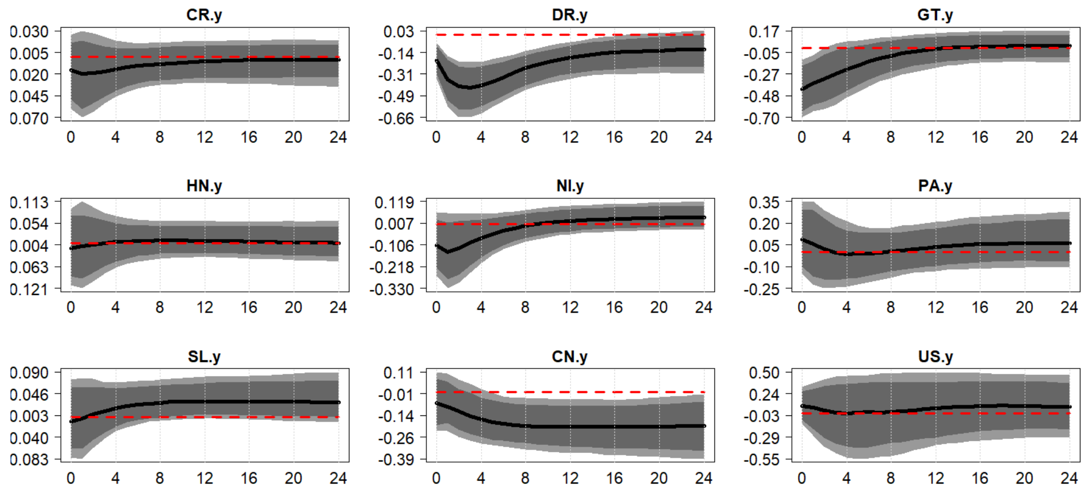
Figure 8: Generalized Impulse Response Functions of a U.S. Monetary Policy Uncertainty Shock on Central American Economic Activity



Note: The figure presents generalized impulse response functions of real economic activity in Central American countries following a one-standard-deviation shock to U.S. monetary policy uncertainty. Responses are computed using the BGVAR model with the Minnesota prior. The horizontal axis represents the forecast horizon in periods, and the vertical axis measures the magnitude of the response. Generalized impulse responses are invariant to variable ordering and reflect the dynamic effects of monetary policy uncertainty shocks transmitted from the United States to Central American economies.

Source: Authors' estimates.

Figure 9: Generalized Impulse Response Functions of a U.S. Fiscal Policy Uncertainty Shock on Central American Economic Activity



Note: The figure reports generalized impulse response functions of real economic activity for Central American countries following a one-standard-deviation shock to U.S. fiscal policy uncertainty. Responses are obtained from the BGVAR model estimated with the Minnesota prior. The horizontal axis represents the forecast horizon in periods, while the vertical axis shows the magnitude of the response. Generalized impulse responses are invariant to variable ordering and capture the dynamic transmission of fiscal uncertainty shocks from the United States to Central American economies.

Source: Authors' estimates.

of a homogeneous or robust impact of U.S. fiscal uncertainty on their economic activity.

From an economic perspective, these results suggest that U.S. fiscal uncertainty operates mainly through heterogeneous and, in many cases, indirect channels. First, when fiscal uncertainty reduces U.S. industrial demand or raises global risk premiums, countries with greater trade exposure, sectoral linkages, or reliance on income flows tied to U.S. demand (industrial exports, tourism, and remittances) are more likely to transmit that shock to real activity. The robust significance in the Dominican Republic may reflect precisely this exposure profile.

Second, the low significance for most GIRFs indicates that economies with higher financial resilience, diversified export destinations, or macro-fiscal buffers do not strongly transmit external fiscal uncertainty in the short term. For policymakers, the lesson is twofold: mitigating transmission requires both internal policies to strengthen fiscal buffers and liquidity networks, as well as strategies to diversify external demand.

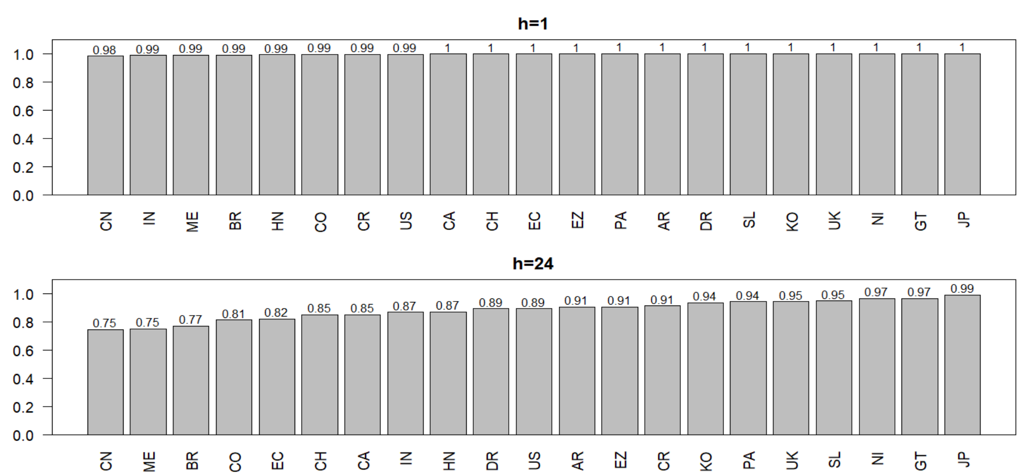
Furthermore, since the indirect channel via U.S. activity contraction has already been identified in previous estimates, it is advisable to integrate the interaction between fiscal uncertainty and the decline in external demand into risk assessments, especially when designing targeted countercyclical measures and macroprudential tools for small open economies.

4.3 GFEVD: Quantifying the Contribution of U.S. EPU (Total and Monetary)

The generalized forecast error variance decompositions (GFEVD) for the model with a Minnesota prior show that, in contemporaneous terms (h = 0), most of the variance in economic activity is explained by internal factors specific to each country (see Fig. 10, top panel), so that the portion attributable to U.S. EPU is generally small at the initial period. However, extending the horizon to h = 24, the GFEVD reveals a heterogeneous increase in the share explained by U.S. EPU across several countries. This suggests that the external effect accumulates and becomes more relevant in the medium term for economies with higher trade or financial exposure to the United States. These patterns are consistent with the literature (Largaespada, 2021; Miescu, 2023). Consequently, the empirical evidence supports the idea of an indirect channel whereby U.S. economic policy uncertainty reduces domestic activity in the United States and subsequently propagates an increasing share of variance to certain recipient economies over longer horizons.

The alternative version using the uncertainty indicator associated exclusively with U.S. monetary policy (see Fig. 11) presents a different contribution configuration; the portion of forecast error explained by monetary EPU tends to concentrate in countries with stronger financial linkages or exposure to capital flows, and in some cases explains a larger fraction of the GFEVD at h = 24 compared to the aggregate

Figure 10: Forecast Error Variance Decomposition of a U.S. Economic Uncertainty Shock on Central American Economic Activity



Note: The gure presents the generalized forecast error variance decomposition (GFEVD) of real economic activity attributable to shocks in U.S. economic policy uncertainty (EPU) across countries included in the BGVAR model. Results are derived from the Minnesota prior under the 21-country specication with 20,000 posterior draws. The top panel (h=1) reports the contemporaneous decomposition and the bottom panel (h=24) the decomposition at a 24-period horizon.

Source: Authors' estimates.

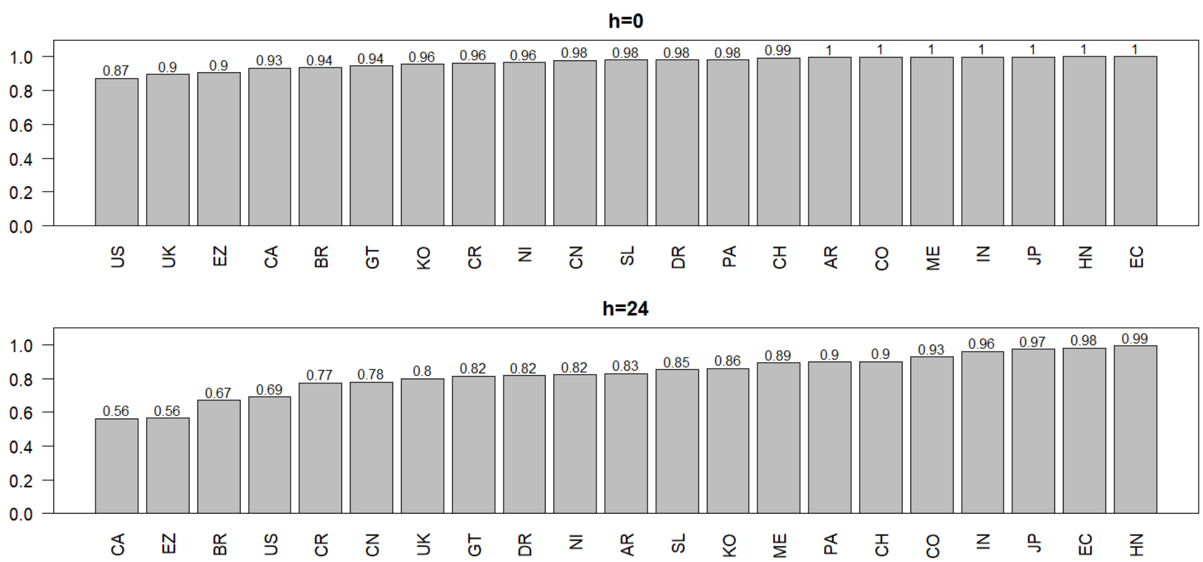
EPU. This finding aligns with studies on GVAR applied to international networks, which show that monetary policy and financial uncertainty shocks transmit prolonged effects through credit and asset price channels (Dées et al., 2007).

From a public policy perspective, differentiating between aggregate EPU and monetary EPU matters, while general political uncertainty may operate through external demand and gradually affect exporting countries, uncertainty about monetary policy can amplify financial vulnerabilities and increase macroeconomic fragility in economies with high cross-border capital flows. To strengthen these conclusions, it is recommended to report, alongside GFEVD, posterior inclusion probabilities (in the case of SSVS) and out-of-sample predictive measures, as suggested by recent empirical applications in the GVAR literature (Dées et al., 2007).

4.4 Comparative GIRF evidence for a US economic policy uncertainty shock

Figure 12 presents a cross-model comparison of generalized impulse responses to an identified US economic policy uncertainty shock for nine countries. Four specifications are shown. Two specifications use a Minnesota prior, and two use stochastic search variable selection. For each prior family, one specification is estimated

Figure 11: Forecast Error Variance Decomposition of a U.S. Monetary Policy Uncertainty Shock on Central American Economic Activity



Note: The gure presents the generalized forecast error variance decomposition (GFEVD) of real economic activity attributable to shocks in U.S. monetary policy uncertainty across countries in the BGVAR model. Results are derived from the Minnesota prior. The gure reports results at forecast horizons h=0 (top panel) and h=24 (bottom panel).

Source: Authors' estimates.

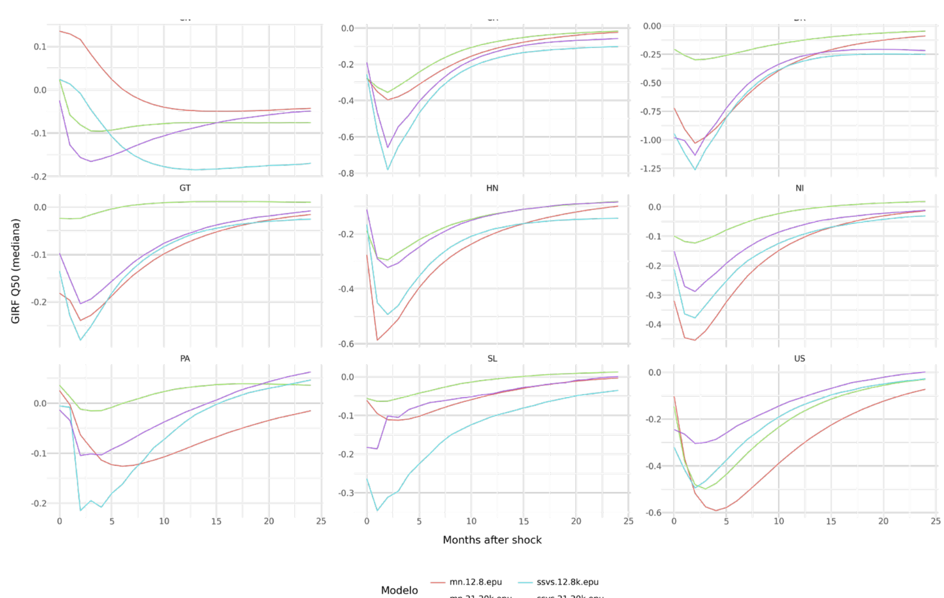
on a 12-country panel with 8,000 draws and thinning 2, and the other on a 21-country panel with 20,000 draws and appropriate burn-in. The plotted series are posterior medians for horizons 0 to 24 months.

The visual evidence highlights three robust empirical regularities and several important caveats.

First, responses are heterogeneous across countries but largely negative in impact. Most panels show an immediate contraction in the activity variable following the uncertainty shock, with the peak negative effect concentrated in the first three months. Recovery begins after the early peak and continues through the one to two-year horizon. This temporal pattern is consistent across priors and cross-sectional sizes, which suggests that a transitory contraction is a common response to elevated US policy uncertainty for the sample considered.

Second, magnitude and shape vary systematically with prior choice and cross sectional dimension. Specifications that employ stochastic search variable selection tend to produce larger and at times more persistent deviations relative to the Minnesota prior. Expanding the cross section from 12 to 21 countries reduces short run volatility and yields smoother posterior medians. These differences are empirically plausible because SSVS places less shrinkage

Figure 12: US Economic Activity Shock GIRF comparative from a 12 country model and a 21-country model, and Minnesota prior and SSVS prior after a US Economic Activity Shock

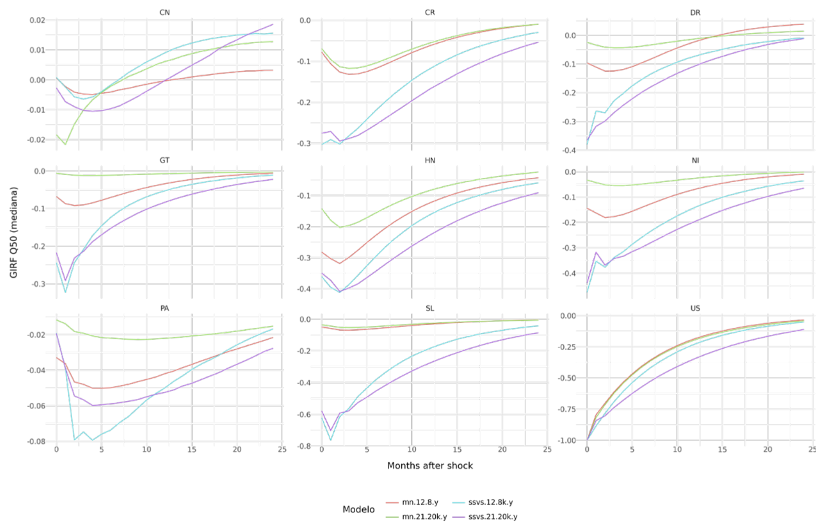


Note: Model labels correspond to the specifications reported in Table \ref{tab:geweke}. Only the models exhibiting the best convergence performance are reported. mn.12.8k.epu denotes the 12-country BGVAR estimated using the Minnesota prior with 8,000 posterior draws; mn.21.20k.epu denotes the 21-country BGVAR estimated using the Minnesota prior with 20,000 posterior draws; ssvs.12.8k.epu denotes the 12-country BGVAR estimated using the stochastic search variable selection (SSVS) prior with 8,000 posterior draws; and ssvs.21.20k.epu denotes the 21-country BGVAR estimated using the SSVS prior with 20,000 posterior draws.

Source: Authors' estimates.

https://doi.org/10.62407/maatfce56

Figure 13: GIRF comparative from a 12 country model and a 21 country model and Min-nesota prior and SSVS prior after a US EPU Shock



Note: Model labels correspond to the specifications reported in Table 4. Only the models exhibiting the best convergence performance are reported. mn.12.8k.epu denotes the 12-country BGVAR estimated using the Minnesota prior with 8,000 posterior draws; mn.21.20k.epu denotes the 21-country BGVAR estimated using the Minnesota prior with 20,000 posterior draws; ssvs.12.8k.epu denotes the 12-country BGVAR estimated using the stochastic search variable selection (SSVS) prior with 8,000 posterior draws; and ssvs.21.20k.epu denotes the 21-country BGVAR estimated using the SSVS prior with 20,000 posterior draws.

Source: Authors' estimates.

on potentially informative coefficients while a larger cross section supplies additional information that stabilizes posterior estimates.

Third, heterogeneity in economic exposure is clear. Some economies display pronounced immediate declines followed by gradual recovery, while others exhibit muted or ambiguous responses. The between-country variation likely reflects differences in trade exposure to the United States, financial openness, commodity links, and domestic policy buffers. The diversity of shapes across panels cautions against single-country generalizations and highlights the value of comparative model ensembles. Convergence diagnostics favor the 21 country specifications due to larger draw counts, which implies greater confidence in their smoothness, but this does not eliminate sensitivity to identification choices.

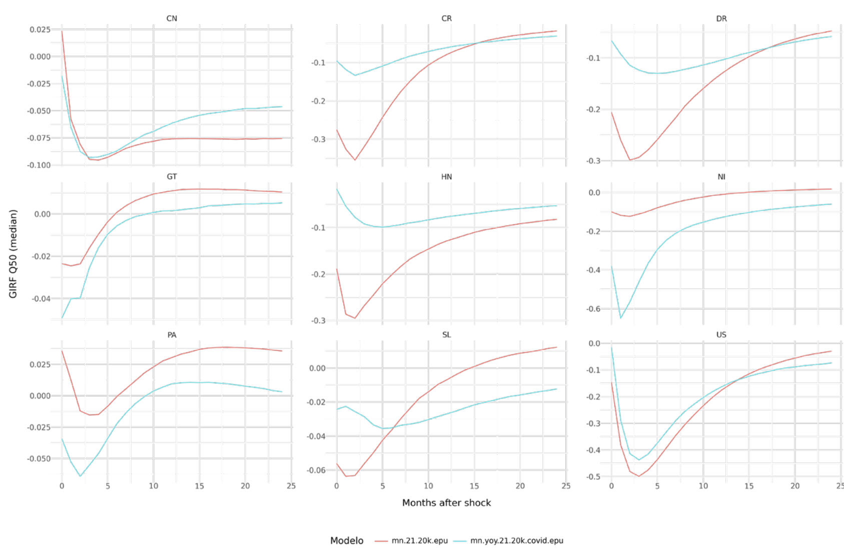
In policy terms, the evidence indicates that episodes of elevated US policy uncertainty transmit negative effects to activity in the region with variable intensity. Policy responses that shorten the duration of real effects through timely stabilization and that reduce external vulnerability may materially attenuate these spillovers.

Similar evidence is found when comparing the US economic activity shock to Central American economies using 2 different priors. The size of the SSVS shock are bigger than with MN prior. However, the reduction of the number of countries included in the model does not affect the size nor the direction and persistence of the GIRF as shown in Figure 13. The size of the response on the economic activity from the SSVS prior for Costa Rica is similar to the results from the International Monetary Fund (2019). Similarly, previous evidence from Giraldo et al. (2025) on the size of the US EPU shock to Costa Rica allign with our results.

4.5 Effect of sample selection and prior choice on GIRFs

The panel in the Figure 14 shows the median generalized impulse response functions for up to twenty four months following a one unit increase in United States Economic Policy Uncertainty as estimated by two versions of the best performing model in our selection exercise. One version is estimated with a monthly sample that ends in December 2024 and the other version is estimated using the identical setup but truncating the sample in December 2019 so that the Covid period is excluded. Both estimates are computed from a 21 country Bayesian global VAR with two lags and with 20,000 draws after burn in using the Minnesota type prior. The plot makes two points that are important for interpretation and robustness assessment. First the size and persistence of the estimated responses to a US EPU shock differ across the two samples for some countries. For several Central American economies the responses estimated on the full sample are larger in absolute value and decay more slowly than the responses estimated on the pre-COVID sample. Second, the cross-country pattern of signs and timing remains broadly similar across the two samples for most economies but not for all.

Figure 14: GIRF comparative from a truncated sample model with Minnesota prior and 21 economies



mn.21.20k.epu: is the model with 21 countries, run 20,000 draws, Minnesota prior and entire sample 2003-2024; mn.yoy.21.20k.covid.epu: is the model with 21 countries, run 20,000 draws, Minnesota prior and shorter sample 2003-2019.

Source: Authors' estimates.

Major global episodes such as the 2008 financial crisis and the COVID-19 shock produce extreme observations and abrupt changes in second moments. When such episodes are included in the estimation sample, the posterior distribution of coefficients and of volatility parameters responds to the additional evidence. If the sample includes COVID, there is an increase in evidence of transient but very large co-movements among countries and in global volatility. Bayesian estimators that permit time variation and that do not enforce very strong shrinkage therefore shift posterior mass towards parameter constellations that imply stronger cross-country transmission and larger responses to a US uncertainty shock. By contrast, if the COVID window is excluded, the estimator places more weight on the pre-COVID dynamics that are typically lower in volatility and in cross-border synchronisation.

The net effect is that impulse responses estimated on the truncated sample are often smaller and less persistent. This point is emphasized in the literature on structural stability in GVAR-style frameworks where co-breaking and changes in error variances affect estimated transmission patterns, and where robust inference on impulse responses benefits from explicit consideration of sample splits or robust standard errors (D´ees et al., 2007; Primiceri, 2005).

Finally, there is an empirical interpretation that links the statistical facts to economic channels. The Covid period featured synchronized global demand and supply disruptions, large policy actions, and acute changes in trade and financial linkages. For small open economies with strong trade exposure to the United States, such episodes increase the elasticity of domestic activity to foreign uncertainty.

4.6 Discussion and Policy Implications

Our results confirm that shocks in U.S. economic policy uncertainty (EPU) generate significant macroeconomic effects on Central American economies, consistent with evidence for other small open markets. In particular, uncertainty shocks are transmitted intensely through the trade channel. When the U.S. EPU rises, it often coincides with lower U.S. industrial demand, which dampens Central American exports. This aligns with Swiston (2010), who identifies demand fluctuations in advanced economies as the main channel affecting growth in Central America, and with Coronado et al. (2020), who show that increases in U.S. EPU tend to cause depreciations in Latin American currencies and weaken output in economies with higher exposure to U.S. trade. Our BGVAR-based GIRFs indicate a dual impact, EPU shocks reduce U.S. activity while simultaneously raising uncertainty, exposing the region to a combined effect of lower external demand and tighter financial conditions.

Recent studies on U.S. uncertainty in Latin America support this pattern; for instance, Giraldo et al. (2025) document that real and financial uncertainty shocks from the U.S. strongly depress Latin American GDP, with effects that can be more persistent than conventional policy or monetary shocks. We also find that uncertainty regarding U.S. monetary policy amplifies these effects, consistent with Lastauskas et al. (2023), who show that a Fed rate normalization accompanied by greater policy volatility produces sharper declines in output and inflation than a rate hike in the absence of increased uncertainty. In summary, the heterogeneous responses observed in Central America are consistent with the notion that small open economies "import" external uncertainty (Giraldo et al., 2025; Luk et al., 2017) and that EPU shocks tend to induce depreciations, capital outflows, and declines in activity in emerging markets (Bhattarai et al., 2018; Coronado et al., 2020).

It is important to note that the BGVAR architecture constructs country equations conditional on trade-weighted foreign aggregates, so a foreign shock is decomposed and allocated across many bilateral exposure channels rather than being concentrated in a single direct foreign regressor as in a single-country VAR. This mechanical dispersal reduces the local marginal effect of a foreign shock on any given domestic variable because part of the foreign disturbance is explained by other countries' endogenous responses and by common global factors rather than by a single direct transmission path. The econometric consequence is a smaller marginal GIRF in the conditional multi-country setup relative to a smaller dimensional VAR where the same foreign shock is modeled as a directly exogenous driver.

Cross-country results highlight differing degrees of vulnerability among Central American economies. As emphasized by Luk et al. (2017), small economies with high financial integration absorb a considerable fraction of global uncertainty. In practice, the more open or highly leveraged countries, such as those oriented toward manufacturing exports or heavily dependent on remittances, exhibit stronger responses to U.S. uncertainty shocks, whereas economies with greater diversification or macroeconomic buffers show more resilience. In line with Bhattarai et al. (2018), some Latin American markets experienced more pronounced output declines and capital outflows following U.S. EPU shocks, depending partly on the speed and nature of policy responses and the degree of financial integration. In our sample, countries such as Honduras or Costa Rica, with relatively high trade-to-GDP ratios, exhibit more pronounced output fluctuations, while Panama, benefiting from an international banking sector with substantial buffers, displays attenuated real effects. Panamanian authorities report that in 2025, banks maintained robust capital and liquidity ratios above regulatory minima, contributing to resilience amid a more uncertain external environment (Superintendencia de Bancos de Panama, 2025).

This evidence aligns with Swiston (2010), who argue that financial resilience and export diversification mitigate the transmission of external shocks; additionally, studies such as Carvalho Filho and Ng (2023) show that counter-cyclical macroprudential tools reduce credit volatility and its transmission to the real economy. Consequently, policy responses should be calibrated to structural heterogeneity, uniform measures may be ineffective, and the intensity of monetary, fiscal, and macroprudential stabilizers should be adjusted according to each country's exposure and vulnerability profile.These conclusions have clear and complementary public policy implications. First, it is essential to strengthen macrofinancial buffers and deploy macroprudential instruments, as dynamic provisions, countercyclical capital requirements, and other tools reduce the transmission of external shocks to credit and financing costs (Carvalho Filho & Ng, 2023; Imam et al., 2012). Second, trade and export destination diversification diminishes dependence on U.S. industrial cycles and, therefore, vulnerability to EPU episodes; similarly, the IMF documents that greater productive diversification contributes to higher sustained growth rates and lower volatility in developing economies (International Monetary Fund, 2014).

Third, maintaining credible macroeconomic frameworks with countercyclical space, prudent fiscal policies, adequate reserves, and price-stability-oriented monetary regimes provides maneuvering room for timely responses without compromising sustainability. Finally, leveraging regional and international support mechanisms, such as precautionary lines, liquidity networks, and disaster insurance, should complement national policies to reduce systemic risks.

Overall, the BGVAR findings support a broad strategy strengthen macroprudential tools to contain financial channels, promote market diversification, and consolidate fiscal and monetary frameworks, all calibrated to the observed heterogeneity of exposures in the region.

5. Conclusion

This study addressed the research question of how U.S. economic policy uncertainty influences growth and inflation dynamics in Central American countries, using a Bayesian Global Vector Autoregressive (BGVAR) model. By defining a regional trade-weighting matrix and estimating the model with Minnesota priors, we captured both direct effects of the economic policy uncertainty (EPU) index and the indirect channel through the contraction of U.S. industrial activity.

The main results reveal heterogeneous transmission. Costa Rica, Honduras, El Salvador, and Nicaragua experience initial declines in their annual growth that are statistically significant for several months, while Panama, and Guatemala show imprecise or weakly significant responses. Furthermore, the EPU shock first generates a moderate but significant contraction in U.S. production, which then serves as a more consistent indirect channel than the direct effect of uncertainty on the region.

The findings indicate that U.S. political uncertainty affects regional activity through a composite mechanism. EPU reduces U.S. industrial production, and this contraction acts as an indirect channel that transmits heterogeneous shocks to Central America, Brazil, and Chile, depending on the strength of trade and financial linkages; however, the magnitude and statistical significance of these transmissions vary systematically across countries. The GIRFs and GFEVDs show that, contemporaneously, most of the volatility in activity is explained by domestic factors, but over the medium term, the share attributable to U.S. EPU increases in economies with high exposure to external demand and financial flows, whereas for countries with external diversification or macrofinancial buffers, the effect is imprecise.

Robustness checks indicate that directional conclusions hold when using alternative priors and index versions, and in pre-covid sample, although posterior uncertainty increases in selection specifications such as SSVS, which requires caution in asserting high sensitivity in specific cases. In terms of channels, the estimates suggest the simultaneous existence of an external demand channel affecting industrial exports, tourism, and remittances, and a financial channel operating through asset prices, credit conditions, and risk premia, the latter being relatively more pronounced when using the monetary-policy-specific EPU. Moreover, implementing macroprudential tools and monitoring cross-border risks should be prioritized in economies with higher financial exposure.

These findings have immediate policy relevance, central banks and policymakers in Central America should complement external uncertainty monitoring with internal countercycli-cal stabilization measures, including market diversification, strengthening fiscal buffers, and promoting regional value chains. Such strategies would help mitigate volatility induced by remote shocks and reduce the risk of pronounced economic slowdowns.

Key limitations of the study include the temporal horizon (monthly data from January 2003 to December 2024), potential sensitivity to the Bayesian prior specification, and the absence of alternative uncertainty measures, such as those based on financial markets, survey expectations, or disaggregated monetary and fiscal uncertainty, which could enrich shock identification. Additionally, the use of a single fixed weighting matrix may underestimate structural changes in trade patterns.

Future research could extend this analysis by incorporating variables that better capture the idiosyncratic behavior of local economic activity, exploring nonlinear or asymmetric effects in uncertainty transmission, employing news-based or implied volatility indicators, and analyzing subsamples by sector or institutional development. This line of work will deepen the understanding of international contagion mechanisms and refine macroeconomic policy tools for small open economies.

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