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Impact of Globalization on Macroeconomic Dynamics Using a Time-varying Bayesian VAR

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Abstract

In the last two decades, advances in globalization have evolved remarkably and countries have become more integrated with the entire world. The implications of this process have attracted interest of researchers and monetary policy authorities. This paper provides an assessment of the impact of globalization on macroeconomic dynamics at the level of four representative Central Eastern European countries (Czechia, Hungary, Poland and Romania) between 2003Q2 and 2022Q4. The method proposed in this study is a time-varying vector autoregressive model with stochastic volatility estimated using the Metropolis-Hastings algorithm through Bayesian inference. The results of the impulse response functions suggest a slight decrease in inflation and an increase in economic activity for some of the analysed countries, after a positive shock in the degree of trade openness, a proxy used for the advance in globalization.

Keywords: globalization, time-varying vector autoregressive, Bayesian estimation, inflation

JEL Classification: C11, C32, C52

1. Introduction

Globalization has become a widely used concept in research papers in the last two decades, given that global processes that influence countries through deepening international linkages and interactions determine a more significant part of economic development. The globalization hypothesis, or the "global slack" hypothesis, states that the internationalization of goods and financial markets through global value chains and trade openness increases the importance of international factors over domestic ones in explaining the dynamics of macroeconomic outcomes such as business cycles or inflation. Moreover, increased competitiveness has the potential to spur productivity growth between interconnected economies and, as a consequence, reduce inflation. This aspect is essential for monetary policy authorities to form their decisions, given that the widely used econometric models for open economies assign a crucial role to external sources in determining domestic macroeconomic dynamics.

However, the main research question in this direction is whether and to what extent this recent and accelerated process of economic globalization has been reflected over time in macroeconomic developments. This aspect has become more critical in recent periods, which are marked by heightened uncertainty and risks related to the pandemic crisis, energy crisis, hikes in consumer prices and the Russo-Ukrainian War, which have affected Central Eastern European (CEE) countries directly and indirectly through commercial links. Nowadays, the main challenge comes from the difficulty of distinguishing, even more so in times marked by uncertainty, between the impact of different shocks (demand, supply, monetary policy, uncertainty, energy or external shocks). While there is a consensus regarding the importance of globalization as a real phenomenon, there is less agreement on its empirical implications for the economy.

To address this issue, we use an approach based on a vector autoregressive model with time-varying parameters and stochastic volatility (TVP-SV-VAR). We evaluate using impulse response functions at each moment in time, between 2003Q2 and 2022Q4, the effect of a positive trade openness shock, being equivalent to an increase in globalization, on the macroeconomic dynamics (i.e., economic growth, core inflation and exchange rate of domestic currency relative to the euro). We follow the OECD (2010) recommendation to define globalization through the ratio of international trade to GDP. According to economic theory, globalization should offer notable benefits to worldwide economies through increased competition (which leads to technological innovation), trade linkages and foreign direct investment. Moreover, this process may contribute to income inequality between countries. All of these improve efficiency and increase economic output. However, in times of uncertainty and crisis, globalization can increase countries' vulnerabilities to adverse shocks, hindering economic development. In theory, regarding the impact on inflation, developments in trade integration and participation for low-cost producers in global value chains are expected to have, in general, a disinflationary effect.

The reason behind the choice of this class of models derives from the possibility of using the Bayesian econometric techniques, which permit the estimation of parsimonious models with a small number of variables and lags. Moreover, the existing literature that tests for predictive accuracy proves that TVP models outperform classical VAR in terms of inflation forecasts (see, e.g., Bekiros, 2014). This approach provides easily and intuitively interpreted results from an economic point of view.

The motivation for this analysis of four emerging CEE countries (i.e., Czechia, Hungary, Poland and Romania) derives from the fact that they are small open economies and non-euro area countries that aim to join the currency union after fulfilling the convergence criteria. They have similar characteristics, such as inflation targeting and floating exchange rate regimes. Therefore, they are susceptible to suffering from the impact of globalization or trade integration that might affect euro area requirements. Recently, in the context of heightened uncertainty and increasing risks associated with the pandemic, energy crises or the war in Ukraine, the urgency for monetary policy to stabilize inflation and prevent recession has become extremely important. The results presented in this paper suggest a slight decrease in inflation and an increase in economic activity after a rise in the degree of trade openness.

The remainder of the paper is organized as follows. The following section presents a review of literature. Sections 3 and 4 are dedicated to the data and empirical results of the economic framework using a TVP-SV-VAR approach. The last section summarizes the key ideas.

2. Literature Review

In the 1990s, global inflation fell significantly, and it became more stable in the same decade with the hyper-globalization phase, which led to questions about the linkages between these two processes. In addition to the growing credibility and efforts by central banks worldwide to maintain price stability, some theories related to other factors emerged. According to Miskin (2009), globalization might help countries reduce inflation in some specific ways. Thus, globalization promotes the benefits of achieving price stability by stimulating interactions between central banks from different countries, governments, public institutions, academia and the public. According to Ferreira de Mendoça et al. (2020), inflation-targeting countries with high levels of development and low risk of political pressures perform better in terms of monetary policy efficiency.

Existing empirical evidence reveals that the effects of globalization at the macroeconomic level are mixed. Many studies concentrate on inflation, while the evidence related to the output or other macroeconomic indicators is limited. The literature introduced the "global slack hypothesis", which claims the progressive influence of global factors at the expense of domestic ones to explain economic dynamics. To summarize some of these pieces of evidence, Borio and Filardo (2007) concluded that the role of external factors has increased over time in addition to the growing credibility of monetary policy. The process of integration into the world economy has gained momentum and global factors seem to undermine the role of domestic ones in explaining the economic slack. Even if inflation results from global factors, it does not imply that monetary policy should remain on the sidelines.

Additionally, Benigno and Faia (2016) added that competition on international markets puts pressure on firms, making their pricing decision process more dependent on foreign factors. In this sense, the authors presented two relevant channels that explain these pronounced linkages between domestic inflation and global factors. First is the rise in the impact of the import prices on the headline price index due to the increase of foreign competitors on markets. Secondly, given the intensification of competition, there is also an increase in the dependence of the pricing strategies of domestic firms on foreign firms. Therefore, the degree of exchange rate pass-through is directly proportional to changes in the number of foreign competitors. Guilloux-Nefussi (2020) added that another force may act in the opposite direction in response to globalization: the reallocation effect. The increase in the concentration of the largest firms and their strategies to set the markup reduce the responsiveness of inflation to marginal cost shocks. As regards the pass-through of costs into prices, the process implies an imperfect pass-through, a feature also discussed in Ravn et al. (2010, 2012).

Ihrig et al. (2010) tested various predictions of the direct impact of globalization on inflation and provided little support for this theory. According to their results, there is no evidence regarding an increase over time in the responsiveness of inflation. However, the volatility of GDP has fallen in the last decades by more than the volatility of domestic demand, so net exports act as a buffer to attenuate fluctuations in domestic demand. Consequently, globalization could stabilize real GDP growth and inflation. Other studies, such as Woodford (2007) and Gali (2010) have admitted that external developments in a more integrated economy are regarded as sources of economic disturbances but have not detected a significant relationship between globalization and inflation. Gao et al. (2023) investigated the effect of globalization on 15 emerging markets using a non-linear smooth transition regression of a Philips curve. Their measure of the foreign output gap incorporated in a Philips curve as a measure of globalization is not a key determinant for inflation.

The effects of globalization on the output in emerging economies were studied by Acheampong et al. (2021). The authors suggested that globalization improves energy efficiency and economic growth in the short run. However, in the long term, there is a need to implement policies in sectors that have comparative advantage in potential productivity growth. The technological effect of globalization needs to be transferred to productive sectors where technological innovations are limited.

Time-varying effects from a VAR perspective have been examined in papers such as Bianchi and Civelli (2015); the analysis has only focused on the effects of foreign output gap as a measure of globalization. Their research concentrated on eighteen countries, emerging and advanced economies, but all were non-CEE countries. The results suggest that global slack affects the inflation dynamics in most countries analysed, but the effects do not become stronger over time. Our model provides a similar methodological framework, but we contribute to the existing literature by extending the framework to CEE countries and assessing the implications of globalization over economic activity. To the best of our knowledge, a similar approach to CEE countries has not been carried out in other research studies.

Many models proposed in the literature (see, among others, Bianchi and Civelli, 2015; Borio and Filardo, 2007; Forbes, 2019) focus on domestic core inflation to control for international developments. Recently, Forbes (2019) explained the role of globalization in inflation puzzles. If the inflation is driven by global factors (commodity prices, exchange rates, oil prices or global value chains), the central bank's ability to stabilize it may be limited and further adjustments to interest rates are required. In this paper, we use core inflation as a measure of price dynamics rather than headline inflation to remove the effect of high and persistent energy prices in recent periods.

Economic theory suggests that the globalization effect could be equivalent to an expansion of openness to trade, the measure used to carry out this analysis. Some recent multidimensional approaches suggest that the KOF globalization index, constructed by Dreher et al. (2009), can measure globalization along economic, social and political dimensions for almost every country in the world (see Aluko and Opuku, 2022). However, it is important to note that the frequency of the data is annual, and questionnaire surveys used to construct this measure are conducted one or two years prior to release. Moreover, this indicator might not be appropriate for empirical studies due to limitations related to data availability, which can significantly reduce the accuracy of classical econometric methods.

3. Time-varying VAR Model with Stochastic Volatility

The vector autoregressive model with time-varying parameters and stochastic volatility proposed in this paper follows the seminal works of Primiceri (2005), Benati and Mumtaz (2007) and Blake and Mumtaz (2017). This methodology with stochastic volatility is frequently used to study the time-varying impact of structural shocks in the economy.

We consider the VAR model with time-varying parameters having the following specifications:

$$\boldsymbol{Y}_t = \boldsymbol{c}_t + \sum_{j=1}^{P} \boldsymbol{B}_{j,t} \boldsymbol{Y}_{t-j} + \boldsymbol{\upsilon}_t, \qquad var(\boldsymbol{\upsilon}_t) = \boldsymbol{R}_t \tag{1}$$

$$\boldsymbol{\beta}_t = \{\boldsymbol{c}_t, \boldsymbol{B}_{1,t} \dots \boldsymbol{B}_{P,t}\} \tag{2}$$

$$\boldsymbol{\beta}_t = \boldsymbol{\mu} + \boldsymbol{F}\boldsymbol{\beta}_{t-1} + \boldsymbol{e}_t, \qquad var(\boldsymbol{e}_t) = \boldsymbol{Q} \tag{3}$$

We denote with $\boldsymbol{Y_t}$ the T×1 vector of observed variables, $\boldsymbol{Y_{t-j}}$ is a T×1 vector containing the regressors and $\boldsymbol{\beta_t}$ is the vector of the time-varying parameters: the constant term c_t with the dimension Tx1 and the coefficients $\boldsymbol{B_{1,t}} \dots \boldsymbol{B_{P,t}}$, where P denotes the lag length included in the model, which in this case is equal to one lag. The time-varying parameters are defined in a generalized form in Equation (3). However, we follow the assumption that $\boldsymbol{\mu} = 0$ and $\boldsymbol{F} = 1$, which is frequently used in practice (see Primiceri, 2005; and Nakajima, 2011).

The variance-covariance matrix of the error terms defined by $\boldsymbol{\upsilon_t}$ is $\boldsymbol{R_t}$, which varies over time. Regarding the specification of this matrix, most of the studies in the literature simply assume the following structure for $\boldsymbol{R_t}$, so the factorization from the below equation ensures a positive definite matrix.

$$\boldsymbol{R}_t = \boldsymbol{A}_t^{-1}\boldsymbol{H}_t\boldsymbol{A}_t^{-1'} \tag{4}$$

where $\boldsymbol{A_t}$ is a lower triangular matrix having the elements $\boldsymbol{a_{ij,t}}$, and $\boldsymbol{H_t}$ is a diagonal matrix with the terms $\boldsymbol{h_{i,t}}$ on the main diagonal. For a vector autoregressive with four time series, we define:

$$\boldsymbol{A}_t = \begin{pmatrix} 1 & 0 & 0 & 0 \\ a_{12,t} & 1 & 0 & 0 \\ a_{13,t} & a_{23,t} & 1 & 0 \\ a_{14,t} & a_{24,t} & a_{34,t} & 1 \end{pmatrix}, \quad \boldsymbol{H}_t = \begin{pmatrix} h_{1,t} & 0 & 0 & 0 \\ 0 & h_{2,t} & 0 & 0 \\ 0 & 0 & h_{3,t} & 0 \\ 0 & 0 & 0 & h_{4,t} \end{pmatrix}, \tag{5}$$

where $\boldsymbol{a}_{ij,t} = \boldsymbol{a}_{ij,t-1} + \boldsymbol{u}_{i,t}, \qquad var(\boldsymbol{u}_{i,t}) = \boldsymbol{D}$

$$\mathbf{ln}\boldsymbol{h}_{i,t} = \mathbf{ln}\boldsymbol{h}_{i,t-1} + z_{i,t}, \qquad var(z_{i,t}) = \boldsymbol{G} \tag{6}$$

$(\forall)\ i = \overline{1,4}$. The random walk specification for both $\boldsymbol{a_{ij,t}}$ and $\mathbf{ln}\boldsymbol{h_{i,t}}$ has the advantage to present persistent structural changes in the variance of residuals and the logarithmic form of $\boldsymbol{h_{i,t}}$ diagonal term ensures a positive definite matrix $\boldsymbol{H_t}$. Therefore, the model contains two different sets of time-varying parameters: the first are $\boldsymbol{\beta_t}$, $\boldsymbol{a_{ij,t}}$ and secondly, we derive the stochastic volatility of the diagonal elements $\boldsymbol{h_{i,t}}$, respectively. To solve the system of equations, we apply the algorithm developed by Carter and Kohn (1994) to estimate the parameters $\boldsymbol{\beta_t}$ and $\boldsymbol{a_{ij,t}}$ and the independent Metropolis-Hastings algorithm to assess the stochastic volatility of structural shocks. For details, see Appendix 1.

For the identification of the structural shocks, we use the Cholesky factorization recursive identification. The specification is preferred in the literature because the TVP-VAR model has a sizeable number of parameters to estimate; therefore, we may want to avoid the over-identification of the model by imposing additional restrictions (Nakajima, 2011). From an economic interpretation, this factorization is essential for the order in which the variables are included in the model. The short-run restrictions of the lower triangular identification matrix (i.e., zero elements above the main diagonal) constrain variables to be introduced in the following order. First, we include "slow-moving" variables and then "fast-moving" ones since in this identification scheme "slow-moving" variables do not contemporaneously affect "fast-moving" ones.

Thus, the relation between the structural shocks and the residuals is as follows:

$$\boldsymbol{A_t v_t} = \boldsymbol{\varepsilon_t} \tag{7}$$

where the variance of the structural shocks is assumed to be $var(\boldsymbol{\varepsilon_t}) = \boldsymbol{H_t}$. For a model with four time series of data, we have the following set of equations:

$$\begin{pmatrix} 1 & 0 & 0 & 0 \\ a_{12,t} & 1 & 0 & 0 \\ a_{13,t} & a_{23,t} & 1 & 0 \\ a_{14,t} & a_{24,t} & a_{34,t} & 1 \end{pmatrix} \begin{pmatrix} v_{1,t} \\ v_{2,t} \\ v_{3,t} \\ v_{4,t} \end{pmatrix} = \begin{pmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \\ \varepsilon_{3,t} \\ \varepsilon_{4,t} \end{pmatrix} \tag{8}$$

or, in more detail:

$$v_{1,t} = \varepsilon_{1,t}$$

$$v_{2,t} = -a_{12,t}v_{1,t} + \varepsilon_{2,t}$$

$$v_{3,t} = -a_{13,t}v_{1,t} - a_{23,t}v_{2,t} + \varepsilon_{3,t}$$

$$v_{4,t} = -a_{14,t}v_{1,t} - a_{24,t}v_{2,t} - a_{34,t}v_{3,t} + \varepsilon_{4,t} \tag{9}$$

where

$$var(\boldsymbol{\varepsilon_{i,t}}) = \boldsymbol{h_{i,t}}$$

and $a_{21,t} = a_{21,t-1} + u_{1t}$, $\qquad var\left(u_{1t}\right) = \boldsymbol{D}_1$

$$\begin{pmatrix} a_{13,t} \\ a_{23,t} \end{pmatrix} = \begin{pmatrix} a_{13,t-1} \\ a_{23,t-1} \end{pmatrix} + \begin{pmatrix} u_{2t} \\ u_{3t} \end{pmatrix}, \qquad var\left(\begin{pmatrix} u_{2t} \\ u_{3t} \end{pmatrix}\right) = \boldsymbol{D}_2$$

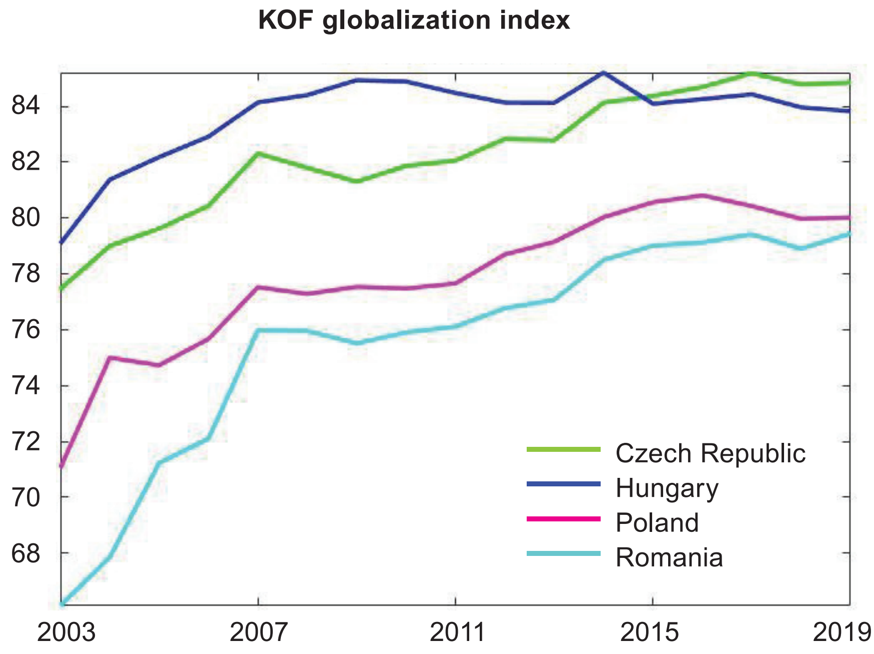
$$\begin{pmatrix} a_{14,t} \\ a_{24,t} \\ a_{34,t} \end{pmatrix} = \begin{pmatrix} a_{14,t-1} \\ a_{24,t-1} \\ a_{34,t-1} \end{pmatrix} + \begin{pmatrix} u_{2t} \\ u_{3t} \\ u_{4t} \end{pmatrix}, \qquad var\left(\begin{pmatrix} u_{2t} \\ u_{3t} \\ u_{4t} \end{pmatrix}\right) = \boldsymbol{D}_3 \tag{10}$$

where $\boldsymbol{a}_{ij,t}$ are time-varying coefficients for the equations of the vector autoregression. We estimate all three types of time-varying parameters with the Gibbs sampling algorithm (i.e., Markov chain simulator) and the Metropolis-Hastings algorithm. To examine the stability of parameters, we check the convergence of the algorithm using a diagnostic test. This procedure is based on the recursive means of the Metropolis-Hastings draws, calculated at every 20 draws for $\boldsymbol{\beta}_t$, $\boldsymbol{a}_{ij,t}$ and $\boldsymbol{h}_{ij,t}$ This is recommended for models with time-varying parameters by Mumtaz and Sunder-Plasssman (2010), Mumtaz and Zanetti (2013), Mumtaz et al. (2011). We represent the results of the convergence of the Metropolis-Hastings algorithm in Appendix 4. The small fluctuations on recursive means suggest convergence for $\boldsymbol{\beta}_t$, and $\boldsymbol{a}_{ij,t}$, but indicate some variation (only in some draws of the algorithm) in the means of $\boldsymbol{h}_{ij,t}$,as a result of the extraordinary COVID-19 shock, which increased the volatility in the system. To assess the impact of the developments in globalization over the last two decades, we compile the impulse response functions to positive trade openness shocks. A description of this procedure is presented in Appendix 2.

4. Data

For this paper, we attain the analysis of the impact of globalization on the macroeconomic dynamics by studying the effects of an increase in trade openness on key macroeconomic dynamics (i.e., economic growth, inflation and exchange rate of domestic currency relative to the euro). As mentioned before, even if the annual KOF globalization index is not appropriate for our analysis with VAR models, we plot it for illustrative purposes in Figure 1. Thus, we can see a visible tendency of globalization in these countries in recent periods.

Figure 1: Evolution of KOF globalization index between 2003 and 2019 (last available data)



Source: author's calculations

The Eurostat database provides all the indicators covering the period between 2003Q2 and 2022Q4. The first indicator in the model is the quarterly variation in the trade openness indicator, calculated as the sum of exports and imports divided by GDP (see Eurostat data code namq_10_gdp). We add quarterly economic growth calculated as a percentage change in real GDP measured in chain-linked volumes with the base year 2015, million units of national currency (see also data code namq_10_gdp). The quarterly inflation rate is calculated by averaging monthly data from consumer prices at constant taxes (HICP), excluding energy, food, alcohol and tobacco (see data code prc_hicp_cind). The last indicator is the quarterly dynamics of the bilateral exchange rate of the national currency relative to the euro for each of the selected group of CEE countries: Czechia, Hungary, Poland and Romania (data code ert_bil_eur_q). The inflation measure described above was preferred to present core inflation, which is more stable relative to headline inflation, and which does not suffer from volatile prices (e.g., energy prices). All the indicators are represented as percentage changes and the time series are stationary (the results of stationarity tests are presented in Table A1, Appendix 3).

In Figure 2, we represent the evolution of the dataset. For all four analysed countries, one can note the volatile developments surrounding the international financial crisis and the COVID-19 pandemic. During periods of high volatility, contagion and regime switches might affect developing countries. After the outbreak of the pandemic crisis and the lockdown measures, the countries experienced not only a significant economic contraction but also partial and temporary disturbances in trade, which in turn caused global chain bottlenecks. In terms of inflation, one can note the tendency to reduce inflation and the persistence at low rates. However, this trend was interrupted in the last year, when we witnessed a sharp increase in inflation induced by the COVID-19 pandemic, global supply chain disruptions and the war in Ukraine. For the exchange rate, the period after the financial crisis exhibits a relative stabilization of domestic currencies. Significant exchange rate fluctuations occurred during the global financial crisis because the external environment influenced the domestic economy and price levels.

Figure 2: Evolution of data series in 2003Q2–2022Q4

Figure 2a: Czechia

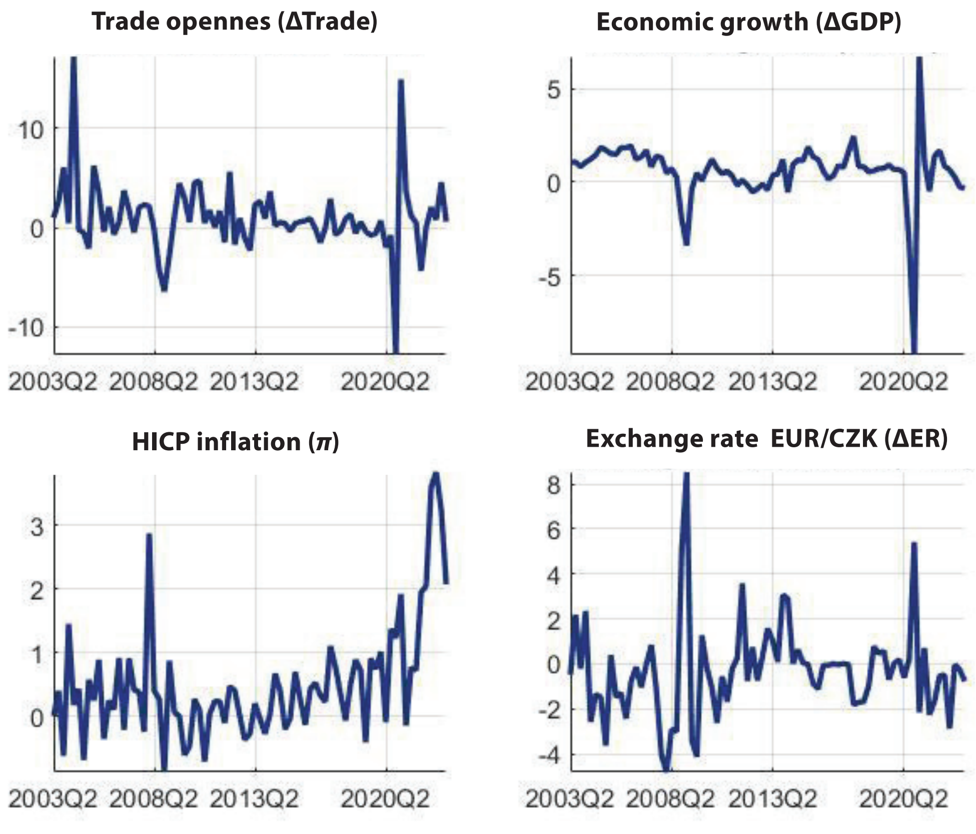


Figure 2b: Hungary

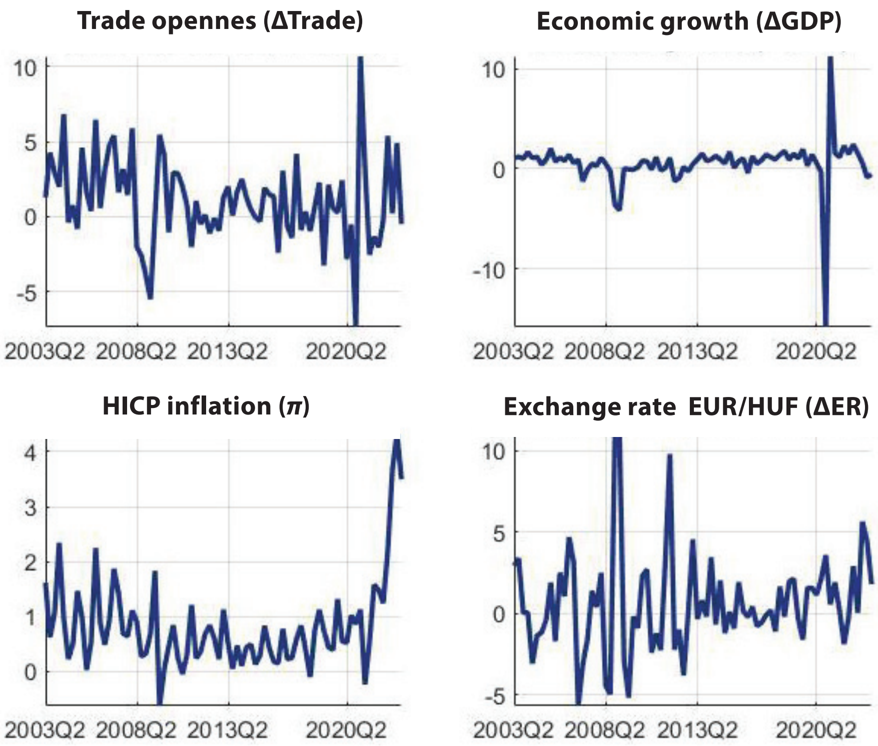


Figure 2c: Poland

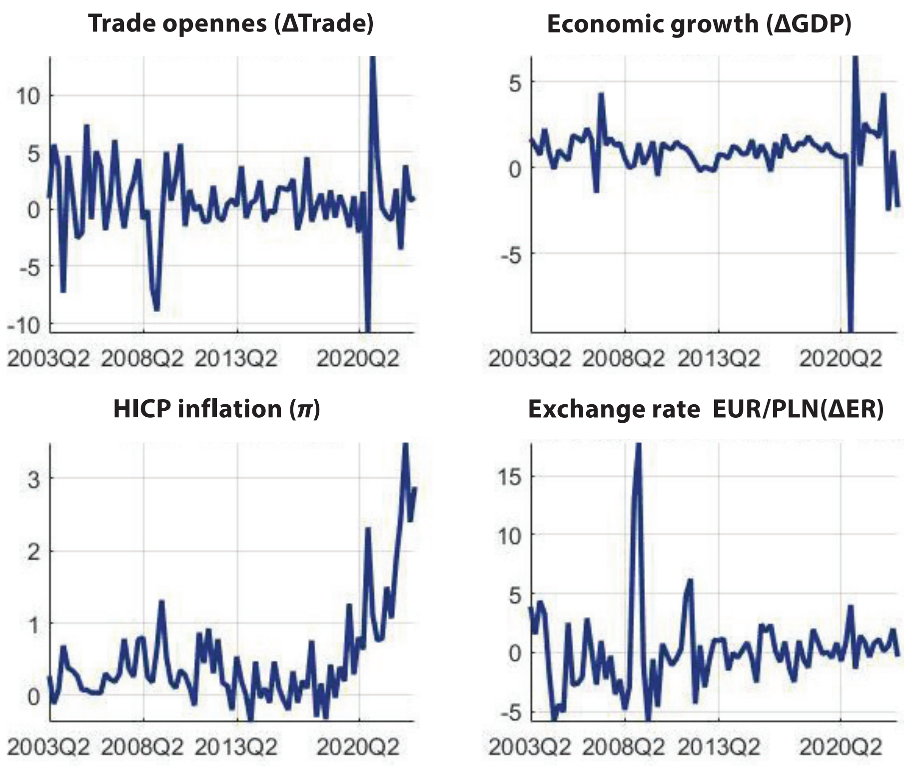
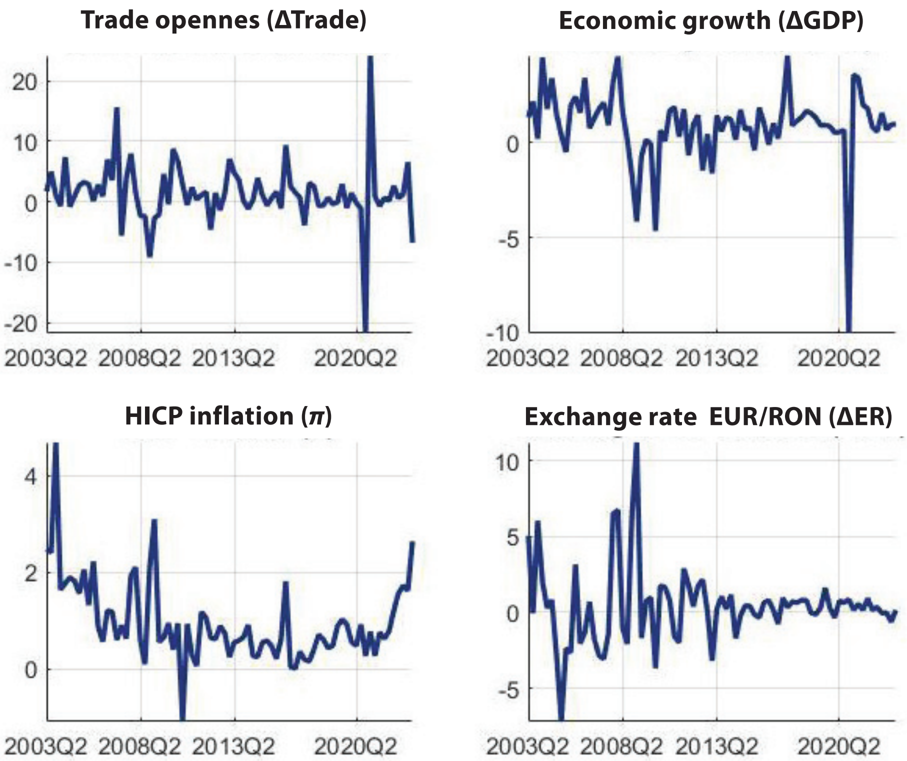


Figure 2d: Romania



Source: Eurostat database, author's calculations

5. Empirical Results

We estimate the class of TVP-SV-VAR models with the Metropolis-Hastings algorithm through Bayesian inference using Gibbs Sampler, a numerical Markov chain Monte Carlo algorithm. The nonlinear nature of the model has two sources of variation: time-varying parameters, from where we assess the influence over time of the indicators introduced in the system, and the stochastic volatility of structural shocks. To identify the optimum number of lags to be used in the VAR model, we apply lag length criteria and choose one lag according to the parsimony principle and Hannan-Quinn information criterion (see Table A2, Appendix 3).

Figure 3: Stochastic volatility of structural shocks of VAR model

Figure 3a: Czechia

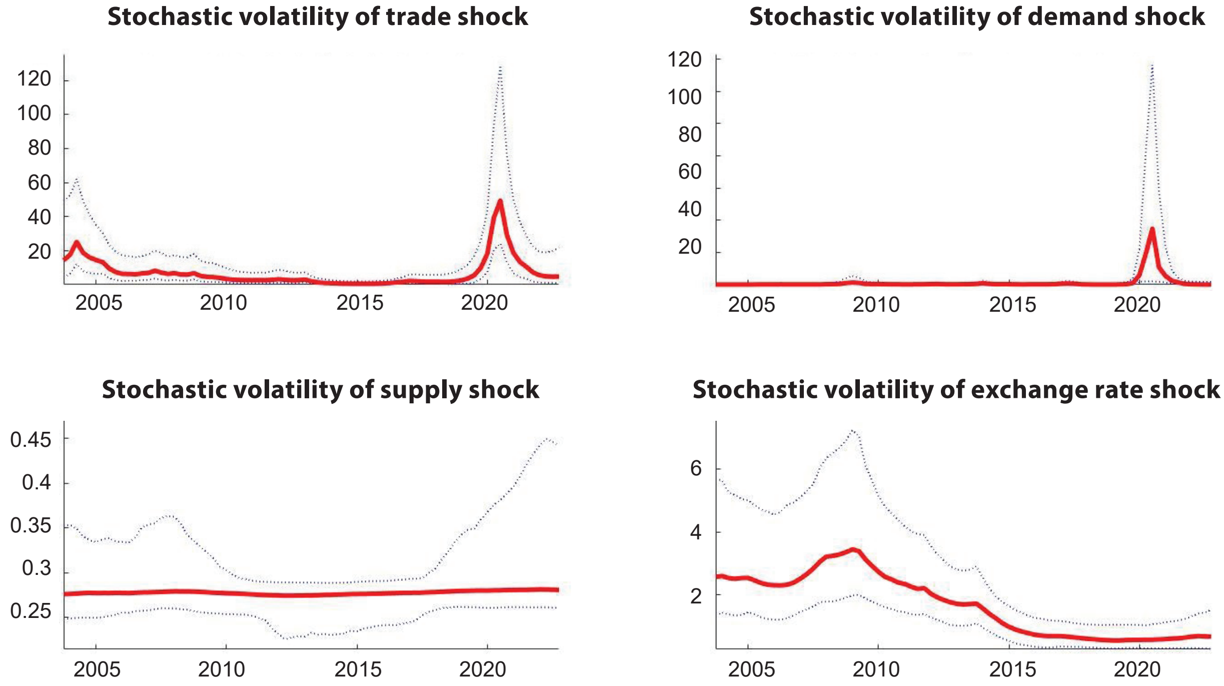


Figure 3b: Hungary

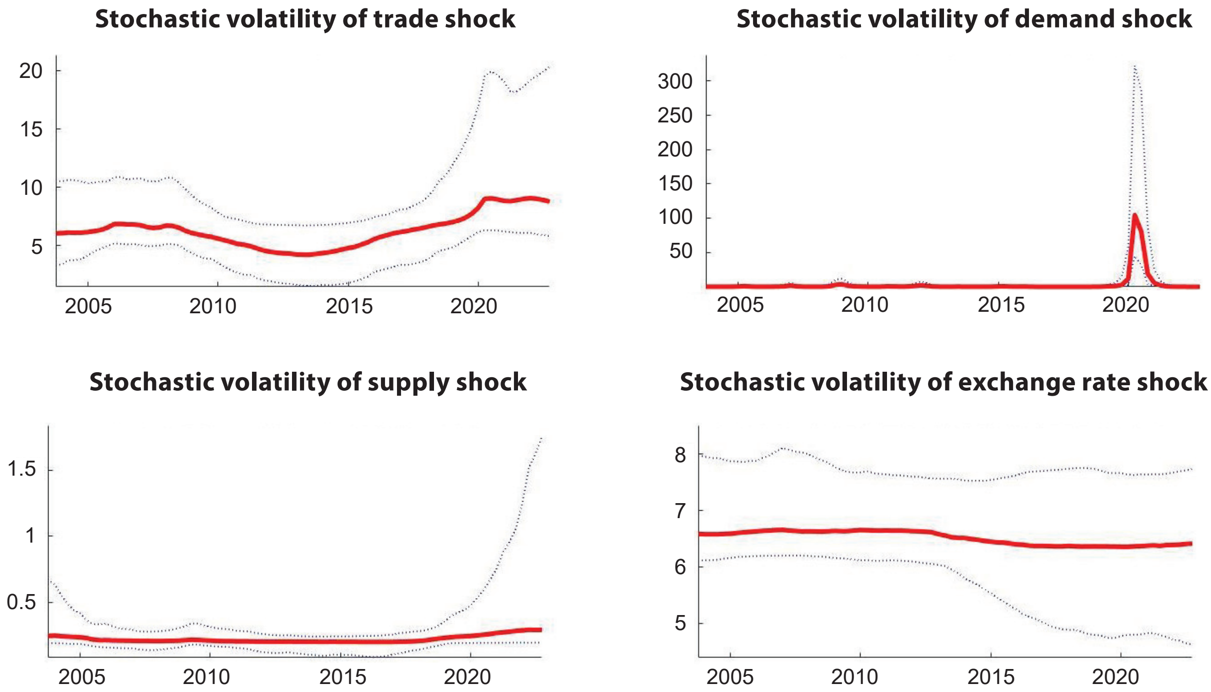


Figure 3c: Poland

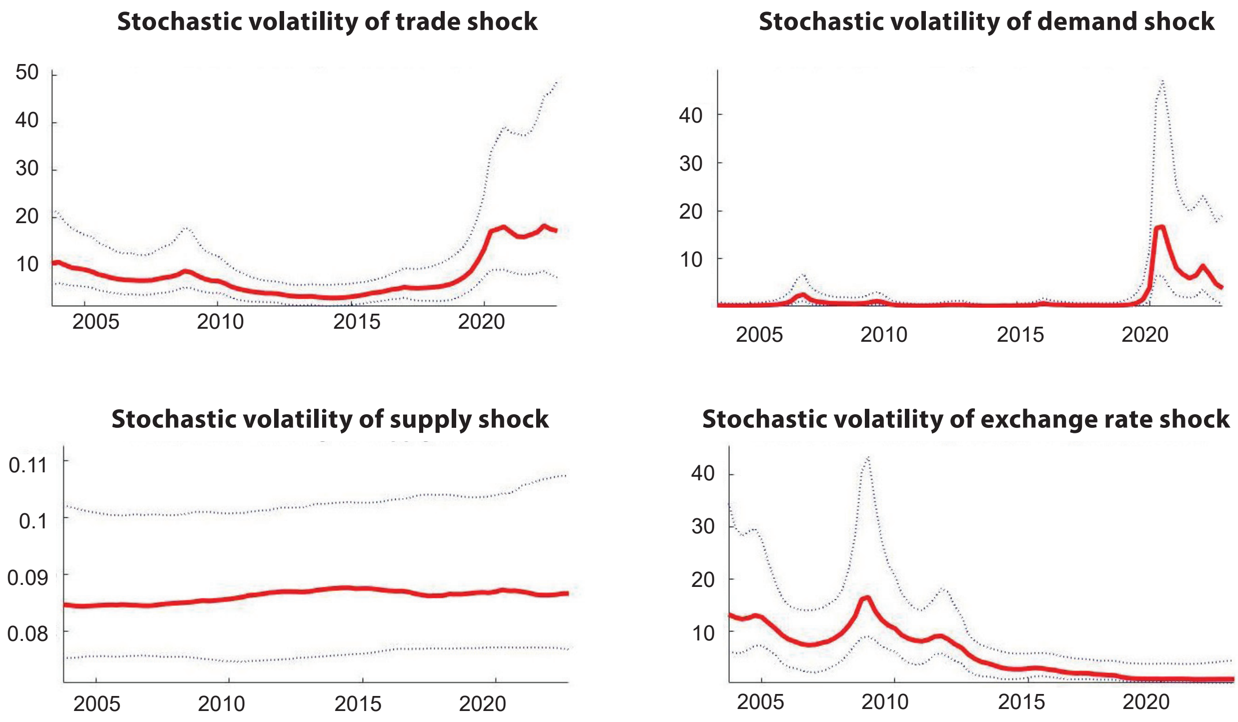
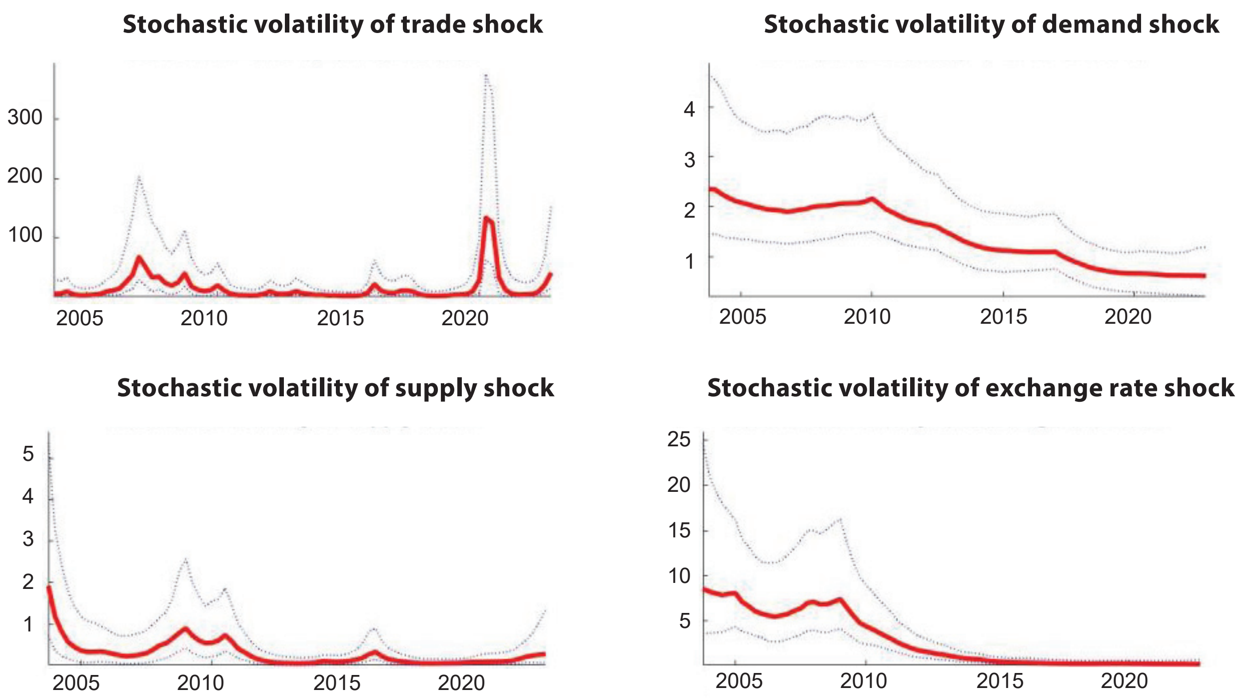


Figure 3d: Romania



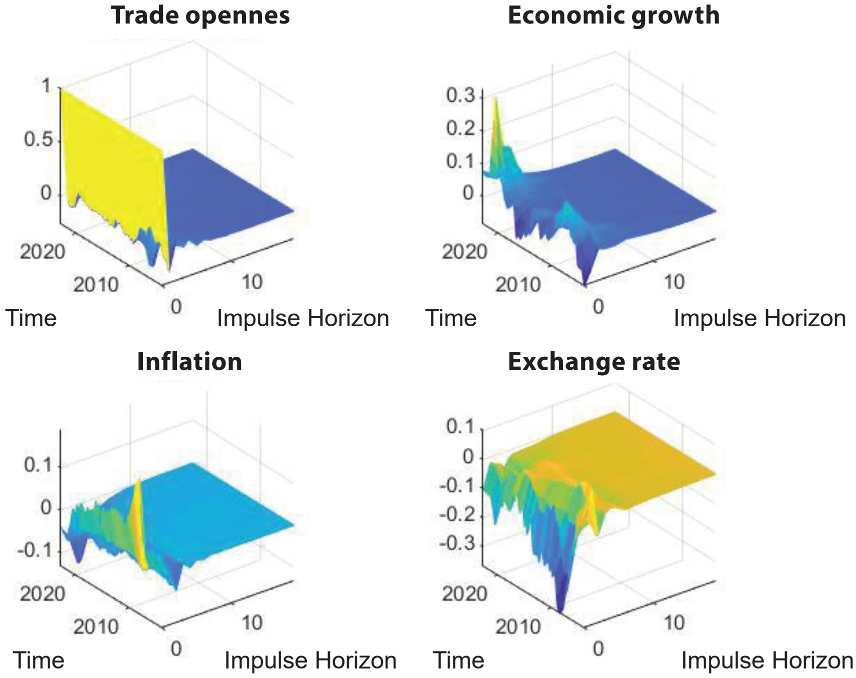
Source: author's calculations

Thus, this representation is one of the most flexible instruments for studying the effect of economic shocks over time. In Figure 3, we plot the stochastic volatility of structural shocks for this vector autoregressive model with a 90% confidence interval. The stochastic volatility of trade and demand shocks have high values not only during the financial crisis from 2008 but especially during the pandemic, at the beginning of 2020.

During the COVID-19 pandemic, support measures implemented to help citizens and companies in all European countries fostered a quick V-shaped economic recovery. Consequently, the volatility started to decline slowly in the last part of 2020. However, the confidence intervals continued to be wider, indicating uncertainty related to the economic development in the context of multiple recent shocks. The energy crisis, sharp increases in domestic prices, supply chain bottlenecks and the war in Ukraine are recent challenges that continue to disrupt economic activity and pose downside risks. The stochastic volatility of supply shocks did not fluctuate broadly between 2003Q2 and 2022Q4. This issue could be a result of including the inflation, measured by the harmonized index of consumer prices at constant taxes, excluding volatile components. Additionally, the low variability in supply shocks is consistent with the fulfilment of the optimal policy under the commitment to stabilize inflation. As for the exchange rate, the volatility of the shocks is close to zero in the last seven years due to stable domestic currencies and higher effectiveness of synchronized policy action.

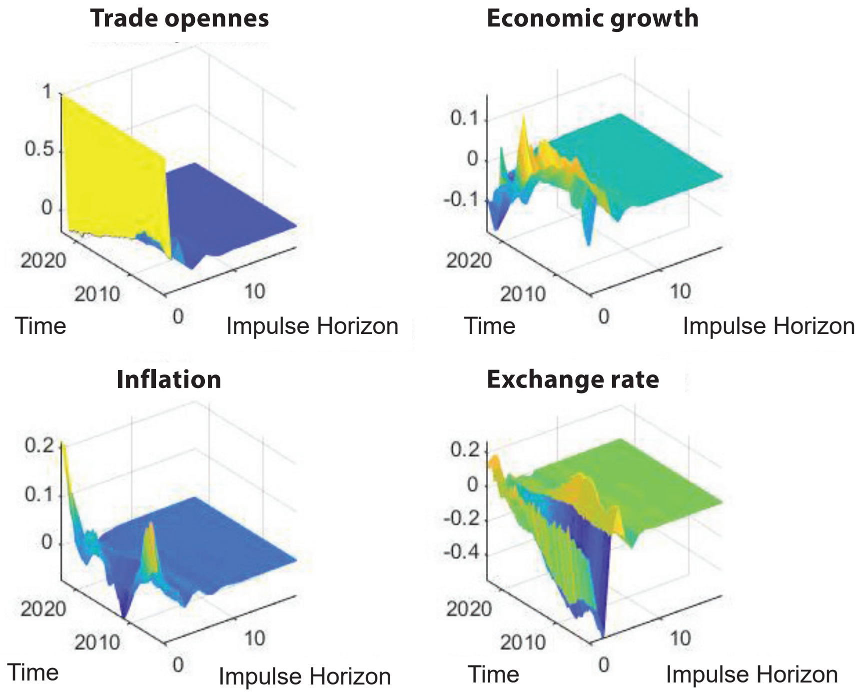
For 2003Q2–2022Q4, we represent in Figures 4 to 7 the time-varying responses of macroeconomic variables to a favourable trade openness shock, used as a proxy for advance in globalization. Given that we use the Cholesky decomposition to identify the impulse responses, the order in which we introduce the time series in the VAR is essential for contemporary impact. Therefore, we introduce the data in the following order: trade openness, economic growth and inflation, which are "slow-moving" economic indicators to which we add the exchange rate dynamics, which is a "fast-moving" indicator with changes taking place at a faster pace. In this way, trade openness does not influence the exchange rate contemporaneously but with a delay, while the reverse is true. The impulse horizon of this shock, quantified by a one-percentage point increase in trade openness, is 20 quarters, a common practice for impulse response function analysis. The results confirm, in some cases, especially in recent periods, the hypothesis of reduction of inflation and increase in economic growth because of globalization.

Figure 4: Time-varying impulse response functions (IRF) for trade openness shock (Czechia)



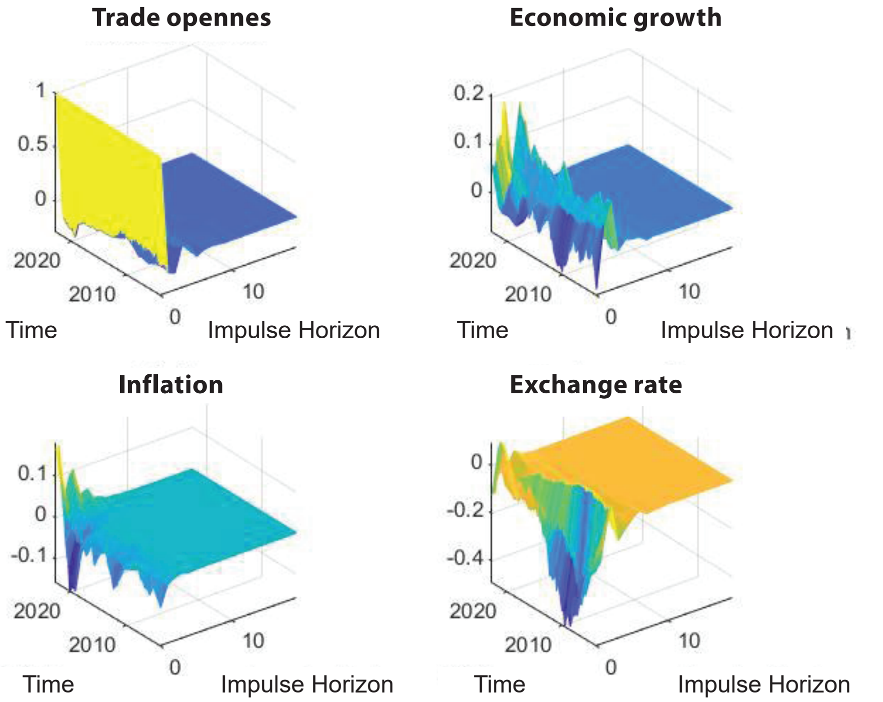
Source: author's calculations

Figure 5: Time-varying impulse response functions (IRF) for trade openness shock (Hungary)



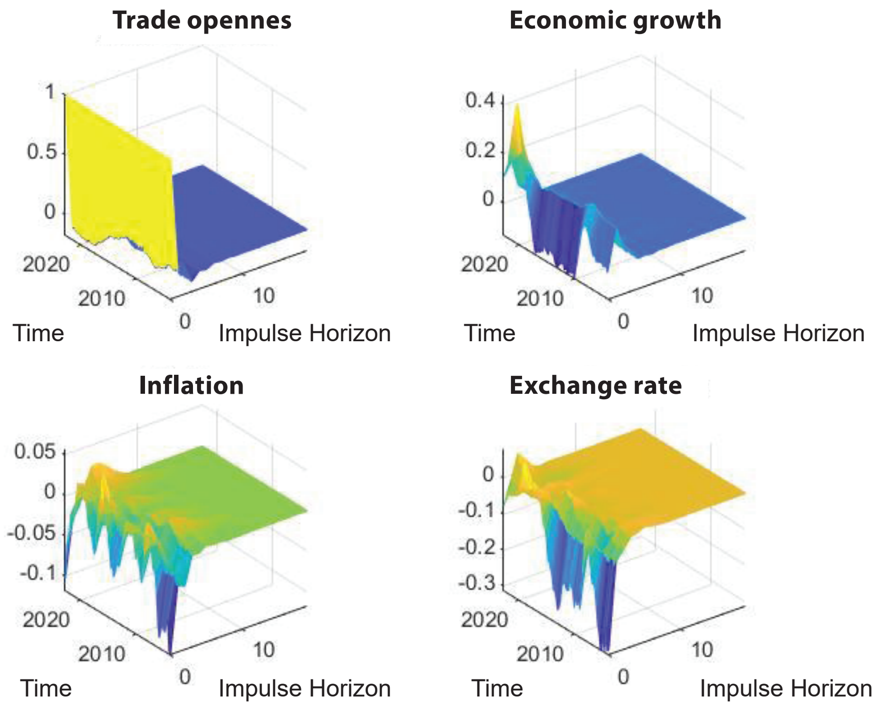
Source: author's calculations

Figure 6: Time-varying impulse response functions (IRF) for trade openness shock (Poland)



Source: author's calculations

Figure 7: Time-varying impulse response functions (IRF) for trade openness shock (Romania)



Source: author's calculations

Globalization affects prices through greater competition, participation in global value chains and cheaper imports. While in the period immediately following the financial crisis (2012–2015), a one-percentage point increase in trade openness has the effect of increasing inflation, in more recent periods (2021–2022), the same shock tends to lower prices. However, the effects are estimated to be small, i.e., less than 0.1–0.2 percentage points. The effects of a one-percentage point increase in trade openness are also estimated to improve economic activity in the last two years by approximately 0.3 percentage points (the case of Romania and Czechia).

To assess the impact of developments in globalization over time, we need to have a closer look at some distinct moments in time. In particular, under the assumption of a one-percentage point increase in trade openness, we compare the impulse response functions in 2019Q4 (before the outbreak of the pandemic crisis) and 2022Q4 (the last moment observed in the data, equal to the period after the pandemic crisis); see Figures 4 to 7. We notice that the effects of economic growth in the first period after the shock were close to zero before the pandemic crisis in 2019Q4. At the same time, in 2022Q4, they were 0.3 percentage points for Czechia and 0.4 for Romania. For Poland, we observe a similar impact in both quarters. At the same time, for Hungary in the last period, the impact is even negative in the first quarter after the shock. Regarding inflation, a clear diminishing effect in the recent period is noticeable in the case of Romania, where we see a decrease of −0.1 percentage points in inflation in 2022Q4 compared with 2019Q4, while the evidence from this model is mixed in the case of the other countries.

The results are consistent with the ECB Strategy Review (2021), which argues that globalization contributed only moderately to the reduction of the responsiveness of inflation to changes in activity. In addition, the process has made the role of the exchange rate in monetary policy more complex by introducing new mechanisms and has strengthened exchange rate valuation effects on external balance sheets. In this regard, we see in recent periods only a slow depreciation of the exchange rate for the Romanian leu, Polish zloty and Hungarian forint, which translates into higher import prices and, consequently, an increase in exports. In this way, net aggregate demand will increase, stimulating economic growth.

However, there might be effects that could mitigate the positive impact on aggregate demand. The evolution of import prices is an essential factor that might dampen the development of economic growth. In the last few years following the pandemic, the global economy has been hit hard by high levels of inflation, which consequently increased the prices of imported goods and the costs of imported inputs. On the one hand, if the economy is consumption-oriented, which is usually the case with emerging markets, as in our analysed countries[1], the rise in import prices might offset gains for increased exports. Among the analysed countries, Romania is the most heavily consumption-oriented country; therefore, the increases in import prices due to global inflation might limit the impact of globalization.

On the other hand, according to economic theory, countries that are export-oriented and, at the same time, rely on imported inputs might suffer from increases in prices and dampen the effects of increased exports and trade openness. Looking at the economic structure, in terms of exports and imports of goods, these countries have relatively close levels of exports and imports of goods and services; especially after the global financial crisis, exports slightly dominate imports. Romania is the exception in that the country is highly dependent on imports[2].

Nevertheless, the effects are not only limited but also isolated in specific periods in time. For a more general overview of the results and robustness check, we split the period into four relevant subperiods from economic perspectives and events (before, during and after global financial crisis and the period that started with the outbreak of the pandemic). We represent the average of impulse response functions over these periods to quantify the effects in each of these regimes. We plot the impulse response functions following an increase in trade openness by one-percentage point in Figures 8 to 10.

The response of economic growth as a result of developments in globalization is estimated to be elevated, especially for Czechia, but also for Poland and Romania. At the same time, for Hungary, we estimate higher effects in response to globalization during the financial crisis than during the pandemic, where the results are negative in the mean. Thus, we observe that while globalization produces favourable effects on macroeconomic activity for Czechia, Poland and Romania (by reducing inflation and increasing economic growth), the process did not improve these dynamics for Hungary. However, in some cases (e.g., response to inflation), we observe minimal effects, given that inflation did not respond in the same way over the entire period (in the 3D representations, we saw that the globalization shock was inflationary in some periods and deflationary in others). In this context, the effects should also be interpreted with a certain degree of caution because of the many other shocks that hit the worldwide economies, especially in recent times.

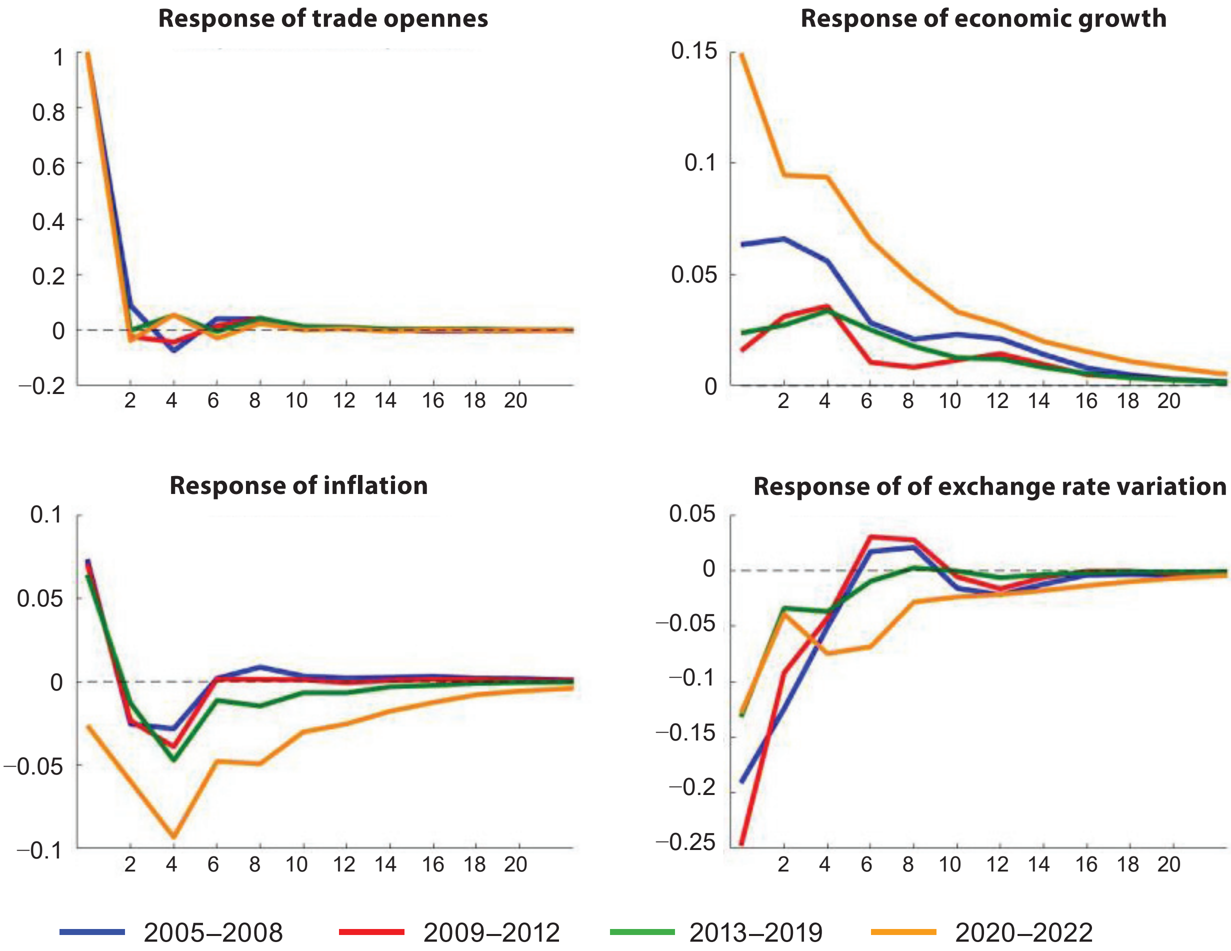
Even if Hungary's results might contradict economic theory, a few particularities about this country could have influenced the effect of globalization in recent years. One explanation is

1 In the last ten years (2013–2023), the share of final consumption in GDP is, on average, 65% in Czechia, 72% in Hungary, 76% in Poland and 80% in Romania.

2 From 2021 to 2023, Romania's imports were noticeably higher than exports (with around ten percentage points of GDP).

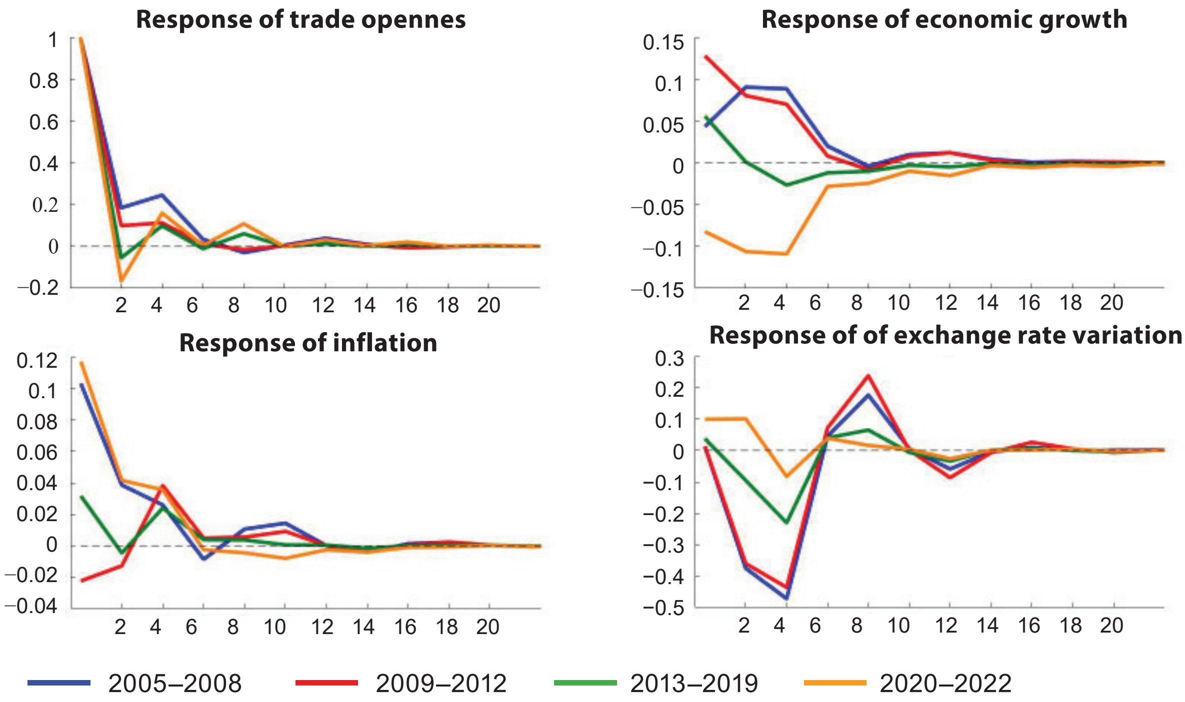
related to the abovementioned structure of the economy. Hungary is one of the most important exporters of cars in Europe, the automotive sector being crucial for Hungary's exports, accounting for around 20% of total exports in 2022. This sector is mainly dependent on inputs such as microchips and semiconductors; therefore, despite the developments of trade openness, the global supply chain bottlenecks that were exacerbated in 2021–2022 led to microchip shortages and a lack of semiconductors, which consequently increased production costs and negatively affected output. Moreover, the issue of Hungarian economic growth and inflation, which are exposed to global value chains, was recently documented by Koppany et al. (2023). The authors pointed out that during the recent energy crisis, Hungary was highly exposed to changes in prices of energy from Germany, Austria and Russia, to which add crude oil and natural gas price booms. These issues result in significant increases in consumer price levels that might dampen any favourable effects of globalization in times of crisis.

Figure 8: Responses of variables to trade openness shock – subperiods (Czechia)



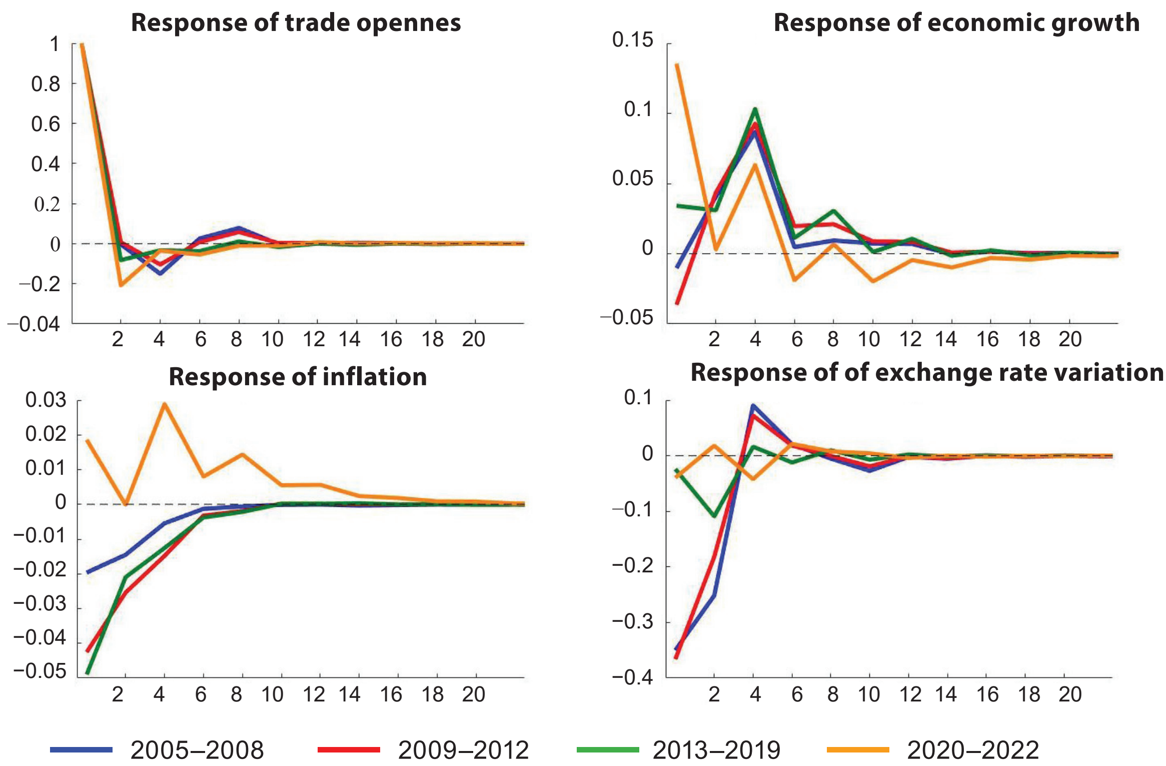
Source: author's calculations

Figure 9: Responses of variables to trade openness shock – subperiods (Hungary)



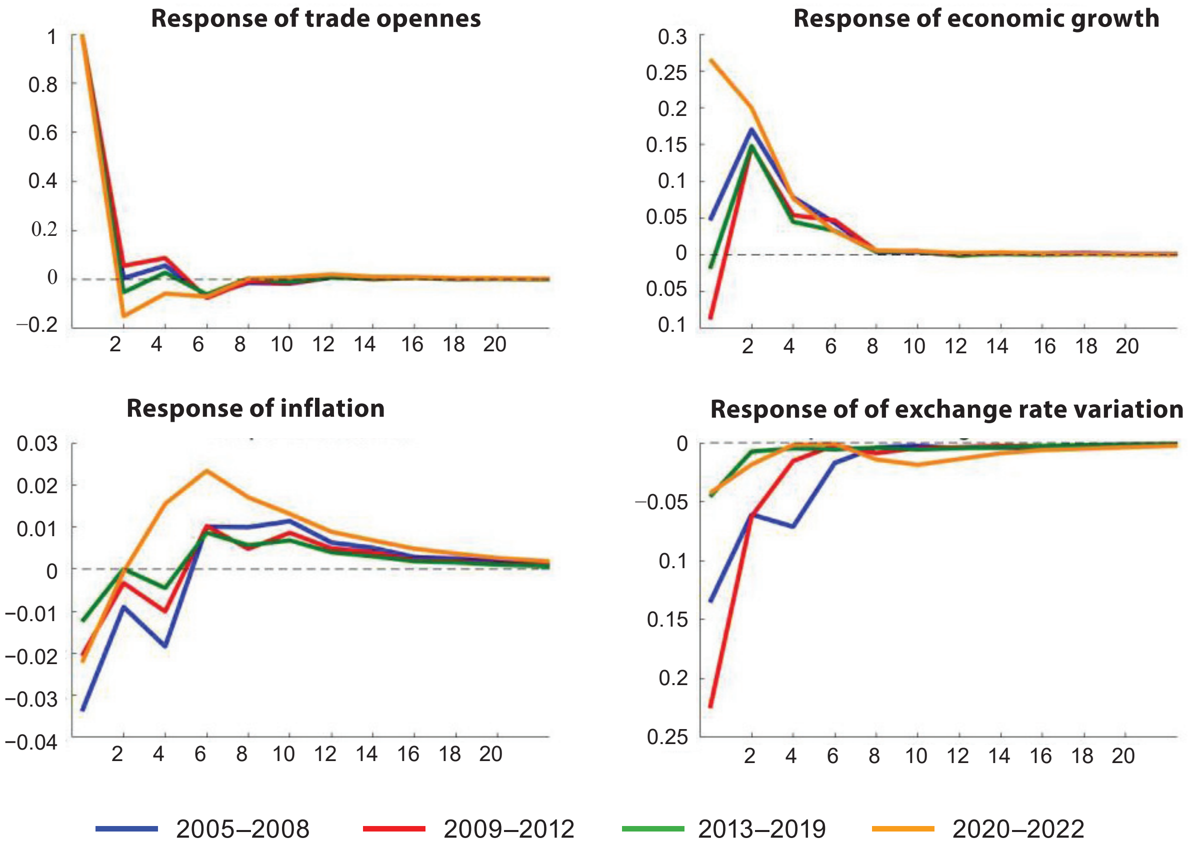
Source: author's calculations

Figure 10: Responses of variables to trade openness shock – subperiods (Poland)



Source: author's calculations

Figure 11: Responses of variables to trade openness shock – subperiods (Romania)



Source: author's calculations

The empirical results point out more clearly improvements in inflation, given a trade openness shock during the coronavirus pandemic for Czechia and Romania. After developments in globalization quantified by a one-percentage point increase in trade openness, we expect a small decrease in inflation of −0.02 for Czechia and −0.05 for Romania. In the case of Hungary, we note a positive effect on inflation and a trajectory of the indicator after the shock that is much more volatile until it returns to the pre-shock level. However, the effects of the shocks are not persistent; the level of inflation gradually returns to the one before the shock after six periods, except for the pandemic period, when the responses exhibit a more pronounced degree of persistence. Regarding the exchange rate dynamics, the degree of currency appreciation is smaller in recent periods, meaning that the loss in price competitiveness due to currency appreciation might be mitigated.

6. Concluding Remarks

In this paper, we assessed the impact of globalization in four CEE emerging countries with similar economic characteristics (Czechia, Hungary, Poland and Romania). These are all small open economies that are not yet euro area members and have inflation targeting and floating exchange rate regimes, whose trade linkages might affect macroeconomic dynamics. This analysis was conducted for the period 2003Q2–2022Q4, covering some essential extreme events such as the global financial crisis in 2008 and the coronavirus outbreak at the beginning of 2020. As an econometric technical tool, we used a time-varying vector autoregressive model with stochastic volatility and Bayesian estimation inference. This proved to be one of the most appropriate methods to study over time the effects of increasing trade openness as a proxy for progress in globalization.

The findings of this research hold significant implications for understanding the impact of globalization on inflation in emerging economies. The results reveal a moderate reduction in inflation as a result of globalization for countries such as Czechia, Poland and Romania. The responses from impulse response functions show pronounced effects in the last two years after the pandemic crisis. However, the estimations remain subject to a certain degree of uncertainty given the recent events and multiple other shocks that hit the worldwide economies (i.e., the energy crisis, inflation, supply chain bottlenecks and the war in Ukraine). However, in the case of Hungary, we did not find evidence for favourable effects of globalization on the macroeconomic dynamics (economic growth, inflation and exchange rate). Hungary was highly affected by the global supply bottlenecks and energy crisis because it is one of Europe's most important car exporters and is highly dependent on natural gas and oil imports from Russia.

To summarize, while there have been some notable changes in macroeconomic dynamics due to globalization in the last two decades, the empirical models used to make projections need to account more for the implications of this phenomenon and to incorporate global factors into their representations. This aspect is crucial for transmission mechanism channels to understand and assess which shocks are drivers of economic activity and consequently, to develop appropriate policy measures. Furthermore, given that some studies demonstrate that globalization produces disruptions in inequality, this aspect remains a compelling area for further research.

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Appendix 1: Gibbs sampling and Metropolis-Hastings algorithm steps

According to Blake and Mumtaz (2017), first, we need to set a prior for **Q** and initial values for the Kalman filter. This prior is defined by an inverse Wishart distribution: $p(\boldsymbol{Q}) \sim IW(\boldsymbol{Q}_0, \boldsymbol{T}_0)$. The procedure is done by using a training sample (we follow Primiceri, 2005, who proposed $\boldsymbol{T}_0$ = 40 observations) and the first $\boldsymbol{T}_0$ observations are used to estimate via OLS the coefficients of a standard VAR model with the parameters $\boldsymbol{\beta}_0 = (\boldsymbol{X}'_{0,t}\boldsymbol{X}_{0,t})^{-1}(\boldsymbol{X}'_{0,t}\boldsymbol{X}_{0,t})$, where the training repressors are $\boldsymbol{X}_{0,t} = \{\boldsymbol{Y}_{0,t-1} \ldots, \boldsymbol{Y}_{0,t-p}, 1\}$ and the variance-covariance matrix is given by $\boldsymbol{\Sigma}_0 = \frac{(\boldsymbol{Y}_{0,t} - \boldsymbol{X}_{0,t}\boldsymbol{\beta}_0)'(\boldsymbol{Y}_{0,t} - \boldsymbol{X}_{0,t}\boldsymbol{\beta}_0)}{\boldsymbol{T}_0 - \boldsymbol{K}}$. The scale matrix $\boldsymbol{Q}_0$ is important because it allows time variation in the vector autoregressive model, meaning that a large value is equivalent to more fluctuations in $\boldsymbol{\beta}_t$ (Blake and Mumtaz, 2017). We impose the prior to be equal to $\boldsymbol{p}_{0|0} \times \boldsymbol{T}_0 \times \boldsymbol{\tau}$, where the scale factor τ is set to $\boldsymbol{\tau} = 10^{-4}$ according to Primiceri (2005) – a very small value to underline the reduced effect of a typical short training sample. Therefore, estimating a complex model with time-varying parameters becomes challenging when the dataset sample is not large enough. The initial state values for coefficients are $\boldsymbol{\beta}_{0|0} = vec(\boldsymbol{\beta}_0)$ and the initial covariance is given by the $p_{0|0}$ elements. The priors used for $\boldsymbol{D}_i$, $i = \overline{1,3}$ are inverse gamma $p(\boldsymbol{D}_i) \sim IG\ (\boldsymbol{D}_{i,0}, T_0)$. Benati and Mumtaz (2006) proposed the following values: $\boldsymbol{D}_1 = 0.0001$, $\boldsymbol{D}_2 = 0.001\boldsymbol{I}_2$, $\boldsymbol{D}_3 = 0.001\boldsymbol{I}_3$. We impose initial values for $\boldsymbol{h}_{ij,t}$ and $\boldsymbol{a}_{ij,t}$ as in Benati and Mumtaz (2017).

Conditional on $\boldsymbol{A}_t, \boldsymbol{H}_t, \boldsymbol{Q}$, sample $\boldsymbol{\beta}_t$, we use Carter and Kohn's (1994) algorithm, where $\boldsymbol{v}_t$ varies at each point in time. Then, we use those draws to calculate the residuals for the transition equation $\boldsymbol{\beta}_t - \boldsymbol{\beta}_{t-1} = \boldsymbol{e}_t$ and sample **Q** from the inverse Wishart distribution with scale the parameter $\boldsymbol{e}_t'\boldsymbol{e}_t + \boldsymbol{Q}_0$ and $\boldsymbol{T} + \boldsymbol{T}_0$ degrees of freedom.

Conditional on $\boldsymbol{\beta}_t$, $\boldsymbol{A}_t$, $\boldsymbol{D}_1$, $\boldsymbol{D}_2$, $\boldsymbol{D}_3$, we draw $\boldsymbol{a}_{ij,t}$ (the elements below the main diagonal of A_t). The state-space formulation for each element $\boldsymbol{a}_{ij,t}$ is determined by the repressors of the time-varying equations and the algorithm is applied individually.

We simulate residual values for $\boldsymbol{u}_{i,t}$, $i = \overline{1,4}$. The posteriors for $\boldsymbol{D}_1$ are generated from inverse Gamma with the scale parameter defined by $\frac{\boldsymbol{u}'_{1,t}\boldsymbol{u}_{1,t} + \boldsymbol{D}_{1,0}}{2}$ and $\frac{\boldsymbol{T} + \boldsymbol{T}_0}{2}$. The other $\boldsymbol{D}_i$ for i = 2,3, are simulated in the same way with the scale parameter $\boldsymbol{u}'_{i,t}\boldsymbol{u}_{i,t} + \boldsymbol{D}_{i,0}$ and $\boldsymbol{T} + \boldsymbol{T}_0$ degrees of freedom.

Using the draws of A_t we calculate the vector $\varepsilon_t = (\varepsilon_{1,t}\ \ \varepsilon_{2,t}\ \ \varepsilon_{3,t}\ \ \varepsilon_{4,t})^T$, which is contemporaneously uncorrelated. After these calculations, we can draw $h_{ij,t}$ by applying the independent Metropolis-Hastings algorithm. Conditional on $h_{ij,t}$, we draw elements of the variance-covariance matrix G of the stochastic volatility from the inverse gamma distribution with the scale parameter defined by $\frac{\left(\ln h_{i,t} - \ln h_{i,t-1}\right)'\left(\ln h_{i,t} - \ln h_{i,t-1}\right) + g_0}{2}$ and $\frac{T + v_0}{2}$ degrees of freedom. For convergence, we use the widely known Bayesian econometric procedure "burn-in" and discard the first 95,000 iterations from a total of 100,000. The algorithm and the procedures explained above were implemented in MATLAB R2018a.

Appendix 2: Impulse response functions (IRFs) of VAR model

Impulse response functions are used to assess the relations between variables in a VAR model. If we consider a stationary VAR(p) process for the dataset Y_t, then it has a Wold moving process as in Lütkephol (2005):

$$Y_t = \sum_{j=0}^{\infty} \Phi_j \upsilon_{t-j} \tag{11}$$

where υ_t are residuals, potentially correlated, and the Φ_j are the (K × K) coefficient matrices. The impulse response functions represent the marginal response $y_{n,t+j}$ to a unit change in y_{mt}, holding constant all past values of y_t given by (n, m)th elements of the matrices Φ_j, viewed as a function of *j*. In other words, the elements of Φ_j represent responses called forecast error impulse responses.

For stationary processes, $\Phi_j \rightarrow 0$ and $j \rightarrow \infty$, meaning that the effect of an impulse is transitory and vanishes over time. As stated in Lütkepohl (2005), forecast error impulse responses may not reflect the actual true responses of a system given that υ_t are contemporaneously correlated and the variance covariance matrix Σ_υ is not diagonal. Therefore, orthogonalized shocks are often considered for the impulse response analysis. As we mention in the Section 3, we can use a nonsingular matrix A with the property $AA' = \Sigma_\upsilon$ to define the orthogonalized shocks $\varepsilon_t = A^{-1}\upsilon_t$. These shocks are contemporaneously uncorrelated and the responses to them are given by the coefficients of the moving-average representation:

$$Y_t = \sum_{j=0}^{\infty} \Phi_j A A^{-1} \upsilon_{t-j} = \sum_{j=0}^{\infty} \Psi_j \varepsilon_{t-j} \tag{12}$$

In this representation, the matrix A is not unique and there exist many different specifications for $\Phi_j - \Phi_j A$. Thus, to identify the impulses that are relevant from an economic point of view, one needs to impose a restriction on ***A***, which results in a unique representation for impulse responses.

According to the literature, a popular choice of ***A*** is a lower triangular matrix given by Cholesky factorization. As we mentioned in the text, this setup has implications for the contemporaneous effects of the impulse responses over the variables introduced in the model.

Appendix 3: Stationarity tests for time series and lag length criteria for VAR model

Table A1: Stationarity tests for time series

Table 1a: Czechia

Test	Test statistic	Trade openness	Economic growth	Inflation	Exchange rate
Augmented Dicky Fuller	t-statistic	−9.312***	−7.925***	−5.230***	−7.100***
Phillips Perron	adjusted t-statistic	−9.335***	−7.942***	−5.003***	−6.948***
KPSS	LM-statistic	0.321***	0.207***	0.439**	0.137***

Table 1b: Hungary

Test	Test statistic	Trade openness	Economic growth	Inflation	Exchange rate
Augmented Dicky Fuller	t-statistic	−8.547***	−10.978***	−4.589***	−8.373***
Phillips Perron	adjusted t-statistic	−3.095**	−10.751***	−4.801***	−8.836***
KPSS	LM-statistic	0.409***	0.104***	0.283***	0.327***

Table 2c: Poland

Test	Test statistic	Trade openness	Economic growth	Inflation	Exchange rate
Augmented Dicky Fuller	t-statistic	−9.976***	−11.932***	−2.945**	−6.933***
Phillips Perron	adjusted t-statistic	−10.233***	−11.864***	−2.931**	−6.572***
KPSS	LM-statistic	0.073***	0.0719***	0.507**	0.14***

Table 2d: Romania

Test	Test statistic	Trade openness	Economic growth	Inflation	Exchange rate
Augmented Dicky Fuller	t-statistic	−10.859***	−7.785***	−5.187***	−6.471***
Phillips Perron	adjusted t-statistic	−10.856***	−7.798***	−5.087***	−6.899***
KPSS	LM-statistic	0.177***	0.15***	0.548***	0.049***

Notes: KPSS is Kwiatkowski–Phillips–Schmidt–Shin test.

Statistical significance is given by the following notation: * at the 10% level, ** at the 5% level and *** at the 1% level.

Source: author's calculations

Table A2: Lag length criteria for VAR model

Table 2a: Czechia

Lag	LogL	LR	FPE	AIC	SC	HQ
0	−534.815	NA	45.897	15.178	15.305*	15.229
1	−510.416	45.363	36.259*	14.941*	15.579	15.194*
2	−496.084	25.029	38.178	14.988	16.136	15.445
3	−485.450	17.375	44.909	15.139	16.797	15.798
4	−464.147	32.404*	39.506	14.990	17.157	15.852

Table 2b: Hungary

Lag	LogL	LR	FPE	AIC	SC	HQ
0	−594.911	NA	249.441	16.871	16.998*	16.921
1	−570.815	44.798	198.760	16.643	17.280	16.896*
2	−551.637	33.493	182.571	16.553	17.700	17.009
3	−532.541	31.199	169.212	16.466	18.123	17.125
4	−506.581	39.489*	130.553*	16.185*	18.352	17.047

Table 2c: Poland

Lag	LogL	LR	FPE	AIC	SC	HQ
0	−576.357	NA	147.906	16.348	16.476	16.399
1	−532.763	81.047	68.049	15.571	16.208*	15.824*
2	−510.492	38.897	57.289	15.394	16.541	15.850
3	−502.228	13.501	72.043	15.612	17.269	16.271
4	−476.562	39.042	56.046	15.340	17.507	16.202

Table 2d: Romania

Lag	LogL	LR	FPE	AIC	SC	HQ
0	−582.665	NA	176.666	16.526	16.653*	16.576
1	−551.325	58.266	114.787	16.094	16.731	16.347*
2	−535.031	28.457*	114.359*	16.085*	17.233	16.542
3	−527.167	12.847	145.442	16.315	17.972	16.974
4	−512.460	22.372	154.069	16.351	18.518	17.213

Notes: * indicates lag order selected by the criteria
LR: sequential modified LR test statistic (each test at the 5% level)
FPE: Final prediction error
AIC: Akaike information criterion
SC: Schwarz information criterion
HQ: Hannan Quinn information criterion

Source: author's calculations

Appendix 4: Convergence of Metropolis-Hastings algorithm

Figure 12: Recursive means for key parameters of time-varying VAR

Figure 12a: Czechia

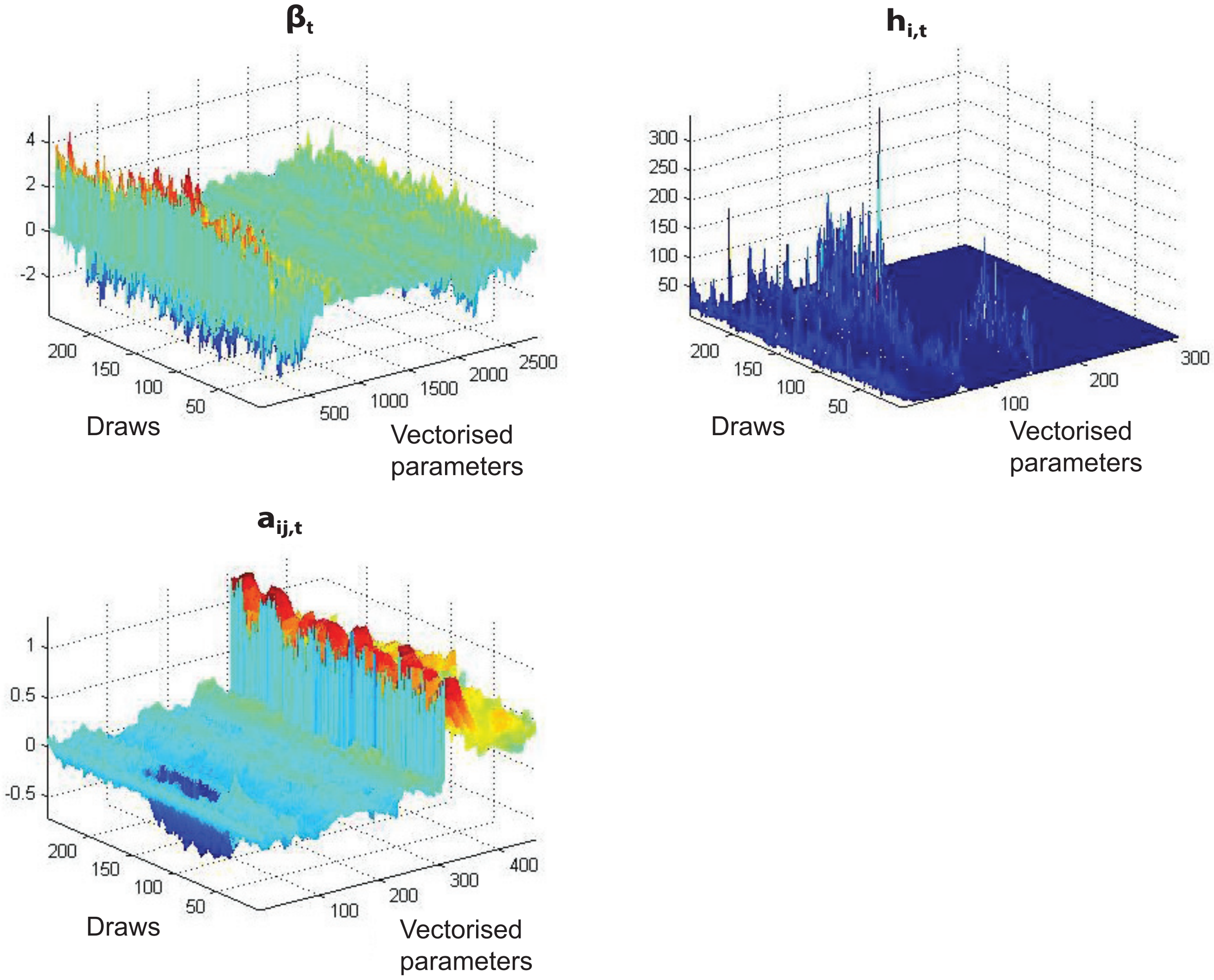


Figure 12b: Hungary

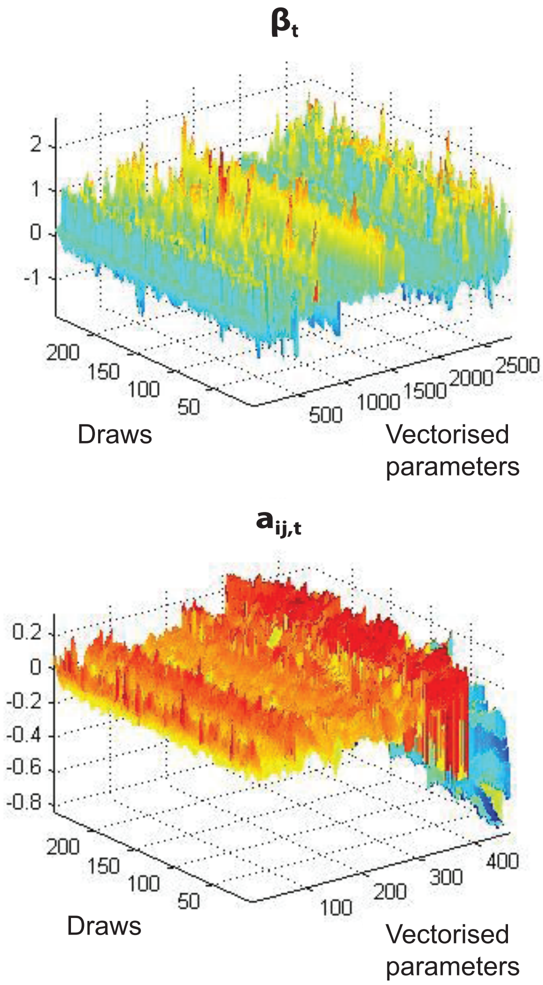


Figure 12c: Poland

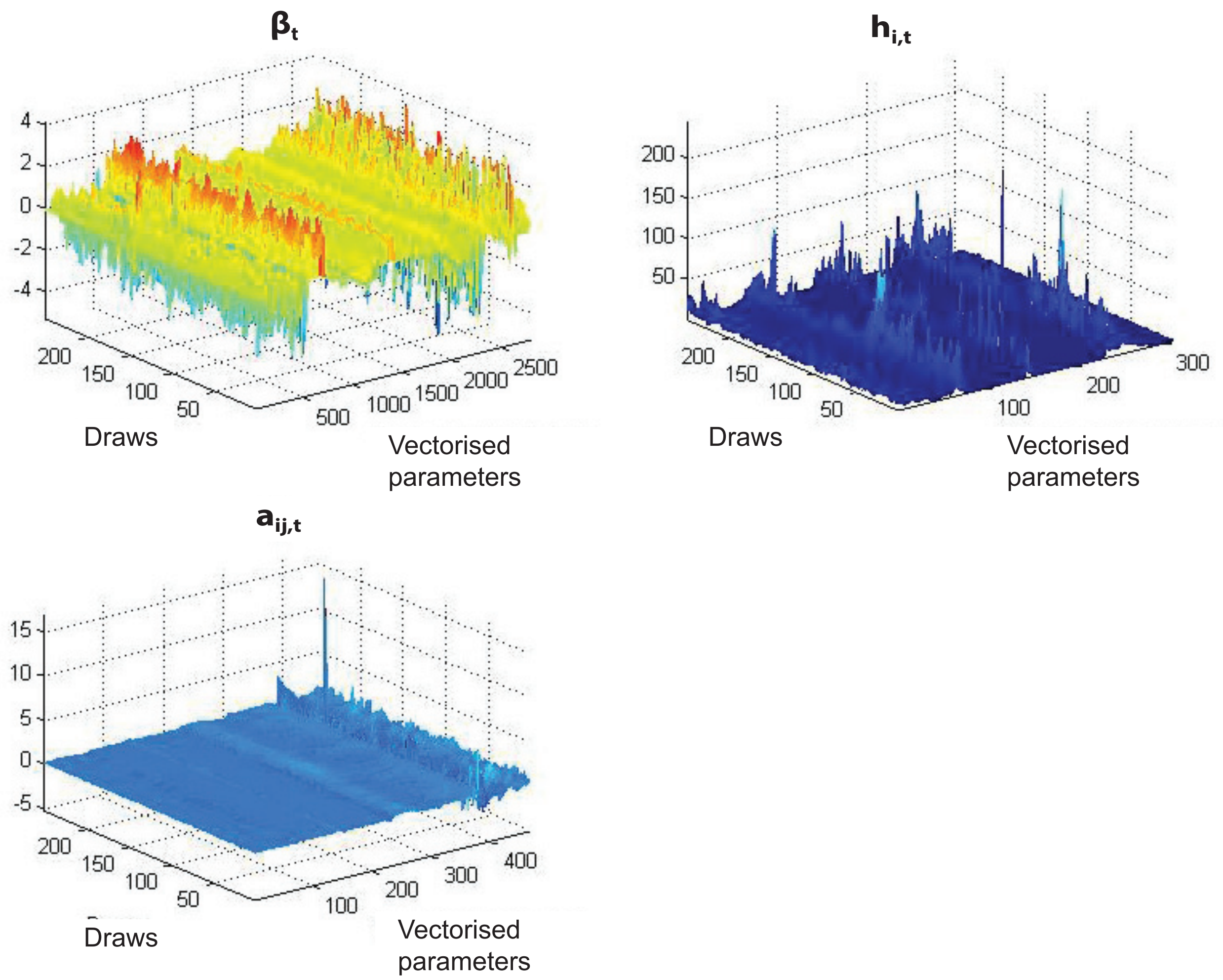
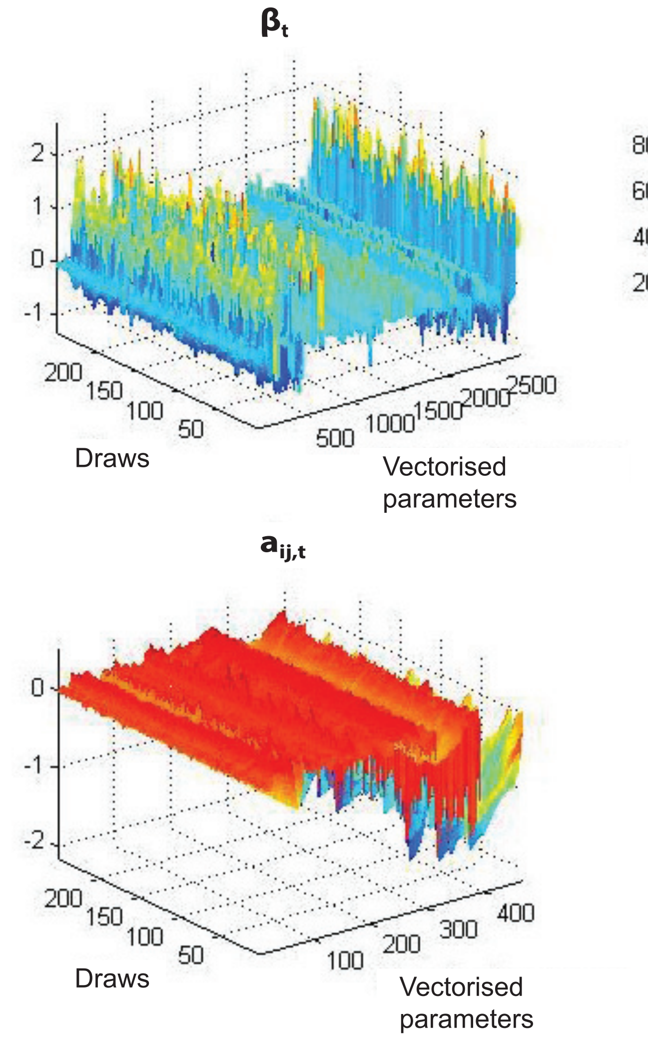


Figure 12d: Romania



Source: author's calculations