Insulating property of the flexible exchange rate regime: A case of Central and Eastern European countries

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Abstract

We examine the insulating property of flexible exchange rate in CEE economies using the fact that they have adopted different regimes. A set of Bayesian structural VAR models with common serial correlations is estimated on data spanning 1998q1-2015q4. The long-term identifying restrictions are derived from a macroeconomic model. We find that irrespective of the exchange rate regime output is driven mainly by real shocks. Its reactions to these shocks, however, are substantially stronger under less flexible regimes, whereas the responses to nominal shocks are similar. Hence, the insulating property of flexible regimes can reduce the costs from economic shocks.

Keywords: open economy macroeconomics; exchange rate regimes; real and nominal shocks; Bayesian structural VAR; common serial correlation

JEL Classification: F33; C11; F41; E44

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1 Introduction

The insulating properties of the exchange rate regime together with its impact on policy effectiveness and importance for the adjustment to trade imbalances constitute one of 'three main strands' in the literature on the choice of exchange rate regime (Ghosh et al., 2010).[1] One of the central findings of this strand is that the floating exchange rate better insulates output against real shocks as it facilitates adjustment in the face of nominal rigidities, whereas under the fixed exchange rate nominal shocks are automatically offset by foreign exchange reserves movements (see, e.g., Lahiri et al., 2008). The flexible exchange rate can, therefore, serve as a shock absorber unless 'disruptions come from asset markets or unstable money demand' (Klein and Shambaugh, 2010). In a nutshell, using words of Ghosh et al. (2002), 'the relative incidence of nominal and real shocks becomes a key criterion in choosing the exchange rate regime.'

Basically there are two main empirical approaches to study the insulating properties of the exchange rate regime. The first one examines the relationship between volatility of output growth and the exchange rate flexibility. The results on this relationship are mixed: although there is some evidence of greater output volatility under pegged exchange rates than under either intermediate or floating regimes, the result holds for advanced and developing economies but not for emerging market economies (EMEs) where output volatility is lower under pegged and intermediate regimes (Ghosh et al., 2010, p. 17). In an earlier study Levy-Yeyati and Sturzenegger (2003) found a negative link between output volatility and exchange rate flexibility for nonindustrial countries but not for industrial ones (see also Ghosh et al., 2002, pp. 71 and 99). Similar conclusion has been drawn in more recent studies by Edwards (2011) and Erdem and Özmen (2015). They found that the impact of external shocks on economic activity is less pronounced in economies under flexible exchange rate regimes.[2] Likewise, Obstfeld et al. (forthcoming) found that economies under fixed regimes experience greater output volatility than those under floating regimes.

The second approach uses structural vector autoregressive (VAR) models to identify shocks hitting an economy, to assess their importance for the output variance and in some studies to examine output reactions to real and nominal shocks. The results are again inconclusive. After examining five Central and Eastern European

[1] The other two prominent strands weigh benefits and costs of adopting a pegged exchange rate or a common currency and consider an exchange rate peg as a precommitment device helping a central bank to disinflate by disciplining credit expansion and by engendering confidence in the currency.

[2] See also Bohl et al. (2016) and Ihnatov and Căpraru (2012) who found that pegged exchange rate regimes exert a negative impact on economic growth.

(CEE) countries Borghijs and Kuijs (2004) concluded that 'the exchange rate appears to have served as much or more as an unhelpful propagator of monetary and financial shocks than as a useful absorber of real shocks.' A similar point was raised by Shevchuk (2014) who found that the exchange rate variability was driven mainly by neutral shocks in 14 CEE countries. In turn, Stążka-Gawrysiak (2009) found that the flexible exchange rate of the Polish zloty had been 'a shock-absorbing rather than a shock-propagating instrument.'[3] In line with her study are the results obtained by Dąbrowski and Wróblewska (2016) who found that 'the hypothesis that the flexible exchange rate is not a shock absorber rests on the imperfect identification of shocks' and demonstrated that the floating exchange rate of the Polish zloty better insulated a real economy against real shocks than the relatively fixed rate of the Slovak koruna.

The research objective of this paper is to examine whether there are discernible differences across exchange rate regimes in CEE economies with respect to insulating output against economic shocks. We follow the second empirical approach and construct a set of Bayesian structural VAR models for each out of eight CEE countries in our sample and focus on comparison of output response functions to real and nominal shocks.

Our contribution to the literature can be summarized in four main points. First, we provide new empirical evidence on the usefulness of the flexible exchange rate regime in CEE countries. The importance of this issue stems from the fact that all these countries are small open economies and the choice of an inferior regime can potentially result in suboptimal reactions of real output, employment and other real variables. Our main finding is that the exchange rate flexibility in CEE economies indeed provides the greater insulation against real shocks. Hence, our finding is in line with a more general point made by Obstfeld et al. (forthcoming) that the exchange rate regimes in EMEs *do* matter and it is complementary to their finding that fixed exchange rate regimes are more prone to financial vulnerabilities and more sensitive to the global financial shocks than more flexible regimes.

Second, the existing studies on CEE countries are deficient in one of two respects: either they include many countries, but allow for too few shocks (see, e.g., Shevchuk, 2014, who analyses 14 countries but has just two shocks) or vice versa, they allow for more shocks, but analyse just few countries (see, e.g., Stążka-Gawrysiak, 2009, who has four shocks but analyses just one country). Our approach overcomes these deficiencies: we allow for four shocks and examine eight CEE countries.[4] In their comparison we go beyond forecast error variance decomposition and find that output

[3] See also Yılmaz (2012) who found that the flexible exchange rate adopted in Turkey after 2001 acted as a shock absorber.

[4] A related approach was adopted by Jarociński (2010) to investigate the effects monetary shocks in four CEE countries and five euro area member states. His methodology was, however, different

reactions to real shocks have been indeed more muted in countries with more flexible exchange rates.

Third, the focus on relatively homogeneous group of CEE countries contributes to the reliability of the results obtained. A similar observation was exploited by Hegerty (2017) in his analysis of relations between inflation and exchange rate regimes in CEE countries. The point is that we compare countries which are similar in many respects and especially they belong to the group of EMEs, are small and open, share similar economic history, but have adopted different exchange rate regimes. Such circumstances can be considered a natural experiment which our analysis inquiries into.

Fourth, an important factor that enhances the reliability of our results is the econometric methodology adopted. Its main strength is that we do not have to rely on a single specification of an empirical model but can take into account results obtained from the most probable specifications through the Bayesian model averaging technique. Moreover, our approach allows for serial correlation that is common to variables included in the analysis. This makes empirical models more parsimonious and improves the short-run structural analysis (see, e.g., Centoni and Cubadda, 2011). Both these features of our approach are important since we found out that the model posterior probability is diffused and the data favour models with common serial correlation between analysed variables. We use sign and zero restrictions to identify the structural shocks and we employ the method proposed by Arias et al. (2018) to draw from the posterior distribution of the model parameters.

The rest of the paper is structured as follows. The next section lays out theoretical model that underlies our analysis. It is used to derive restrictions that are imposed on structural VAR models in the empirical part. Empirical methodology is presented in Section 3. The details of the division of CEE countries into peggers and floaters together with the data description are discussed in Section 4. Main empirical results are reported in Section 5. In the last section we offer conclusions.

2 Theoretical issues

The seminal papers of Fleming (1962) and Mundell (1963), published well before the demise of the Bretton Woods system, have reignited economists' interest in choices and consequences of exchange rate regimes. It was, however, only after this system broke down in the 1970s that their analyses gained great policy relevance (The Economist, 2016). Their main finding was that the choice of an exchange rate arrangement was

from ours. Not only did he investigate just one shock, but he also used a different algorithm to deal with sign restrictions.

not neutral to macroeconomic policy effectiveness unless capital flows were controlled. In a study summarizing the state of the art on the exchange rate regimes Ghosh et al. (2002, p. 30 ff) built a simple stylized model in the spirit of the Barro-Gordon approach and demonstrated that the floating exchange rate is preferable to pegged rate if shocks are real but worse if shocks are nominal. This conclusion rests on the assumption of high capital mobility. If capital mobility is low, e.g. due to capital controls, aggregate demand shocks are partly offset under fixed exchange rate, whereas floating rate amplifies such shocks (Ghosh et al., 2002, p. 26). While their model is neat and elegant, it allows for two sources of shocks only: (real) productivity shocks and (nominal) monetary shocks. Although it is just enough to convey the main theoretical point about insulating properties of the exchange rate regime it seems too parsimonious to be used as a theoretical framework in empirical research.

A fully microfounded model of a small open economy with sticky prices was developed by Galí and Monacelli (2005). Using its tractable version they analysed properties of three alternative monetary regimes: a domestic inflation-based Taylor rule, a CPI-based Taylor rule and an exchange rate peg. They found that there is a trade-off between the stabilization of nominal exchange rate and output gap. Their conclusion was that 'an exchange rate peg generates substantially higher welfare losses than a Taylor rule, due to the excess smoothness of the terms of trade that it entails' (Galí and Monacelli, 2005, p. 727). In other words, the real exchange rate was excessively smooth under a fixed exchange rate regime and prices could not compensate for the constancy of the nominal exchange rate since they were sticky (Galí, 2015). In comparison to the model developed by Ghosh et al. (2002) the (dynamic) model with fully specified microeconomic foundations and stochastic shocks is – due to its sophisticated theoretical structure – on the other extreme of a spectrum of economic models. Their common feature, however, is that they are rather inconvenient as a theoretical framework for empirical application.[5]

A tractable macroeconomic model of a small open economy hit by two real shocks was offered by Rødseth (2000, pp. 325-336). He examined output variability under an exchange rate target and under a price level target, assuming that in both cases a central bank used the interest rate appropriately. This two ways of setting monetary policy correspond to fixed and floating exchange rate regimes, respectively. Rødseth (2000, p. 331) demonstrated that a floating exchange rate regime was superior to a fixed exchange rate in preserving output stability under a high volatility of demand shocks and low volatility of supply shocks and vice versa. One should, however, be

[5]It is quite common in the literature that DSGE models are calibrated and not estimated. In fact, that was the method adopted by Galí and Monacelli (2005). For more on the problems related to DSGE modelling see, e.g., Blanchard (2016).

careful with interpretation of these findings. First, his definition of a demand shock was quite broad: it encompassed 'a genuine demand shock' as well as changes to foreign variables and a stochastic risk premium (Rødseth, 2000, p. 319). As such it included both real and financial shocks. Second, Rødseth (2000, p. 332) admitted that some output variability might be desirable in the face of supply (productivity) shocks. Thus, even if shocks are predominantly supply in origin, the floating rate regime is more desired.[6]

The framework that allows for four types of shocks was used by Agénor and Montiel (2008). They employed the model developed by Genberg (1989) to examine the relation between the extent to which monetary policy reacts to changes in the exchange rate and output variability. Adopting the fixed exchange rate was an optimal choice when shocks originated exclusively from the domestic money market, whereas exchange rate flexibility was preferable when an economy was hit by demand, financial and/or supply shocks. Thus, they concluded that 'from the perspective of providing automatic stabilization to domestic output, fixed exchange rates will rarely be optimal' (Agénor and Montiel, 2008, p. 308).

All shocks in their model, however, were purely transitory and due to an assumed lack of any nominal rigidities an economy remained continuously in the flexible-price equilibrium. Free of these flaws is the macroeconomic model of an open economy developed by Obstfeld (1985) and subsequently extended by Clarida and Galí (1994). Their model was used in empirical work, among others, by Mumtaz and Sunder-Plassmann (2013). The model we use in this paper is, in principle, taken from Clarida and Galí (1994), but we extend it to allow for financial shocks.

The model consists of four underlying equations. The first two describe equilibria in the goods market and money market with the conventional IS and LM relations:

$$y_t^d = d_t - \eta\left(s_t + p_t\right) - \sigma\left[i_t - E_t\left(p_{t+1} - p_t\right)\right], \qquad \eta, \sigma > 0, \tag{1}$$

$$m_t^s = p_t + y_t - \lambda i_t, \qquad \lambda > 0. \tag{2}$$

Each variable represents a log-difference between domestic and foreign levels, so for example y_t^d is the relative demand and equals the (log of the) domestic demand minus the (log of the) foreign demand. The exception is the nominal interest rate differential which is based on plain levels, i_t. The sum of the nominal exchange rate, s_t, which is defined as a price of domestic currency in terms of a foreign currency, and the relative price level, p_t, is equal to the real exchange rate, q_t. The relative money supply is denoted as m_t^s and the relative demand disturbance as d_t.

[6] This was so if the price elasticity of aggregate demand was greater than one (Rødseth, 2000, p. 333). Taking into account that imperfect competition models require firms to operate on the elastic segment of demand, the condition does not seem to be too restrictive.

The equilibrium in the foreign exchange market is represented with the UIP equation which we extend to allow for the stochastic risk premium term, x_t:

$$i_t = -\left(E_t s_{t+1} - s_t\right) + x_t. \tag{3}$$

This extension is important as it enables us to introduce explicitly financial shocks into the model. In the original model these shocks have not been separated from demand shocks. This in turn undermines economic interpretation of empirical results because the restrictions used to identify demand shocks are not the same as those needed to identify financial shocks (see Dąbrowski, 2012a).

The fourth equation, the PS relation, is used to capture the price stickiness: the price level is an average of the flexible-price equilibrium level, p_t^e, and the level expected in $t-1$ period to clear the goods market in time t, $E_{t-1}p_t^e$:

$$p_t = \theta p_t^e + (1-\theta)\, E_{t-1} p_t^e, \qquad 0 < \theta < 1. \tag{4}$$

Due to price stickiness the adjustment process to shocks hitting an economy is not instantaneous, so the flexible-price equilibrium is attained only in the long term.

Four stochastic processes, i.e. the relative supply of output, y_t^s, the relative demand disturbance, d_t, the relative stock of money, m_t, and the risk premium, x_t, are defined as follows:

$$h_t = h_{t-1} + u_t, \tag{5}$$

where $h_t = [y_t^s, d_t, m_t, x_t]'$. A vector of structural shocks u_t includes supply, demand, financial and monetary shocks $\left[u_t^s, u_t^d, u_t^f, u_t^m\right]'$.[7]

The long-run solution of the model can be obtained recursively starting with the observation that in the flexible-price equilibrium $y_t^e = y_t^s$ (see Clarida and Galí, 1994). The real exchange rate adjusts to keep the goods market in equilibrium, the price level changes in order to maintain equilibrium in money market and the interest rate follows the changes in the risk premium. Concisely, in equilibrium:[8]

$$z_t^e = A h_t = z_{t-1}^e + A u_t, \tag{6}$$

[7]Each shock has a permanent component only. A transitory component can be added in a straightforward way. Then $h_t = h_{t-1} + u_t + \Gamma u_{t-1}$, where Γ is a diagonal matrix with positive and less than unity $\gamma_s, \gamma_d, \gamma_f, \gamma_m$ on the main diagonal.

[8]In a model with both permanent and transitory shocks the 'long-run' flexible-price equilibrium corresponds to the state in which transitory shocks have already died out. Thus, it should be defined in terms of $E_t z_{t+j}^e = z_{t-1}^e + A(I-\Gamma)u_t$, for $j \geq 1$ (assuming that $u_{t-1} = 0$), rather than z_t^e. The point is that the flexible-price *equilibrium* levels change in two steps, from z_{t-1}^e to z_t^e and then to $E_t z_{t+1}^e$, because additional period is needed for transitory components to die out. The sign restrictions based on the solution for z_t^e remain the same.

Table 1: Model-based long-run identifying restrictions

Variable	Shock			
	supply	demand	financial	monetary
Relative output	+	0	0	0
Real interest rate differential	−	+	+	0
Real exchange rate	−	+	−	0
Relative price level	−	+	+/−	+

Note: The reaction of the relative price to financial shock depends on the exchange rate regime: it is positive if the exchange rate is floating and negative the rate is pegged. In the empirical part this sign restriction is not imposed.

where $z_t^e = [y_t^e, r_t^e, q_t^e, p_t^e]'$ and A is a matrix that includes the long-run multipliers:

$$A = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ -\frac{1}{\eta} & \frac{1}{\eta} & -\frac{\sigma}{\eta} & 0 \\ -1 & 0 & \lambda & 1 \end{bmatrix}. \tag{7}$$

The solution is the same under both flexible and fixed exchange rate regimes. The only difference is that for the latter regime the equilibrium price level is

$$p_t^e = q_t^e - \bar{s}, \tag{8}$$

where $\bar{s}$ is the level which the nominal exchange rate is pegged to. The difference stems from the fact that under fixed rate regime the required adjustment in the real exchange rate needs to be attained via the price level changes. Thus, the last row in matrix A under the fixed rate regime is simply $\left[-\frac{1}{\eta}, \frac{1}{\eta}, -\frac{\sigma}{\eta}, 0\right]$.

The matrix A can be used to justify the long-term restrictions we impose on the reactions of flexible-price equilibrium levels to structural shocks. Table 1 summarizes these restrictions. For example, the long-run reaction of the real exchange rate is positive to a demand shock ($\frac{1}{\eta} > 0$), negative to supply and financial shocks ($-\frac{1}{\eta} < 0$ and $-\frac{\sigma}{\eta} < 0$, respectively) and nil to a monetary shock.

The restrictions depicted in Table 1 are not as strict as those suggested in equation (7). First, according to that equation the real interest rate is driven exclusively by financial shocks in the long term. We are a bit skeptical about such a strong assertion: it would be appropriate only in a limiting case of an extremely open economy and perfect capital mobility. It is straightforward to demonstrate that in a closed economy case the real interest rate reacts negatively to supply shocks and positively to demand shocks. Thus, taking into account that CEE economies are open but capital mobility is less than perfect, we decided not to impose possibly oversimplifying zero-restrictions on the long-term reactions of the real interest rate to supply and demand shocks.

Second, the definition of matrix A implies that the equilibrium price level under floating exchange rate is unaffected by a demand shock in the long run. The required adjustment is attained via a change in the nominal exchange rate only. Such a restriction, however, would be far-fetched as it is justified for clean floats only which are rather poor description of actual exchange rate arrangements.

Third, if the exchange rate is fixed, then the stock of money is no longer exogenous. Monetary shocks, however, are still possible, e.g. changes in money multiplier, although their impact is absorbed by an instantaneous adjustment in monetary base. Thus, one can argue that the relative price level is independent of monetary shocks. Such a claim, however, would be valid under the assumption that the exchange rate is perfectly fixed, which simply seems to be inappropriate with respect to actual pegs. As we will explain below, CEE countries are somewhere between the two extremes of the exchange-rate-regime spectrum rather than exactly at one of the poles.

3 Empirical strategy

The analysis starts with an $n-$dimensional stable Gaussian VAR(k) process:

$$\Delta y_t = A_1 \Delta y_{t-1} + A_2 \Delta y_{t-2} + \cdots + A_k \Delta y_{t-k} + \Phi D_t + \varepsilon_t,\ \{\varepsilon_t\} \sim iiN(0, \Sigma),\ t = 1, 2, \ldots, T, \tag{9}$$

where Σ is a PDS matrix, $\{\varepsilon_t\}_{t=1}^{T}$ form a Gaussian white noise process with a covariance matrix Σ, D_t collects deterministic components and the starting points $y_{-k}, y_{-k+1}, \ldots, y_0$ are treated as known. Matrices A_1 through A_k and Φ stand for parameters of the considered VAR(k) process.

The matrix form of the process (9) reads as follows:

$$Y = X\Gamma + Z\Gamma_d + E = \begin{pmatrix} X & Z \end{pmatrix} \begin{pmatrix} \Gamma \\ \Gamma_d \end{pmatrix} + E = \tilde{X}\tilde{\Gamma} + E, \tag{10}$$

where $\Gamma_{nk \times n} = \begin{pmatrix} A_1 & A_2 & \ldots & A_k \end{pmatrix}'$, $\Gamma_d = \Phi'$, $Y_{T \times n} = \begin{pmatrix} \Delta y_1 & \Delta y_2 & \ldots & \Delta y_T \end{pmatrix}'$, $x_t = \begin{pmatrix} \Delta y'_{t-1} & \Delta y'_{t-2} & \ldots & \Delta y'_{t-k} \end{pmatrix}'$, $X_{T \times nk} = \begin{pmatrix} x_1 & x_2 & \ldots & x_T \end{pmatrix}'$, $E_{T \times n} = \begin{pmatrix} \varepsilon_1 & \varepsilon_2 & \ldots & \varepsilon_T \end{pmatrix}'$, $Z_{T \times l} = \begin{pmatrix} D_1 & D_2 & \ldots & D_T \end{pmatrix}'$, l denotes the number of deterministic components, and $\tilde{X} = \begin{pmatrix} X & Z \end{pmatrix}$, $\tilde{\Gamma}' = \begin{pmatrix} \Gamma' & \Gamma'_d \end{pmatrix}$.

The parameters of the above-presented process will be estimated with the use of

Bayesian methodology. To complete the definition of the considered Bayesian VAR(k) model we have to impose prior distributions for the parameters. We have decided to employ the same Normal-inverted Wishart prior structure as in Dąbrowski and Wróblewska (2016):

1. $\Sigma_{n\times n} \sim iW(S, q_\Sigma)$, where S is a PDS matrix and $q_\Sigma \geq n$,
2. $\Gamma_{nk\times n}|\Sigma \sim mN(0, \Sigma, \underline{\Omega}_\Gamma)$, where $\underline{\Omega}_\Gamma$ is a PDS matrix of order nk,
3. $\Gamma_d|\Sigma \sim mN(0, \Sigma, \underline{\Omega}_d)$, where $\underline{\Omega}_d$ is a PDS matrix of order l.

The above stated prior distributions for Γ and Γ_d lead to the following matrix normal prior $\tilde{\Gamma}_{nk+l\times n}|\Sigma \sim mN(0, \Sigma, \underline{\Omega})$, where $\underline{\Omega} = \begin{pmatrix} \underline{\Omega}_\Gamma & 0 \\ 0 & \underline{\Omega}_d \end{pmatrix}$. In the presented research $\underline{\Omega}$ is of the form $\begin{pmatrix} \frac{\nu_\Gamma}{nk} I_{nk} & 0 \\ 0 & \nu_d I_l \end{pmatrix}$, where the parameters ν_Γ and ν_d are estimated ($\nu_\Gamma \sim iG(s_\Gamma, n_\Gamma)$, $\nu_d \sim iG(s_d, n_d)$, $iG(s_., n_.)$ denotes an inverted Gamma distribution with parameters $s_.$ and $n_.$, i.e. $p(\nu_.) \propto \nu_.^{-n_.-1} \exp(-\frac{s_.}{\nu_.})$), so the hierarchical prior structure is applied (see, e.g., Koop et al., 2010).

The joint prior distribution is truncated by the stability condition imposed on the VAR parameters.

The assumed distributions belong to the so called conjugate priors family. It means that the posterior distributions are of the same form:

1. $\Sigma|., Y \sim iW(S + E'E + \tilde{\Gamma}'\underline{\Omega}^{-1}\tilde{\Gamma}, q_\Sigma + nk + l + T)$, where $E = Y - \tilde{X}\tilde{\Gamma}$'
2. $\tilde{\Gamma}|., Y \sim mN(\overline{\mu}_{\tilde{\Gamma}}, \Sigma, \overline{\Omega})$, where $\overline{\Omega} = (\underline{\Omega}^{-1} + \tilde{X}'\tilde{X})^{-1}$, $\overline{\mu}_{\tilde{\Gamma}} = \overline{\Omega}\tilde{X}'Y$,
3. $\nu_\Gamma|., Y \sim iG(s_\Gamma + \frac{1}{2}tr(nk\Sigma^{-1}\Gamma'\Gamma), n_\Gamma + \frac{n^2k}{2})$,
4. $\nu_d|., Y \sim iG(s_d + \frac{1}{2}tr(\Sigma^{-1}\Gamma_d'\Gamma_d), n_d + \frac{nl}{2})$.

Additionally we consider models allowing for common serial correlation of the analysed series which, in the framework of VAR, leads to the reduced rank restriction imposed on the parameters A_1 through A_k (see e.g. Engle and Kozicki, 1993 and Vahid and Engle, 1993, for the survey see e.g. Centoni and Cubadda, 2011):

$$\Delta y_t = \gamma\delta_1\Delta y_{t-1} + \gamma\delta_2\Delta y_{t-2} + \cdots + \gamma\delta_k\Delta y_{t-k} + \Phi D_t + \varepsilon_t,\ \{\varepsilon_t\} \sim iiN(0, \Sigma),\ t = 1, 2, \ldots, T, \tag{11}$$

where meaning of the symbols Σ, $\{\varepsilon_t\}_{t=1}^T$, D_t, Φ is left the same as in the process (9). The matrix form of the process (11) reads as follows:

$$Y = X\delta\gamma' + Z\Gamma_d + E, \tag{12}$$

where meaning of Γ_d, $Y_{T\times n}$, $X_{T\times nk}$, $E_{T\times n}$, $Z_{T\times l}$, l is left unchanged (see the explanation under equation (10)). We assume that matrices $\gamma_{n\times(n-s)}$ and $\delta_{nk\times(n-s)}$ are of full column rank. In the present model the matrix Γ introduced in equation (10) is of reduced rank equal to $n-s$ ($\Gamma=\delta\gamma'$), s denotes the number of common features.

In models with the reduced rank parameters we have to deal with the problem of the parameters non-identification which is similar to that one known from the error correction modeling, so the solutions known from VEC models can be adopted.

Specifically, for any non-singular matrix M of order $n-s$ products $\delta\gamma'$ and $\delta MM^{-1}\gamma'$ are equivalent, so we have the problem with identification of matrix parameters γ and δ. To overcome this ambiguity we have decided to adopt the method proposed by Koop et al. (2010) for the VEC models, also employed in Dąbrowski and Wróblewska (2016).

The over-mentioned algorithm switches between two parameterisations of the considered product:

$$\delta\gamma'=\delta O_\Gamma^{-1}O_\Gamma\gamma'\equiv DG', \tag{13}$$

where O_Γ is an $n-s\times n-s$ symmetric positive definite matrix. In the left-hand of (13) it is assumed that δ has orthonormal columns with positive elements in the first row whiles the matrices on the right-hand are left free, i.e. $G\in\mathbb{R}^{n(n-s)}$ and $D\in\mathbb{R}^{nk(n-s)}$. We can now write model (12) in the $G-D$ parameterisation:

$$Y=XDG'+Z\Gamma_d+E=\begin{pmatrix} XD & Z\end{pmatrix}\begin{pmatrix} G' \\ \Gamma_d\end{pmatrix}+E=\tilde{X}_D\tilde{\Gamma}_G+E, \tag{14}$$

where $\tilde{X}_D=\begin{pmatrix} XD & Z\end{pmatrix}$, $\tilde{\Gamma}'_G=\begin{pmatrix} G & \Gamma'_d\end{pmatrix}$. For G and D we settle matrix normal priors of the following form:

1. $D\sim mN(0,\frac{1}{nk}I_{n-s},I_{nk})$, which leads to non-informative prior for δ and for the space spanned by it (see Chikuse, 2002),
2. $G|\nu_G\sim mN(0,\nu_G I_{n-s},\Sigma)$,
3. $\nu_G\sim iG(s_G,n_G)$.

The priors for the remaining parameters are left unchanged. It is easy to see that $\tilde{\Gamma}_G|\Sigma,\nu_G,\nu_d\sim mN(0,\Sigma,\underline{\Omega}_G)$, where $\underline{\Omega}_G=\begin{pmatrix}\nu_G I_{n-s} & 0\\ 0 & \nu_d I_l\end{pmatrix}$.

Similarly to VAR models, the joint prior is truncated by the stability condition.

One of the advantages of the Koop et al. (2010) method is the possibility of using the Gibbs sampler for the simulation from the posterior distribution, because we know the full conditional posteriors (for the $D-G$ parameterisation):

1. $\Sigma|., Y \sim iW(S+E'E+\frac{1}{\nu_G}GG'+\frac{1}{\nu_d}\Gamma_d'\Gamma_d, q_\Sigma+n-s+l+T)$,
2. $G|., Y \sim mN(vec(\overline{\mu}_G), \overline{\Omega}_G, \Sigma)$, where $\overline{\Omega}_G = (\frac{1}{\nu_G}I_{n-s}+D'X'XD)^{-1}$, $\overline{\mu}_G = (Y-Z\Gamma_d)'XD\overline{\Omega}_G$,
3. $vec(D)|., Y \sim N(\overline{\mu}_{vD}, \overline{\Omega}_{vD})$, where $\overline{\Omega}_{vD} = ((G'\Sigma^{-1}G\otimes X'X)+(nkI_{n-s}\otimes I_{nk}))^{-1}$, $\overline{\mu}_{vD} = \overline{\Omega}_{vD}vec(X'(Y-Z\Gamma_d)\Sigma^{-1}G)$,
4. $\Gamma_d|., Y \sim mN(vec(\overline{\mu}_d), \Sigma, \overline{\Omega}_d)$, where $\overline{\Omega}_d = (\frac{1}{\nu_d}I_l+Z'Z)^{-1}$, $\overline{\mu}_d = \overline{\Omega}_dZ'(Y-XDG')$,
5. $\nu_G|., Y \sim iG(s_G+\frac{1}{2}tr(G'\Sigma^{-1}G), n_G+\frac{n(n-s)}{2})$,
6. $\nu_d|., Y \sim iG(s_d+\frac{1}{2}tr(\Sigma^{-1}\Gamma_d'\Gamma_d), n_d+\frac{nl}{2})$.

Samples from the posterior distributions of δ and γ can be obtained by using transformations: $\delta = D(D'D)^{-\frac{1}{2}}O$ and $\gamma = G(D'D)^{\frac{1}{2}}O$, where $O = diag(\pm 1)$ helps to obtain positive elements in the first row of δ.

The shocks are identified *via* zero and sign restrictions with the help of the method proposed by Arias et al. (2018).

To obtain the marginal data density, needed for the model comparison we have to integrate the parameters. Some of them can be integrated analytically (Γ in the model (10), G in the model (14) and Γ_d, Σ in both models), which leads us to the following results:

- the data density conditional on ν_Γ and ν_d in the VAR model (10)

$$\begin{aligned} p(Y|\nu_\Gamma,\nu_d) &= \pi^{-\frac{nT}{2}}\prod_{i=1}^{n}\frac{\Gamma[(q_\Sigma+T+1-i)/2]}{\Gamma[(q_\Sigma+1-i)/2]}|S|^{\frac{q_\Sigma}{2}}|\underline{\Omega}|^{-\frac{n}{2}}|\overline{\Omega}|^{\frac{n}{2}}\times \\ &\times\ |S+Y'M_{\tilde{X}}Y+\hat{\Gamma}'\tilde{X}'\tilde{X}\overline{\Omega}\underline{\Omega}^{-1}\hat{\Gamma}|^{-\frac{q_\Sigma+T}{2}}, \end{aligned} \tag{15}$$

  where $M_{\tilde{X}} = I_T - \tilde{X}(\tilde{X}'\tilde{X})^{-1}\tilde{X}'$, $\hat{\Gamma} = (\tilde{X}'\tilde{X})^{-1}\tilde{X}'Y$ and $\Gamma(\alpha)$ is the gamma function, that is the function defined by the integral: $\Gamma(\alpha) = \int_0^\infty x^{\alpha-1}\exp(-x)\mathrm{d}x$ for $x>0$ (see e.g. Bauwens et al., 1999);

- the data density conditional on D, ν_G and ν_d in the VAR models with common serial correlation (14):

$$\begin{aligned} p(Y|D,\nu_G,\nu_d) &= \pi^{-\frac{nT}{2}}\prod_{i=1}^{n}\frac{\Gamma[(q_\Sigma+T+1-i)/2]}{\Gamma[(q_\Sigma+1-i)/2]}|S|^{\frac{q_\Sigma}{2}}|\underline{\Omega}_G|^{-\frac{n}{2}}|\overline{\Omega}_G|^{\frac{n}{2}}\times \\ &\times\ |S+Y'M_{\tilde{X}_D}Y+\hat{\Gamma}_G'\tilde{X}_D'\tilde{X}_D\overline{\Omega}_G\underline{\Omega}_G^{-1}\hat{\Gamma}_G|^{-\frac{q_\Sigma+T}{2}}, \end{aligned} \tag{16}$$

where $M_{\tilde{X}_D} = I_T - \tilde{X}_D(\tilde{X}_D'\tilde{X}_D)^{-1}\tilde{X}_D'$, $\hat{\Gamma}_G = (\tilde{X}_D'\tilde{X}_D)^{-1}\tilde{X}_D'Y$ and $\overline{\Omega}_G = (\tilde{X}_D'\tilde{X}_D + \underline{\Omega}_G^{-1})^{-1}$.

To obtain marginal data density in the compared models, we have to integrate ν_Γ, ν_G, ν_d and D from the above stated equations, for which we employ the arithmetic mean estimator.

In our analysis we impose the following prior hyperparameters $S = 0.01I_n$, $q_\Sigma = n + 2$, $s_. = 2$, $n_. = 3$ therefore $E(\nu_.) = 1$, $D(\nu_.) = 1$.

4 Data and preliminary analysis

In our approach four variables are used to built structural VAR models: the relative output, the real interest rate differential, the real exchange rate and the relative price level. We construct these variables using quarterly data on real GDP, three-month money market nominal interest rate and average nominal exchange rate spanning 1998q1 to 2015q4 and monthly data on harmonized index of consumer prices for an analogous period. The data have been collected mainly from the the Eurostat database. Their description and sources are depicted in Table A1 in the Appendix.

The (log of the) real GDP is used as a measure of output. The real interest rate is calculated as a difference between the three-month money market nominal interest rate and the actual HICP inflation. The (log of the) real exchange rate is based on the average quarterly nominal exchange rate defined as the price of national currency in terms of the euro, so its rise means an appreciation of the domestic currency. The price level is measured with a harmonized index of consumer prices (the same index is used to calculate inflation rate). The relative output and relative price level are constructed as the log-differences between domestic and foreign (euro area) variables. The real interest rate differential is the difference between domestic and foreign rates.

Our main objective is to examine whether the floating exchange rate insulates output against economic shocks to a greater extent than the fixed exchange rate. Thus, we need to divide eight CEE countries in our sample into two groups: peggers and floaters. The natural thing to do is to look at the de facto exchange rate regime classification. We focus on two popular classifications: the one published by the IMF in the *Annual Reports on Exchange Arrangements and Exchange Restrictions* and the (updated) Reinhart-Rogoff classification (Ilzetzki et al., forthcoming). The results of this exercise are depicted in Table A2 and Figure A1 in the Appendix.

According to the IMF's classification out of eight CEE countries included in our sample only two can be considered as being close to opposite poles of an exchange rate regime spectrum and thus uncontroversial: Bulgaria with its currency board adopted in 1997 and Poland with free floating adopted in 2000. Interestingly, all CEE countries managed their exchange rates in the late 1990s, but by the mid-2000s clear differences between them had emerged. Croatia, Slovenia and Slovakia decided to shift closer to the hard peg option and Bulgaria had already been there. Other countries, i.e. the Czech Republic, Hungary, Poland and Romania, moved in the opposite direction allowing their exchange rates to be more flexible. The reading of the IMF's *Annual Reports on Exchange Arrangements and Exchange Restrictions* makes it reasonable, therefore, to treat the first group as 'peggers' and the second group as 'floaters.'

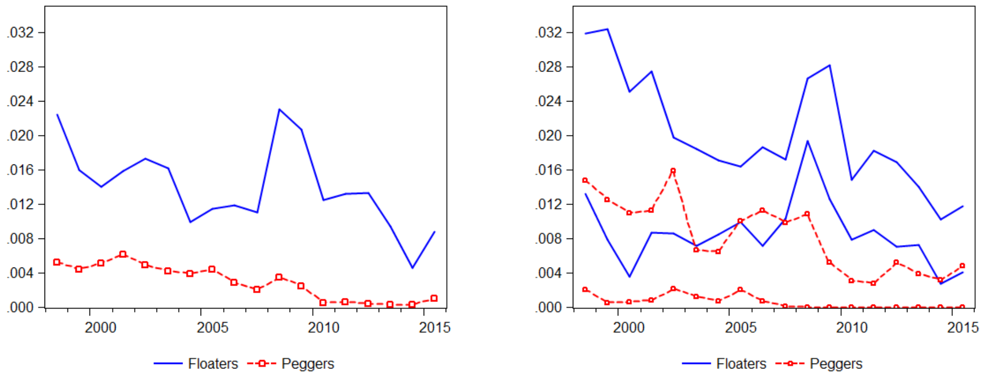
We realize that such a division is imperfect, since for example one can argue that Hungary can be considered a soft pegger rather than a floater till 2008 and as such more similar to Croatia than to the Czech Republic. Thus, in order to provide stronger arguments for our division we checked it against the updated Reinhart-Rogoff classification. Basically, it includes four broad categories: peg, limited flexibility, managed floating and freely floating.[9] In order to provide a full picture of the evolution of exchange rate regimes in CEE countries in 1998-2015 we used the fine Reinhart-Rogoff classification in Figure A1. The lighter the colour, the more flexible the exchange rate regime was. In principle, our division was confirmed: the light grey region can be observed in rows that correspond to our floaters, whereas the region that corresponds to our peggers is in dark grey (or in black).

In order to strengthen our argument the degree of variability of the actual exchange rates is examined. It was defined as the average absolute monthly change of the exchange rate against the euro. The results are presented in Figure 1. In the left-hand panel the median for each group is depicted. It is clear that the median for peggers is well below that one for floaters. In the right-hand panel minimum and maximum variabilities are illustrated. The floaters indeed experienced a greater exchange rate variability than the peggers and the explicit overlap between these two groups can only be observed at the turn of the centuries (1998-2002) and in 2006. The first overlap was mainly due to a relatively low variability of the Hungarian forint at the beginning of the century which is consistent with both exchange rate regime classifications discussed above. The second overlap was related to a gradual appreciation of the Slovak koruna after its entrance into the ERM II which resulted in its variability above that characteristic for other peggers.[10]

Finally, in other studies a similar division of CEE countires into floaters and peggers can be found (see, e.g., Harkmann and Staehr, 2018, Nucu and Anton, 2018).

[9] In fact, their coarse classification includes two additional categories: 'freely falling' and 'dual market in which parallel market data is missing.' These, however, were not observed in CEE countries in 1998-2015. The only exception was Romania before 2001 with 'free falling' exchange rate due to high inflation.

[10] Moreover, till 2007 the Czech and Slovak korunas were quite close one another with respect to the exchange rate variabilities.



(a) Median

(b) Minimum and maximum

Figure 1: Exchange rate variability in CEE countries, 1998-2015

For instance, Nucu and Anton (2018) investigate the monetary condition index in the Czech Republic, Hungary, Poland and Romania, i.e. countires we consider floaters, arguing that the other CEE countries are 'too different in terms of exchange rate regime.'

Two more points require clarification before we turn to empirical results and both are related to openness to capital flows. First, in financially closed economies – as is known from the impossible trinity hypothesis – monetary authorities can retain monetary autonomy irrespective of the exchange rate regime adopted. Thus, even if the exchange rate is pegged, monetary authorities can adjust their policy in order to absorb shocks hitting an economy. In this way the output reaction to shocks, especially real ones, can be quite similar to that under the floating exchange rate regime. This, however, does not seem to be a problem in the group of CEE countries as the data in Figure A2 in the Appendix illustrate. Capital account openness is measured by the Chinn-Ito index that ranges from 0 to 1 (Chinn and Ito, 2006, 2008). Admittedly, the median index for peggers is below that one for floaters, but the difference is not large (see the left-hand panel of Figure A2). Moreover, medians for both groups have remained above the median for the group that includes all the countries at least since 2003, so both CEE peggers and floaters can be considered relatively open to capital flows. In the right-hand panel minimum and maximum indices are depicted. Again, there is not too much difference between two groups, although one can claim that the peggers were slightly lagging behind the floaters with respect to opening capital account.

Second, one should not be misled by the data depicted in Figure A2 into thinking that capital mobility is perfect in CEE countries, especially in those that recorded the Chinn-Ito index equal to one, e.g. Bulgaria, the Czech Republic, Hungary, Romania. The point is important because in Section 2 we argued that less than perfect capital mobility was a reason to refrain from imposing some zero-restrictions. We abide by such a conservative approach since the Chinn-Ito index is a normalized index and the value of 1 corresponds to the highest observed degree of financial openness and not to perfect capital mobility.[11]

5 Empirical results

The empirical analysis starts with the Bayesian model comparison. We assume equal prior probability of each specification. The set of compared models consists of non-nested specifications which may differ in the lag order (k) and the number of co-features (s). There is a constant in each model. Additionally, in models for Slovakia and Slovenia we include a dummy to account for the participation in the ERM II and the euro area.

Models with posterior probabilities higher than 0.05, i.e. the assumed equal prior probability, are displayed in Table 2. Among these models are only those with reduced rank structures, so the data strongly support the hypothesis of the existence of common serial correlation among the analysed variables. Further results are obtained with the advantage of the Bayesian model averaging technique employed in the set of models with the highest posterior probability.

The conventional analysis of insulating properties of floating exchange rate is based on the forecast error variance decomposition that is used to identify the proportions of variability accounted for by real and nominal shocks. In our approach real shocks include supply and demand shocks, whereas financial and monetary shocks are nominal shocks. In Tables A3 and A4 in the Appendix the sources of fluctuations of all four variables in our models are reported.

It is quite clear that irrespective of the exchange rate regime real, especially supply, shocks are behind output variability. A small difference between Bulgaria and other CEE countries in this respect dissipates at longer forecast horizons and at four-year horizon the contribution of real shocks is more than 99 per cent in all countries (not reported).[12] A similar finding was obtained for Turkey by Yılmaz (2012), who examined the consequences of a shift to 'the more flexible' exchange rate regime in 2001.

There is also little difference between peggers and floaters with respect to the

[11] For details on the construction and interpretation of the index see Chinn and Ito (2008).

[12] Results for other forecasting horizons are available from the authors upon request.

Table 2: Models with the highest posterior probability

Bulgaria			Czech Republic			Croatia			Hungary		
k	s	$p(M_{k,s}\|Y)$	k	s	$p(M_{k,s}\|Y)$	k	s	$p(M_{k,s}\|Y)$	k	s	$p(M_{k,s}\|Y)$
5	3	0.098	5	3	0.087	5	3	0.153	5	3	0.095
7	3	0.093	6	3	0.082	6	3	0.130	7	3	0.092
6	3	0.085	8	3	0.079	7	3	0.084	8	3	0.085
8	3	0.085	7	3	0.078	5	2	0.083	9	3	0.082
5	2	0.067	5	2	0.072	8	3	0.075	6	3	0.077
9	3	0.066	9	3	0.072	9	3	0.070	5	2	0.073
6	2	0.064	6	2	0.062				6	2	0.067
			7	2	0.056				5	1	0.057
			8	2	0.054						
			5	1	0.053						

Poland			Romania			Slovakia			Slovenia		
k	s	$p(M_{k,s}\|Y)$	k	s	$p(M_{k,s}\|Y)$	k	s	$p(M_{k,s}\|Y)$	k	s	$p(M_{k,s}\|Y)$
5	3	0.092	5	3	0.133	5	3	0.086	5	3	0.090
6	3	0.091	7	3	0.113	7	3	0.082	7	3	0.087
7	3	0.082	6	3	0.108	6	3	0.080	6	3	0.081
5	2	0.081	5	2	0.106	5	2	0.074	6	2	0.066
8	3	0.074	8	3	0.094	8	3	0.073	8	3	0.066
9	3	0.073	6	2	0.074	9	3	0.072	9	3	0.061
6	2	0.067	6	1	0.060	6	2	0.060	5	2	0.058
9	2	0.061	7	2	0.059	7	2	0.060	5	1	0.055
7	2	0.057	5	1	0.052	5	1	0.059			
6	1	0.055				8	2	0.051			
8	2	0.053									
5	1	0.052									

Note: Prior probability of each specification: $p(M_{(k,s)}) = \frac{1}{20}$, $M_{(k,s)}$ stands for VAR(k) with s co-features.

relative importance of real and nominal shocks to the real exchange rate variability. Both the floating and fixed rates are mainly driven by demand and financial shocks and the contribution of real shocks increases with the forecasting horizon, whereas that of nominal shocks goes down. It is worth emphasizing that the higher exchange rate variability in the group of floaters (see evidence presented in the previous section) cannot be explained with their stronger susceptibility to financial shocks as their contribution is very similar across floaters and peggers, around 43 per cent on average in both groups.

One can hardly see any important differences between floaters and peggers as far as the sources of variability of two remaining variables are concerned. In the case of the real interest rate differential some small differences between countries do not seem to be related in any systematic way to the exchange rate regime adopted. The variability of the relative price level is mainly accounted for by monetary shocks, albeit their contribution is lower in the group of peggers, especially in Bulgaria. At the same time the contribution of demand and financial shocks is greater in that country than in the group of floaters. In general, this observation fits theoretical characteristics of exchange rate regimes: the real exchange rate variability stems from fluctuations of the nominal exchange rate if it is flexible and from changes in the relative price level if the nominal exchange rate is kept fixed.

Overall, our general conclusion from the analysis of forecast error variance decompositions is that there is a lot of similarities between CEE countries in this respect. This finding can be interpreted as evidence against the claim that the floating exchange rate is heavily influenced by financial shocks that are subsequently transmitted into a real economy and as such is a propagator of instability. This is in line with Yılmaz (2012) who found for Turkey that it is the price level that is influenced by nominal shocks rather than the exchange rate and these shocks are not destabilizing for the economy.

In an attempt to examine the insulating properties of the flexible exchange rate one should not rely on the forecast error variance decomposition only. The point is that the forecast error variance contains information about the structure of variability and not about its magnitude and that by definition contributions of all shocks have to sum up to 100 per cent. The claim that the contribution of, for example, financial shocks is similar across CEE floaters and peggers can, therefore, be challenged with the argument that it is uncontroversial that under floating rate regime the nominal exchange rate variability is greater (see Figure 1), so the exchange rate flexibility creates favourable conditions for instability. The important question that arises here is whether the increased exchange rate variability moderates output reactions to shocks hitting an economy. The relevant tool, which we think – following Dąbrowski and

Wróblewska (2016) – is appropriate to answer that question, are impulse response functions of the relative output to structural shocks.

In Figure 2 the mean reaction of output to two real shocks is illustrated with solid lines and (the analogue of) the confidence interval (the posterior mean plus/minus posterior standard deviation) is depicted with broken lines.[13] To keep figure uncluttered countries are compared in pairs: one pegger (lines with squares) and one floater are presented in each panel. The reactions to supply and demand shocks are depicted on the left-hand and right-hand, respectively.

A closer inspection of impulse response functions enables us to make three observations. First, in general, there are more differences in the output reactions to supply shocks than to demand shocks. This observation, together with the finding that the former's contribution to output variability is much greater than that the latter's (see Table A3), implies that differences between peggers and floaters identified below are even more important.

Second, the response of output under a fixed exchange rate regime to at least one real shock is stronger than that under a floating rate regime in each of our four pairs: Croatia, Slovakia and Slovenia react more intensely to a supply shock than their floating-rate counterparts, i.e. Hungary, Poland and the Czech Republic, whereas Bulgaria reacts stronger to a demand shock than Romania. This observation is in line with the view that exchange rate flexibility can be useful in insulating output against real shocks.

Third, in three pairs the mean reaction of output under a pegged exchange rate is outside the confidence interval for the corresponding reaction under a floating exchange rate. The dissimilarity between reactions to a supply shock is observed in the Croatia-Hungary and Poland-Slovakia pairs. Much the same difference is between Bulgaria and Romania in their reactions to a demand shock, but not to a supply shock. The latter response is fairly similar in both countries as the mean reactions are within the counterpart's confidence interval. In the remaining pair it can be observed that the reaction of the Slovenian relative output to a supply shock is stronger than that of the Czech relative output, although both means are relatively close one another.

Even though the contribution of nominal shocks to output variability is almost nil, witness forecast error variance decomposition depicted in Table A3, we compare impulse response functions of output to these shocks. To conserve space the response functions to nominal shocks are reported just for one pair, Croatia and Hungary, in Figure 3. The additional and important reason is that for the other pairs the picture is very much the same: there is hardly any difference between floaters and

[13] Impulse response functions for other shocks and variables are available from the authors upon request.

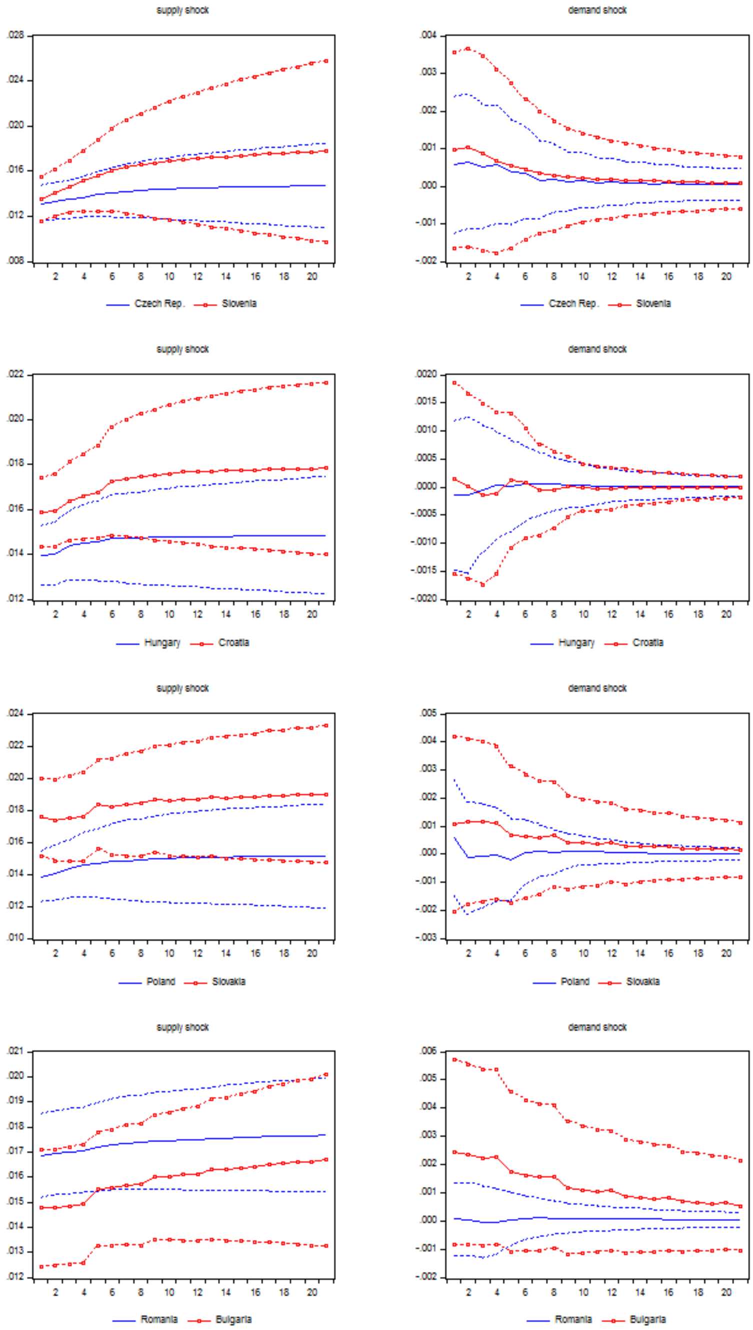


Figure 2: Impulse response functions of the relative output in CEE countries

peggers in this respect. The comparison, therefore, simply does not lend any support to the hypothesis that the increased exchange rate variability in the group of floaters resulted in stronger responses of output to nominal shocks. This is in line with finding by Jarociński (2010) that output responses to monetary shocks in CEE countries and the euro area countries are broadly similar (see especially Figure 4 in his study). Using a different approach (a panel regression framework) and a large sample of about 40 EMEs, Obstfeld et al. (forthcoming) demonstrate that it is fixed rates that make output more volatile in the face of global financial shocks and 'insulation properties afforded by flexible exchange rates can materially reduce the costs to EMEs from [such] shocks.' Similarly, Zeev (2019) finds that the exchange rate fixity has a 'negative effect on macroeconomic stability,' whereas Han and Wei (2018) offer evidence that the flexible rate regime insulates against tightening of foreign monetary policy, but not against its loosening.[14]

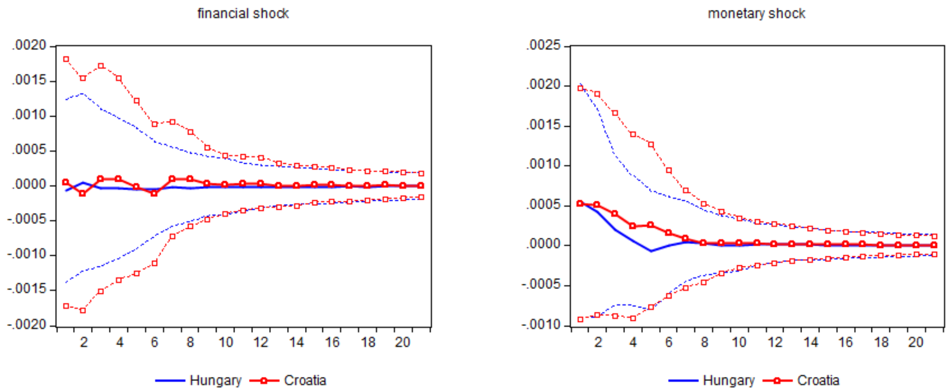


Figure 3: Impulse response functions of the relative output to nominal shocks in Croatia and Hungary

Interestingly, floaters and peggers differed with respect to the responses of the real exchange rate to all but monetary shocks. Again, to conserve space we report the response functions for one pair only. In Figure 4 reactions of the real exchange rate in Croatia and Hungary are compared. As is clear the mean responses of the exchange rate to demand and financial shocks, i.e. main drivers of its variability, in Hungary are much stronger than in Croatia. The means for Hungary are outside the confidence intervals of the corresponding reactions in Croatia. This pattern can also be observed

[14]See also Rohit and Dash (forthcoming) who provide evidence that the floating exchange rate regime insulates against the monetary spillover to a larger extent than less flexible regimes. It seems, however, that they compare advanced economies with EMEs rather than exchange rate regimes.

in the remaining pairs, although it is a bit weaker in the pair Bulgaria-Romania. The finding that CEE floaters experience the greater real exchange rate responsiveness to real shocks in CEE than peggers was also reported in Mirdala (2015).[15]

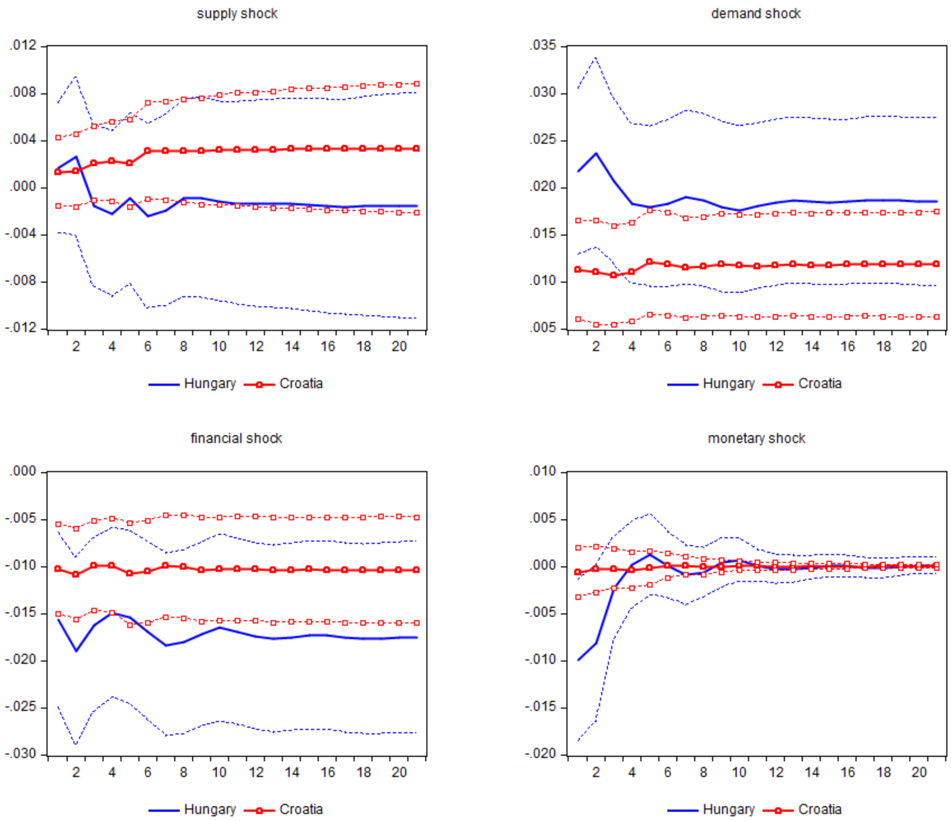


Figure 4: Impulse response functions of the real exchange rate in Croatia and Hungary

One can question our strategy of comparing peggers and floaters in pairs arguing that the changes in pairs would result in different conclusions. We make two things to justify our approach. First, we provide arguments in favour of our pairs, and second we run the robustness check.

Each pair is supposed to include a pegger and a floater. Using the updated Reinhart-Rogoff classification we confirmed that the exchange rate regime observed in a country considered 'a pegger' was indeed less flexible than in a counterpart

[15] His main conclusion about 'higher immediate absorption capabilities of fixed exchange rates' can hardly be considered well-founded, since he neither presents nor discusses the impulse responses of output.

country considered 'a floater.' We found only nine exceptions (6.4% of all pair-year observations): there was no difference between exchange rate regimes in Bulgaria and Romania in 2013-2015, the Czech Republic and Slovenia in 1998-2001 and in the remaining two pairs in 1998. When the fine Reinhart-Rogoff classification was used the results were even sharper: only six exceptions (4.3%), those at the beginning of our sample, survived. It should be emphasized that there was not a single pair-year observation in which our pegger had more flexible exchange rate regime than our floater.

Pairwise comparisons are the more effective, the more similar are compared countries with respect to main economic and institutional characteristics. Thus, basic macroeconomic characteristics were used to provide a closer look at our pairs. In Table A5 in the Appendix the data on income per capita, current account, CPI inflation, unemployment rate and absence of corruption were depicted. There are definitely some similarities between CEE countries. All countries can indeed be classified as middle-income countries with low inflation and moderate levels of unemployment, current account deficit and corruption. At the same time differences can be observed. Bulgaria and Romania are countries with the lowest income and highest inflation, whereas the Czech Republic and Slovenia are the wealthiest with the lowest inflation. The two former countries had the highest current account deficit and the two latter countries and Hungary had the lowest current account deficit. Croatia, Poland and Slovakia recorded the double-digit unemployment rates, whereas the Czech Republic, Romania and Slovenia maintained it below 7.5 per cent. The Czech Republic, Poland and Slovenia had the highest indices of the absence of corruption and the performance of Bulgaria and Romania was much less satisfactory.

Table 3 presents the rankings of CEE countries with respect to basic macroeconomic characteristics. The higher the country on the list the worse its performance in a given dimension. For instance, Bulgaria had the lowest income and Slovenia the highest and Croatia had the highest unemployment rate and the Czech Republic the lowest. Countries in bold are those that are located next to their counterparts in our pairs. For example, Bulgaria is in bold in column for income per capita as it is next to Romania, but Croatia is not in bold since it is separated from Hungary (its counterpart) by Poland. Admittedly, our pairs have not been perfectly confirmed by all macroeconomic criteria adopted – the ranking with respect to the CPI inflation is indeed perfect but the one according to the unemployment rate identifies just one our pair – but the general picture is quite consistent with our pairs.

In order to go beyond the rankings and use more effectively information on all basic macroeconomic characteristics (except for the absence of corruption due to data

Table 3: Rankings of CEE countries according to basic macroeconomic characteristics

Rank	Income per capita	Current account	CPI Inflation	Unemployment rate	Absence of corruption
1	**Bulgaria**	**Bulgaria**	**Romania**	Croatia	**Bulgaria**
2	**Romania**	**Romania**	**Bulgaria**	**Slovakia**	**Romania**
3	Croatia	**Poland**	**Hungary**	**Poland**	**Croatia**
4	Poland	**Slovakia**	**Croatia**	Bulgaria	**Hungary**
5	Hungary	Croatia	**Slovakia**	Hungary	**Slovenia**
6	Slovakia	Czech Rep.	**Poland**	Slovenia	**Czech Rep.**
7	**Czech Rep.**	Hungary	**Czech Rep.**	Romania	Poland
8	**Slovenia**	Slovenia	**Slovenia**	Czech Rep.	

Notes: The higher the country in the ranking, the worse its performance with respect to a given criterion. Bold is used for countries that are located next to their counterparts in our pairs.

availability) a simple clustering analysis was carried out.[16] The results of clustering with the Ward method are presented in Figure A3 in the Appendix. The highest average silhouette width (0.45) was for three clusters. Two of them were in line with our pairs (silhouette width in parentheses): Bulgaria-Romania (0.50), the Czech Republic-Slovenia (0.59), and the third one was more diversified: Croatia, Hungary, Poland, Slovakia (0.35). Again, even though these results are not perfectly in line with our pairs, the degree of agreement is considerable.[17]

The robustness check is the second thing we do to make our results more appealing. We compared the relative output reactions to real shocks in each country with analogous reactions in Poland. The latter country was chosen because its exchange rate was floated de jure in April 2000 and de facto in 1998 when the National Bank of Poland decided to refrain from foreign exchange market interventions. Moreover, out of CEE currencies it is the Polish zloty that has been floating for the longest time. In Figures A4 and A5 in the Appendix the output reactions of floaters and peggers were depicted against those of Poland, respectively.

The reactions of the relative output to supply shocks across floaters were rather similar: the mean reactions in the Czech Republic and Hungary were within the confidence interval for Poland and close to the mean of the latter. Romania, however, stood out from CEE floaters with its relatively high mean output response, especially

[16] We thank our colleague Sławomir Śmiech for his kind assistance on this analysis.

[17] One can buttress this claim with the observation that Slovakia and Croatia turned out to be the most similar in the last cluster, but it makes no sense to have them in one pair since both are the peggers. Out of two reasonable variants of pairs the first, i.e. Croatia-Hungary and Poland-Slovakia, has already been examined. The pair Croatia-Poland will be explicitly examined below as a part of a robustness check. The results for the pair Hungary-Slovakia will not be reported. They can, however, be inferred from Figure 2 and are in line with our main finding.

in the first six quarters. The output reactions to supply shocks of peggers turned out to be stronger than that of Poland: in Slovakia the mean response was well above the upper border of confidence interval for Poland, in Croatia – on the border and in Bulgaria and Slovenia – below the border, but above the mean response for Poland.

The similar pattern was observed for output responses to demand shocks. They were comparable in the group of floaters, but stronger in the group of peggers. The mean response of output in Bulgaria was above the upper bound of the confidence interval for Poland, whereas responses in Slovakia and Slovenia were within the confidence interval for Poland but above the mean. Croatia was the only pegger whose output reaction to a demand shock was similar in magnitude to that of Polish output.

6 Conclusion

In this paper we have examined the insulating properties of the floating exchange rate regime by comparing pegs and floats adopted in eight Central and Eastern European economies. Our findings can be summarized in four points. First, using two popular exchange rate classifications developed by the IMF and Ilzetzki et al. (forthcoming) we divided CEE countries into floaters and peggers and found out that the former indeed experienced a greater exchange rate variability than the latter. The difference was not related to capital account openness as it was relatively high and, more importantly, similar across both groups of CEE countries.

Second, we have used the model of a small open economy to carefully justify the restrictions imposed in empirical analysis in order to identify structural shocks. We found that irrespective of the exchange rate regime real, and in particular supply, shocks were behind the relative output variability in all CEE countries. The variability of the real exchange rate in turn was explained by demand and financial shocks. This finding alone was not sufficient to decide whether the flexible exchange rate was a shock propagator as suggested by the importance of nominal shocks or a shock absorber as suggested by the importance of real shocks.

Third, in order to remove ambiguity about the role of the flexible exchange rate we compared impulse response functions of the relative output to structural shocks in CEE countries. The comparison was between floaters and peggers. The output reactions to supply shocks in the former group turned out to be weaker than those of the latter group, whereas the responses to other shocks were not too different. Taking into account that it was supply shocks that were the main driver of output variability in all CEE countries, we conclude that evidence lends support to the hypothesis that the flexible exchange rate regime insulates the economy against real shocks to a greater extent than the fixed exchange rate regime. At the same time we did not find evidence of enhanced responsiveness of output to financial shocks in the group of floaters. Thus, the claim that the flexible exchange rate acts as a shock propagator should be treated as a theoretical possibility rather than an empirical regularity.

As far as policy implications are concerned a word of caution seems to be in place, especially as all CEE countries examined are formally obliged to join the euro area and Slovakia and Slovenia have already adopted the common currency. Our empirical results weigh in favour of the flexible exchange rate regime as it provides a partial insulation against shocks (see, e.g., Obstfeld et al., forthcoming, for the similar finding). It is not without reason, however, that we are economical with policy recommendations. The point is that our analysis contributes to *one* of the arguments used in the discussion on the choice of the exchange rate regime. Even though this argument is important, there are also others. For instance, the usefulness of autonomous monetary policy in CEE countries may be limited: indeed Nucu and Anton (2018) found that monetary decisions in the euro area had 'a prominent influence on monetary conditions' in four CEE countries we consider here as floaters. Another argument is on currency misalignments in real terms: Fidora et al. (2018) found that the euro area countries recorded smaller misalignment, albeit it was more persistent than in countries outside the euro area 'owing to the absence of a nominal adjustment channel.' Yet another argument is on transaction costs: Chen and Novy (2018), for example, found that currency unions lower trade costs and promote trade, although there is 'a significant amount of heterogeneity across country pairs' and the average trade effect of the euro is modest. Our results should, therefore, be interpreted narrowly: the flexible exchange rate regime insulates the economy against real shocks and does not act as a propagator of nominal shocks in CEE countries. As far as other criteria of exchange rate regimes comparison are concerned we still need more evidence and this looks like a promising avenue of further research.

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Appendix

Below we provide additional information in Tables and Figures that is summarized in the main text.

Table A1: Data description

Variable	Description	Source
Real GDP	Gross domestic product at market prices, chain linked volumes, index 2005=100, seasonally and calendar adjusted data; for Bulgaria (1998:1-1998:4) and Croatia (1998:1-1999:4) unadjusted data data from ESA 1995 (Tramo/seats method used for seasonal adjustment); for Poland (1998:1-2001:4) data from ESA 1995; for Slovakia seasonally adjusted data but not calendar adjusted data	Eurostat
Nominal interest rate	Three-month money market nominal interest rate; for Bulgaria (1998:1-1998:2 and 1999:1-1999:2) and Slovenia (1998:1)the deposit rate used; for Croatia (1998:1-2000:1) lending rate used; average of four adjacent quarters used for missing value for Hungary (2004:3).	Eurostat and IMF/IFS (for deposit and lending rates)
Nominal exchange rate	Quarterly average nominal exchange rate index (2005 = 100); an increase is an appreciation of domestic currency against the euro.	based on Eurostat data
Price level	Harmonized index of consumer prices (HICP); monthly data used to calculate quarterly averages.	Eurostat
Relative output	The log-difference between domestic and the euro area real GDPs.	based on Eurostat data
Real interest rate differential	The difference between domestic and euro are real interest rates. The real interest rate defined as a difference between nominal interest rate and actual HICP inflation.	based on Eurostat data
Real exchange rate	The (log of the) real exchange rate calculated as the nominal exchange rate corrected for price ratio; its rise means an appreciation of domestic currency against the euro in real terms	based on Eurostat data
Relative price level	The log-difference between domestic and euro area price levels.	based on Eurostat data

Table A2: Exchange rate regimes in Central and Eastern European countries, 2000-2015†

Bulgaria	Czech Rep.	Croatia	Hungary	Poland	Romania	Slovakia	Slovenia
Currency board July 1, 1997	Managed floating[a)]	Managed floating[a)] Sept 30, 1999	Crawling band	Crawling band	Managed floating[a)]	Managed floating[a)]	Managed floating[a)]
	Independent-ly floating June 30, 2001		Pegged band[b)] Oct 1, 2001	Independent-ly floating Apr 12, 2000	Crawling band June 30, 2001		Crawling band Feb 1, 2002
	Managed floating[a)] Jan 1, 2002				Managed floating[a)] Nov 2, 2004	Pegged band[b)] Nov 25, 2005	Pegged band[b)] June 27, 2004
	Independent-ly floating Jan 1, 2006	Conventional peg Sept 1, 2006	Independent-ly floating Feb 26, 2008				Currency union Jan 1, 2007
	Free floating* Apr 30, 2008	Stabilized arrangement* Apr 30, 2008	Free floating* Apr 30, 2008	Free floating* Apr 30, 2008	Floating* Apr 30, 2008		
	Other managed[c)] Nov 7, 2013	Other managed[c)] Jan 1, 2009	Floating Mar 1, 2009	Floating Sept 23, 2011		Currency union Jan 1, 2009	
	Stabilized arrangement Nov 19, 2013	Crawl-like arrangement Jun 17, 2010	Free floating Dec 31, 2011				

Notes: † A date below the arrangement corresponds to the introduction of the regime; if there is no date, the regime was adopted before 2000.
Formal categories in the AREAER: [a)] 'Managed floating with no predetermined path for the exchange rate.'
[b)] 'Pegged exchange rate within horizontal bands.' [c)] 'Other managed arrangement.' * No change in the regime: effective Feb 2, 2009 the classification has been changed retroactively to April 30, 2008, due to the revision of the classification methodology.
Source: based on the IMF *Annual Report on Exchange Arrangements and Exchange Restrictions*, various issues from 1999-2015.

Table A3: Forecast error variance decomposition of the relative output and real interest rate differential in CEE countries

Variable	Shock			
and country	supply	demand	financial	monetary
Relative output:				
Bulgaria	90.1	6.2	2.9	0.8
Czech Rep.	95.2	1.7	1.6	1.6
Croatia	97.4	0.9	0.9	0.7
Hungary	98.0	0.7	0.6	0.7
Poland	94.1	1.7	1.8	2.4
Romania	96.3	0.5	1.2	2.0
Slovakia	93.3	3.0	2.2	1.4
Slovenia	91.7	3.0	2.7	2.6
Averages:				
All	94.5	2.2	1.7	1.5
Pegs	93.2	3.3	2.2	1.4
Floats	95.9	1.1	1.3	1.7
Real interest rate differential:				
Bulgaria	4.4	18.1	73.8	3.7
Czech Rep.	3.2	31.8	63.1	1.8
Croatia	4.7	33.0	61.6	0.7
Hungary	2.7	29.5	65.6	2.2
Poland	3.7	50.7	43.5	2.1
Romania	3.6	17.9	64.9	13.6
Slovakia	3.7	29.9	63.8	2.6
Slovenia	5.8	37.1	54.7	2.4
Averages:				
All	4.0	31.0	61.4	3.6
Pegs	4.7	29.5	63.5	2.3
Floats	3.3	32.5	59.3	4.9

Note: Numbers are posterior expected values of forecast error variance decomposition expressed in percentage points at one-year forecast horizon.

Table A4: Forecast error variance decomposition of the real exchange rate and relative price level in CEE countries

Variable	Shock			
and country	supply	demand	financial	monetary
Real exchange rate:				
Bulgaria	7.0	50.2	34.2	8.6
Czech Rep.	4.0	52.2	39.9	3.9
Croatia	4.2	50.2	43.6	2.0
Hungary	4.6	51.1	34.1	10.1
Poland	9.8	44.2	42.7	3.3
Romania	2.7	62.5	29.6	5.1
Slovakia	5.4	47.5	42.1	5.0
Slovenia	5.6	51.5	40.5	2.4
Averages:				
All	5.4	51.2	38.4	5.1
Pegs	5.5	49.9	40.1	4.5
Floats	5.3	52.5	36.6	5.6
Relative price level:				
Bulgaria	7.0	28.2	18.5	46.3
Czech Rep.	6.0	7.6	6.8	79.7
Croatia	2.1	4.7	5.8	87.4
Hungary	4.4	12.2	15.8	67.6
Poland	4.3	6.3	3.7	85.7
Romania	7.0	3.5	20.8	68.6
Slovakia	6.7	7.9	11.1	74.3
Slovenia	5.4	5.2	4.9	84.4
Averages:				
All	5.4	9.4	10.9	74.2
Pegs	5.3	11.5	10.1	73.1
Floats	5.4	7.4	11.8	75.4

Note: Numbers are posterior expected values of forecast error variance decomposition expressed in percentage points at one-year forecast horizon.

Table A5: Basic macroeconomic characteristics of CEE countries, 2005-2015

Country	Income per capita[a)]	Current account[b)]	CPI Inflation[c)]	Unemployment rate[d)]	Absence of corruption[e)]
Bulgaria	14,888	-7.7	4.1	9.6	0.41
Czech Rep.	26,540	-1.8	2.1	6.4	0.54
Croatia	20,083	-2.6	2.4	13.0	0.63
Hungary	21,972	-1.6	3.8	8.9	0.57
Poland	20,905	-3.8	2.2	10.2	0.66
Romania	17,619	-6.2	4.9	6.8	0.51
Slovakia	24,414	-3.3	2.3	13.0	n.a.
Slovenia	28,371	0.5	2.1	7.3	0.60
Averages:					
All	21,849	-3.3	3.0	9.4	0.56
Pegs	21,939	-3.3	2.7	10.7	0.52
Floats	21,759	-3.3	3.2	8.1	0.59

Notes: [a)] Gross national income per capita converted to (constant 2011) international dollars using purchasing power parity rates.
[b)] In percent of GDP. [c)] The annual percentage change of consumer price index. [d)] In percent of the labor force (International Labour Organization estimate). [e)] One of the subindices of the World Justice Project Rule of Law Index that measures the extent to which countries adhere to the rule of law in practice. It ranges from 0 (the lowest score) to 1 (the highest score).

Source: all data from the *World Development Indicators* database except for the absence of corruption index that is from the World of Justice Project website: `www.worldjusticeproject.org`.

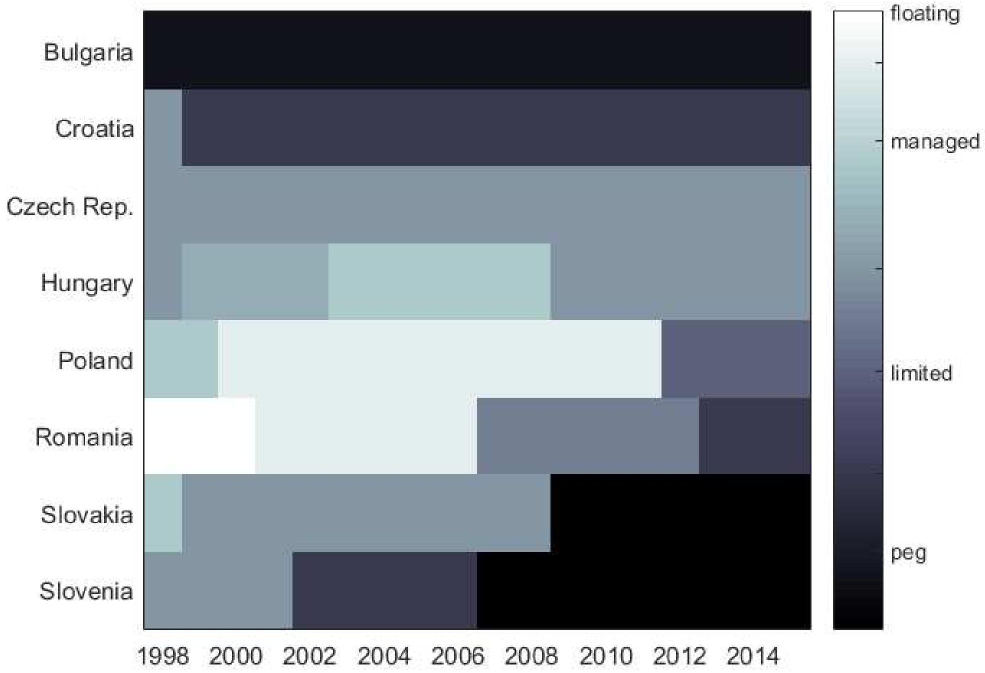
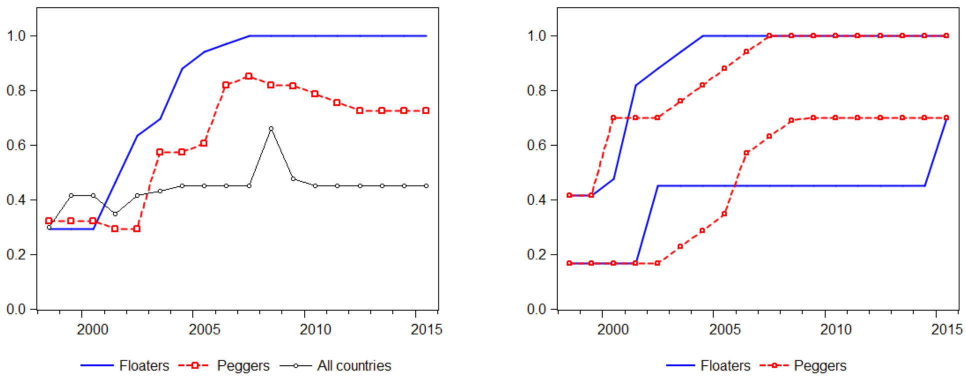


Figure A1: The evolution of exchange rate regimes in CEE countries, 1998-2015
Notes: Exchange rate arrangements: 'peg' stands for a coarse peg category, 'limited' for limited flexibility, 'managed' for managed floating and 'floating' for freely floating. Romania classified as 'freely falling' in 1998-2000.
Source: data from the updated classification of Reinhart and Rogoff (2004).



(a) Median (b) Minimum and maximum

Figure A2: Capital account openness in CEE countries, 1998-2015
Notes: The Chinn-Ito index ranges from 0 to 1.
Source: Data from the dataset developed by Chinn and Ito (2008).

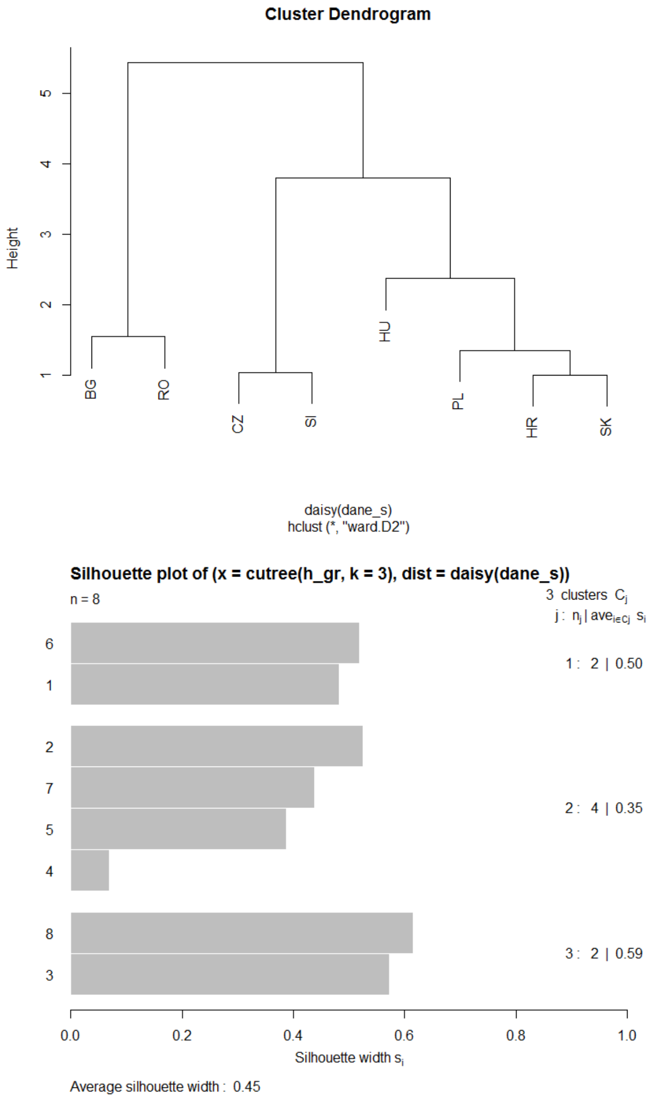


Figure A3: Clustering analysis for CEE countries

Notes: Numbers on the left hand side of the silhouette plot correspond to: 1 – Bulgaria, 2 – Croatia, 3 – the Czech Republic, 4 – Hungary, 5 – Poland, 6 – Romania, 7 – Slovakia, 8 – Slovenia.

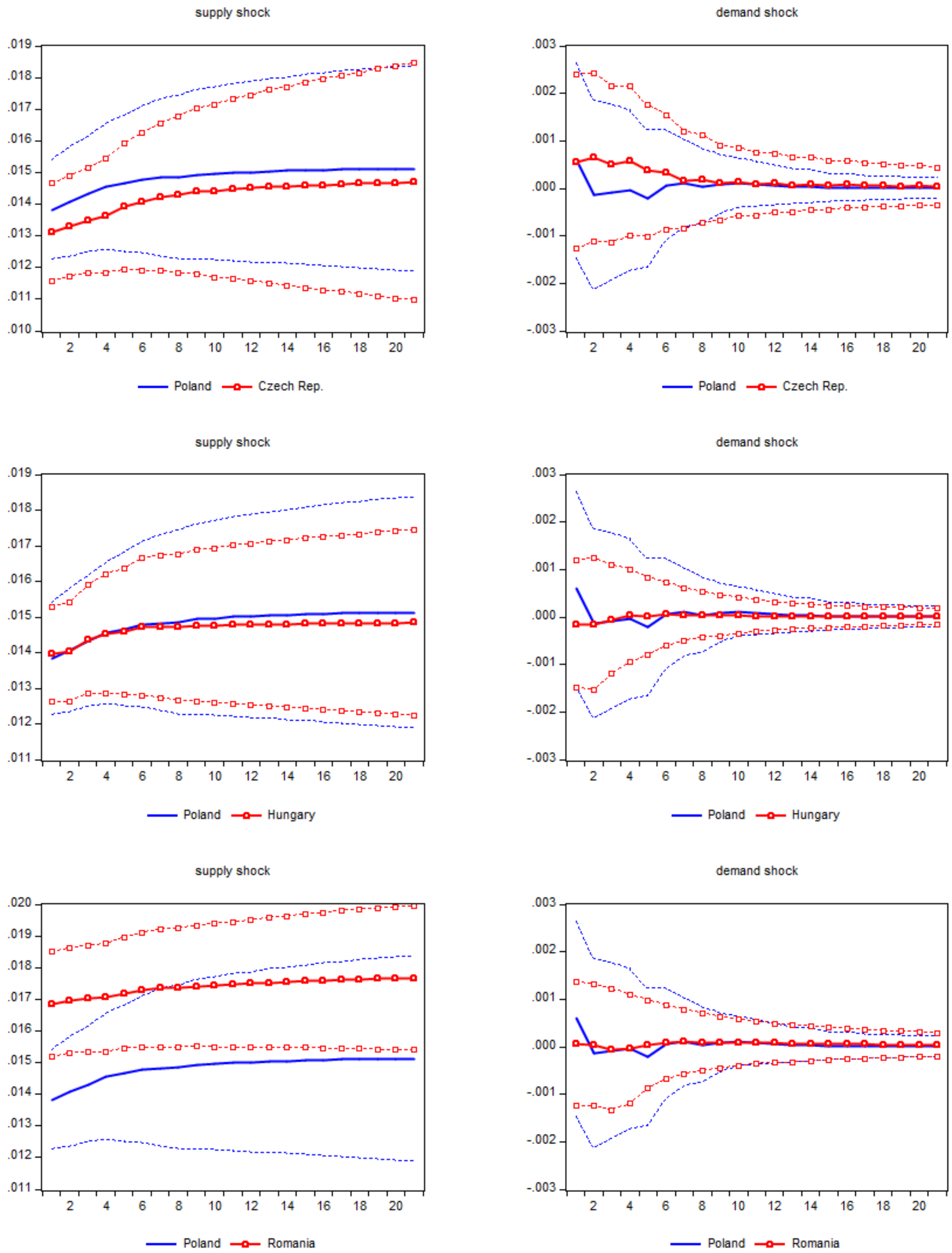


Figure A4: Impulse response functions of the relative output in CEE floaters against Poland

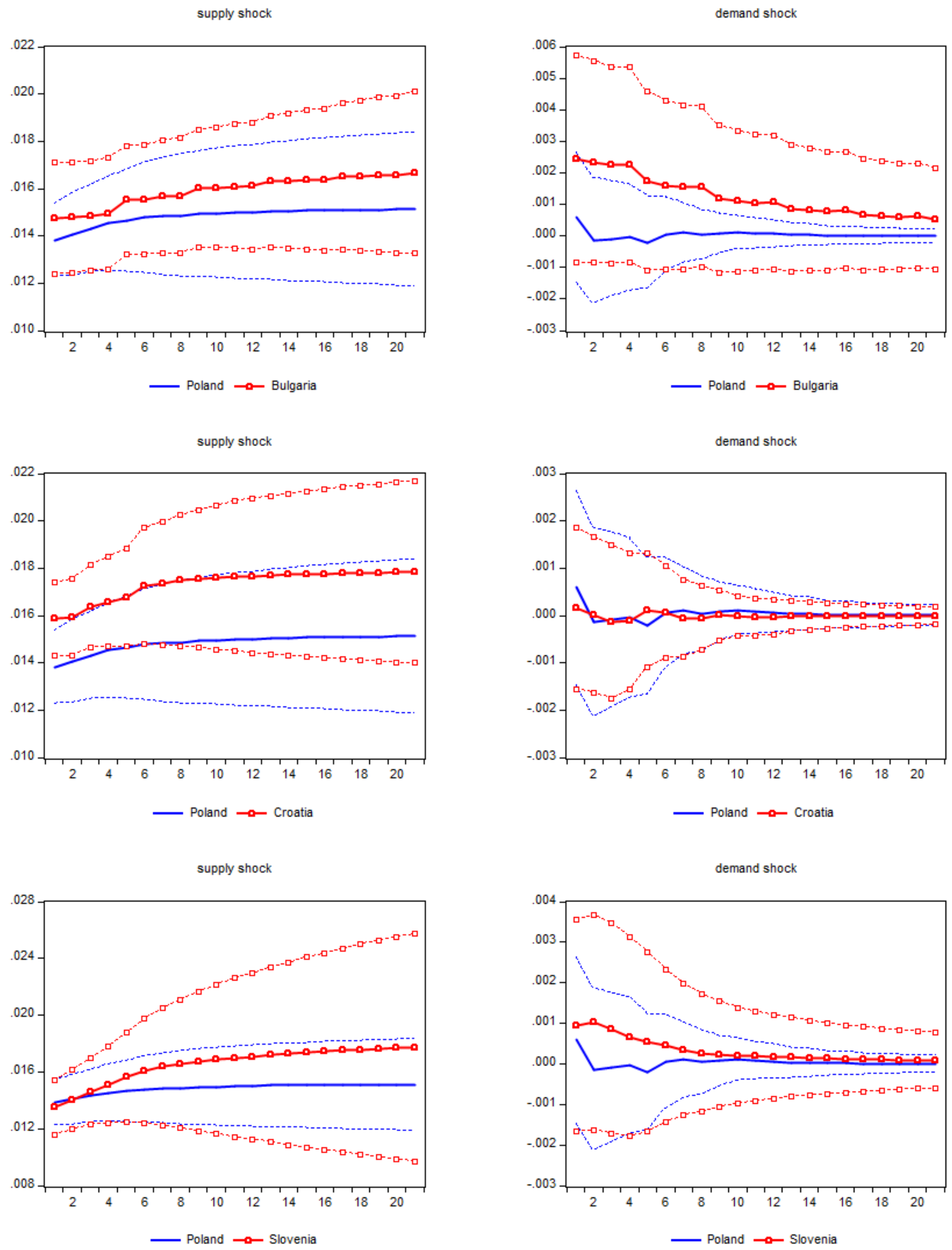


Figure A5: Impulse response functions of the relative output in CEE peggers against Poland

Notes: For Slovakia-Poland pair see Figure 2.