Exchange rate as a shock absorber or a shock propagator in Poland and Slovakia - an approach based on Bayesian SVAR models with common serial correlation

Dąbrowski, Marek A. and Wróblewska, Justyna

Cracow University of Economics, Department of Macroeconomics, Cracow University of Economics, Department of Econometrics and Operational Research

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Marek A. Dąbrowski*† Justyna Wróblewska‡

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Abstract

The paper examines whether exchange rates in Poland and Slovakia acted as shock absorbers or rather shock-propagating mechanisms. A set of Bayesian structural VAR models is built for each country that enables us to identify supply, demand, monetary and financial shocks. Identifying restrictions are derived from the extended stochastic macroeconomic model of an open economy. Sample covers quarterly data 1998-2013. After careful consideration of alternative VAR specifications it is demonstrated that overly parsimonious VARs result in an imperfect identification of shocks that distorts the results. Empirical evidence is found that the higher exchange rate flexibility in Poland than in Slovakia contributed to the absorption of shocks. Though financial shocks had stronger influence on the exchange rate in Poland than in Slovakia, especially in the run-up to the crisis, the participation in the ERM II did not protected the Slovak koruna against the strong and excessive appreciation.

Key Words: open economy macroeconomics; real exchange rate; monetary integration; Bayesian structural VAR; common serial correlation; financial shocks.

JEL Classification: F41; E44; C11

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*Cracow University of Economics, Department of Macroeconomics, Rakowicka 27, 31-510 Kraków, Poland, e-mail: marek.dabrowski@uek.krakow.pl.

†Corresponding author

‡Cracow University of Economics, Department of Econometrics and Operational Research, Rakowicka 27, 31-510 Kraków, Poland, e-mail: eowroble@cyf-kr.edu.pl

1 Introduction

Exchange rate regimes in Poland and Slovakia have evolved along divergent paths: in the late 1990s both countries had intermediate exchange rate regimes (in 1998 it was even the same regime according to Reinhart and Rogoff (2004), i.e. de facto crawling band that is narrower than or equal to ± 5 percent) but then Poland shifted to free floating whereas Slovakia limited fluctuations of the exchange rate and adopted the euro in 2009. Thus, on the one hand, by the time the recent financial crisis has spread on the global economy, these two countries were at opposite poles of exchange rate arrangement spectrum. On the other hand, Poland and Slovakia are relatively similar economies (i.e. at the comparable level of economic and institutional development, similar history[1]). This creates a unique opportunity to examine whether the exchange rate acted as a shock absorber or rather a shock-propagating mechanism and this is the broad objective of our paper. Specifically, we are interested in the relative role of real, financial and monetary shocks in driving the exchange rates in Poland and Slovakia before and during the global financial crisis (GFC). We raise several related questions: What was the nature of shocks that hit Poland and Slovakia before and during the crisis? What were the sources of real exchange rate movements? Were exchange rates acting as a shock absorber or rather a shock-propagating mechanism? Does the ERM II insulate against excessive exchange rate movements?

Empirical evidence on the shock-absorbing property of the flexible exchange rate regime is mixed. Using the SVAR approach Stążka-Gawrysiak (2009) and Dąbrowski and Wróblewska (2013) show that the exchange rate in Poland was driven by real shocks and thus acted as a shock-absorbing mechanism (see also Dąbrowski 2012). In turn Borghijs and Kuijs (2004) find that the nominal shocks had significant contribution in the variation of exchange rates of five Central European countries and conclude that the exchange rate appears on average to have acted as much or more as an unhelpful propagator of nominal shocks than as a useful absorber of real shocks. In a more recent study Shevchuk (2014) confirms these results.

In a related strand of literature on the resilience of emerging market economies (EMEs) to the GFC is examined by cross-country comparisons. Tsangarides (2012) finds that EMEs that peg their currencies weathered the GFC not worse than the those that float but peggers appeared to be faring worse in the recovery period 2010-2011. Similarly Blanchard et al. (2010) find that there is a little direct effect of the exchange rate regime in limiting the decline in the output growth during the crisis. One of policy lessons drawn by Berkmen et al. (2012) after careful examination of factors behind crisis resilience of EMEs is that exchange rate flexibility helped in dampening the impact of the crisis. According to Adler and Tovar (2012) exchange rate flexibility mitigates the impact of adverse financial shocks on EMEs, particularly those that are highly financially integrated. More recently, Bussière et al. (2015)

[1] We share Blanchard's (2010) opinion that 'institutions are central to the workings of a market economy.'

have found that it was the accumulation of foreign reserves prior to the crisis that positively and significantly contributed to the real GDP growth during the crisis and not the exchange rate regime which remained insignificant. The mixed results have been, at least, partly explained by Dąbrowski et al. (2015) who demonstrated that it is not the exchange rate regime *per se* that matters for the crisis resilience, but the specific set of policy tools actually adopted to mitigate the contractionary pressure.

Our contribution to the literature is threefold. First, examining three broad classes of structural VARs (two-, three- and four-variable models) with two types of restrictions (zero and sign restrictions) we provide empirical evidence that the identification scheme used to obtain structural shocks matters.[2] It is demonstrated that the reasonable identification of shocks requires the four-variable VAR. The finding is that the earlier studies that lent support to the hypothesis that the flexible exchange rate regime is chiefly a shock-propagating mechanism (e.g. Shevchuk, 2014) or quite the contrary that it mainly acts as a shock absorber (e.g. Stążka-Gawrysiak, 2009; Dąbrowski, 2012b) rested on the imperfect identification of shocks.

Second, financial shocks are not neglected either in theoretical part or in empirical analysis. Though the importance of these shocks have been pointed out by Stążka-Gawrysiak (2009) our approach is different: restrictions used to identify financial shocks are derived explicitly from the theoretical model. The finding is that financial shocks were an important source of exchange rate variability especially in Poland but the participation in the ERM II did not protected Slovakia against the strong appreciation of the koruna (partly due to financial shocks).

Third, on more technical grounds we employ Bayesian VAR models with common serial correlation restrictions, which enable us to model the common short-run behaviour of the analysed series. Taking account of such commonality leads to more parsimonious models so may increase estimation efficiency. It may also improve the short-run structural analysis. It turned out that the common serial correlation restriction was strongly supported by the analysed data.

The paper is organised as follows. The next section describes monetary background in Poland and Slovakia pointing to differences in the exchange rate regimes adopted. The macroeconomic model is presented in Section 3 and econometric methodology and data are discussed in Section 4. Empirical results are presented in Section 5. The last section concludes.

2 Monetary background

2.1 Monetary policy framework

Poland and Slovakia are quite similar economies in terms of economic and institutional development. They both started their economic transition in the early 1990s and went through deep structural, social and institutional reforms. These efforts were crowned

[2] The point was raised by Artis and Ehrmann (2006) in a study on four advanced economies.

by the accession to the European Union in May 2004. It is interesting to observe that the monetary policy framework in both countries underwent substantial changes at the turn of the millennium: the central banks adopted some form of inflation targeting framework and the exchange rate was officially allowed to float. In Poland changes resulted in an establishment of a fully-fledged direct inflation targeting: in September 1998 the newly-established Monetary Policy Council published the 'Medium-Term Strategy of Monetary Policy' that formally introduced a direct inflation strategy with the medium-term target to reduce inflation 'to below 4% by the year 2003' (NBP, 1998) and the reform was complemented in April 2000 with the de jure shift to a floating exchange rate regime (Table 1).

TABLE 1

The case of Slovakia is less clear-cut. Explicit inflation targeting was introduced by the National Bank of Slovakia (NBS) not until at the beginning of 2005 (NBS, 2004a). Since the koruna was expected to join the ERM II, however, the framework was specified as 'inflation targeting in the conditions of [the] ERM II' (NBS, 2004a). The IMF was more rigorous and – indicating to the fact that Slovakia adopted more than one nominal anchor – classified the Slovak framework as exchange rate anchor and/or inflation targeting (IMF, 2006 and 2007).

It seems, however, that basic elements of inflation targeting were in place since 2001.[3] First, the amendment to the NBS Act of 2001 reformulated the main objective of monetary policy from maintaining stability of currency to maintaining price stability (see the Article 2 of the NBS Act). Second, beginning 2001 the NBS started to publish in its Monetary Programmes the medium-term outlook for inflation development over three succeeding years. Even though the outlook is methodologically different from the genuine target, it had an influence on the inflation expectations of the public. In the words of the NBS (2004a, p. 2) 'the programmed interval of its [inflation] development has represented the implicit monetary policy target.' Third, the NBS took care to communicate its policy to the public through regularly published Monetary Programmes and Monetary Surveys. Nevertheless, the Slovak monetary policy framework was classified by the IMF in the 'Annual Report on Exchange Arrangements and Exchange Restrictions' as 'other' meaning that '[t]he country has no explicitly stated nominal anchor, but rather monitors various indicators in conducting monetary policy' (see IMF, 2001, p. 4). Thus, following the NBS the monetary policy framework before 2005 could be labelled an implicit inflation targeting (NBS, 2014).[4]

A substantial difference between Poland and Slovakia was the significance attached to the exchange rate in monetary policy. The National Bank of Poland sticked by

[3] See, e.g., Mishkin (2003, p. 521-2) and Ortiz Martinez (2008) for a desctirption of inflation targeting framework.

[4] Jonas and Mishkin (2004, p. 380), however, classify monetary policy regime of Slovakia as 'float without an inflation target' and currency regime as 'managed float.'

the free floating regime and in principle restrained from interventions in the foreign exchange market. The typical point in the 'Monetary Policy Guidelines' was that the NBP did 'not rule out foreign exchange interventions should they turn out necessary to ensure domestic macroeconomic and financial stability, which is conducive to meeting the inflation target in the medium term' (NBP 2009, p. 3). Thus, the interventions should be oriented on price stability.

Though there was no predetermined path for the exchange rate of the Slovak koruna, the NBS retained control over the exchange rate: until the ERM II entry in autumn 2005 managed floating regime was employed. The NBS intervened primarily to smooth large fluctuations in the exchange rate and when the exchange rate moved to an unacceptable level. For example, the NBS 'repeatedly intervened [...] in 2003 on the foreign exchange market against excessive appreciation of the Slovak koruna' (NBS, 2004b, p. 22). Moreover, the interest rate policy was, at least partly oriented on exchange rate objective, e.g. in 2002 '[h]eavier pressure on appreciation of the exchange rate of the Slovak koruna in mid-November prompted the NBS Bank Board to reduce the key interest rates by 1.5 percentage points' (NBS, 2003, p. 20).

After the ERM II entry the exchange rate of the koruna was stabilised within the standard fluctuation band of ± 15 per cent around the central parity and the NBS 'paid [attention] to the exchange rate in relation to the development of inflation' (NBS, 2004a, p. 12). In January 2009 the exchange rate was irrevocably fixed and the koruna was replaced with the euro.

2.2 Exchange rate and real economy

A choice of different exchange rate regimes by Polish and Slovak monetary authorities was reflected in exchange rate fluctuations (Table 2). Variability of the nominal exchange rate, measured with the standard deviation of a de-trended component of the quarterly exchange rate over three year intervals, was more than 5 per cent in Poland (except the last sub-period) whereas in Slovakia it was much smaller and remained in an interval of 1-3 per cent. A similar difference could also be observed in the behaviour of real exchange rates, though they were less pronounced. Thus, the relative price level was more volatile in Slovakia than in Poland.

TABLE 2

Both the economic growth and its variability before the global financial crisis (GFC) were greater in Slovakia than in Poland. The crisis changed the ranking with respect to the average growth but not with respect to growth variability. The former declined in both economies but the latter remained almost unchanged in Poland and increased substantially in Slovakia. A related point can be made about fluctuations in the GDP gap: it decreased by less than 1 percentage point in Poland whereas in Slovakia it plummeted by more than 7 percentage points.

Taken together these observations lend some support to the hypothesis that the shocks triggered by the GFC, at least part of them, were absorbed by changes in the nominal exchange rate of the zloty whereas in Slovakia the burden of adjustment was put to a greater extent on changes in the real economy. Greater volatility of the relative price level in Slovakia than in Poland fits this hypothesis as well: due to the rigidity of the koruna's nominal exchange rate larger changes in the relative price level were required in response to real and financial shocks. In other words, the relative price level in Slovakia was under a stronger influence of shocks than its counterpart in Poland where these shocks were rather absorbed by changes in the nominal exchange rate.

It seems, therefore, that the flexible exchange rate in Poland acted as a shock absorber, especially at the time of the crisis, whereas Slovakia had to rely to a greater extent on changes in output. One, however, has to be careful since the observed negative relation between exchange rate variability and GDP growth variability in Poland and Slovakia could be coincidental. Thus, in order to provide more robust evidence on the shock-absorbing property of flexible exchange rate it is useful to uncover the composition of shocks hitting each economy and assess their contribution to the variability of output and exchange rate. At first, however, theoretical model and empirical methodology are presented.

3 Macroeconomic model of an open economy

As a theoretical framework we use the macroeconomic model of an open economy developed by Obstfeld (1985) and Clarida and Galí (1994). This is the stochastic, two-country, rational expectations model of an economy with sluggish price adjustment to supply, demand and monetary shocks. The model, however, is extended to include financial shocks which are defined as unexpected changes in the risk premium. Intuitively, the modification seems to be important because financial shocks are a quite plausible source of exchange rate fluctuations, especially in the GFC, and should not be neglected. Moreover, as demonstrated by Dąbrowski (2012a) if financial shocks are not considered a separate type of shocks, they enter both demand and monetary shocks and distort empirical results (unless risk premium is constant). The second extension of the model is that all shocks (and not only demand shock) are allowed to have permanent and transitory components.

The model consists of four building blocks which are the IS and LM relations, the uncovered interest rate parity (UIP) condition and the price-setting (PS) relation:

$$\begin{aligned} y_t^d &= d_t - \eta\left(s_t + p_t\right) - \sigma\left[i_t - E_t\left(p_{t+1} - p_t\right)\right] && (1)\\ m_t^s &= p_t + y_t - \lambda i_t && (2)\\ i_t &= -E_t\left(s_{t+1} - s_t\right) + x_t && (3)\\ p_t &= \left(1-\theta\right) E_{t-1} p_t^e + \theta p_t^e && (4) \end{aligned}$$

All variables except interest rates are in logs and represent a difference between home (H) and foreign (F) levels, e.g. the relative price level, p_t, is defined as $p_t^H - p_t^F$. Equation (1), the IS relation, says that the relative demand for home goods, y_t^d, depends on the relative demand disturbance (d_t), the real exchange rate ($q_t = s_t + p_t$) and the real interest rate differential (expression in square brackets). Equation (2) is an equilibrium condition for money market (a conventional LM relation). The UIP condition (3) requires the nominal interest rate differential, i_t, to be equal to the expected depreciation of domestic currency adjusted for the risk premium, x_t. Equation (4) is the PS relation: since prices are sticky, the relative price level, p_t, is an average of its equilibrium level expected in time $t-1$ to prevail in period t and the price level that would actually clear the market under fully flexible prices in time t, p_t^e.

There are four stochastic processes that govern the relative supply of output, y_t^s, the relative demand disturbance, d_t, relative money, m_t, and the risk premium, x_t:

$$
\begin{aligned}
y_t^s &= y_{t-1}^s + u_t^s - \gamma_1 u_{t-1}^s && (5)\\
d_t &= d_{t-1} + u_t^d - \gamma_2 u_{t-1}^d && (6)\\
x_t &= x_{t-1} + u_t^f - \gamma_3 u_{t-1}^f && (7)\\
m_t &= m_{t-1} + u_t^m - \gamma_4 u_{t-1}^m && (8)
\end{aligned}
$$

Thus, shocks to supply, demand, risk premium and money are allowed to have both permanent and transitory components: a fraction γ of any shock in period t is expected to be reversed in $t+1$.

One can demonstrate that under the flexible exchange rate regime the long-run flexible price rational expectations equilibrium levels of relative output, real interest rate differential, real exchange rate and relative price level are:[5]

$$
\begin{aligned}
y_t^e &= y_t^s && (9)\\
r_t^e &= x_t + \frac{1}{\eta+\sigma}(-\gamma_1 u_t^s + \gamma_2 u_t^d - \sigma\gamma_3 u_t^f) && (10)\\
q_t^e &= \frac{d_t - y_t^s - \sigma x_t}{\eta} - \frac{\sigma}{\eta(\eta+\sigma)}(-\gamma_1 u_t^s + \gamma_2 u_t^d - \sigma\gamma_3 u_t^f) && (11)\\
p_t^e &= m_t - y_t^s + \lambda x_t + \alpha_1\gamma_1 u_t^s + \alpha_2\gamma_2 u_t^d - \alpha_3\gamma_3 u_t^f - \alpha_4\gamma_4 u_t^m && (12)
\end{aligned}
$$

where α's are functions of parameters and are positive.

If monetary authorities credibly commit to a fixed exchange rate, say $\bar{s}$, then the relative price level in the long-run equilibrium is simply

$$p_t^e = p_t^e - \bar{s} \qquad (13)$$

and the equilibrium levels for y_t^e, r_t^e and q_t^e are the same as under the flexible exchange rate.[6]

[5] For details see the appendix.

[6] For details see Dąbrowski (2012a).

The long-run identifying restrictions can now be derived from the solution of the model and are presented in Table 3. As is clear, zero restrictions are not enough to identify four shocks: the issue is that demand and financial shocks cannot be separated one from another. Fortunately, sign restrictions can be used to solve the problem.[7]

TABLE 3

The table can also be used to describe restrictions placed in a three-variable VAR, i.e. the one used by Clarida and Galí (1994). It is enough to omit a row with the real interest rate differential and a column with financial shock. In such a case shocks can be identified with zero restrictions only.

4 Empirical strategy and data

4.1 Empirical methodology

Two groups of structural VAR models are used: with zero restrictions only and with zero and sign restrictions. In the former case the methodology originally proposed by Blanchard and Quah (1989) is employed to identify structural shocks. The recursive structure of the total impact matrix is obtained by the use of the Cholesky decomposition with the normalizing assumptions based on the diagonal signs displayed in Table 3. In the latter case, i.e. for the group of SVAR models with both zero and sign restrictions, the algorithm proposed by Arias, Rubio-Ramírez and Waggoner (2014) is used to identify structural shocks.

In all cases the analysis starts with an $n-$dimensional stable Gaussian VAR(k) process:

$$y_t = A_1 y_{t-1} + A_2 y_{t-2} + \cdots + A_k y_{t-k} + \Phi D_t + \varepsilon_t,\ \{\varepsilon_t\} \sim iiN(0,\Sigma),\ t = 1, 2, \ldots, T, \quad (14)$$

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 $y_{-k+1}, y_{-k+2}, \ldots, y_0$ 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 (14) reads as follows:

$$Y = X\Gamma + Z\Gamma_d + E, \quad (15)$$

where $\Gamma_{nk\times n} = \begin{pmatrix} A_1 & A_2 & \ldots & A_k \end{pmatrix}'$, $\Gamma_d = \Phi'$, $Y_{T\times n} = \begin{pmatrix} y_1 & y_2 & \ldots & y_T \end{pmatrix}'$, $x_t = \begin{pmatrix} y'_{t-1} & y'_{t-2} & \ldots & y'_{t-k} \end{pmatrix}'$, $X_{T\times nk} = \begin{pmatrix} x_1 & x_2 & \ldots & x_T \end{pmatrix}'$, $E_{T\times n} =$

[7]The reaction of the relative price to financial shock depends on the exchange rate regime: with a floating rate the reaction is positive and under a fixed rate it is negative. In the empirical part this sign restriction is not imposed.

$\begin{pmatrix} \varepsilon_1 & \varepsilon_2 & \dots & \varepsilon_T \end{pmatrix}'$, $Z_{T\times l} = \begin{pmatrix} D_1 & D_2 & \dots & D_T \end{pmatrix}'$ and l denotes the number of deterministic components.

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 commonly known matrix Normal-inverted Wishart structure:

1. $\Sigma_{n\times n} \sim iW(S, q_\Sigma)$, where S is a PDS matrix and $q_\Sigma \geq n$,

2. $\Gamma_{nk\times n}|\Sigma, \nu_\Gamma \sim mN(\underline{\mu}_\Gamma, \Sigma, \nu_\Gamma \underline{\Omega})$, where $\underline{\Omega}$ is a PDS matrix of order nk and the additional parameter ν_Γ can be estimated or determined by the researcher,

3. $\nu_\Gamma \sim iG(s_\Gamma, n_\Gamma)$, if the ν_Γ is to be estimated; $iG(s_\Gamma, n_\Gamma)$ denotes an inverted Gamma distribution with parameters s_Γ and n_Γ, i.e. $p(\nu_\Gamma) \propto \nu_\Gamma^{-n_\Gamma - 1}\exp(-\frac{s_\Gamma}{\nu_\Gamma})$,

4. $\Gamma_d|\Sigma, \nu_d \sim mN(\underline{\mu}_d, \Sigma, \nu_d I_l)$, where I_l denotes an identity matrix of order l and ν_d is an additional parameter which controls the tightness of the above-stated matrix normal distribution,

5. $\nu_d \sim iG(s_d, n_d)$, if the researcher wants ν_d to be estimated.

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. $\Sigma|., Y \sim iW(S + E'E + \frac{1}{\nu_\Gamma}(\Gamma - \underline{\mu}_\Gamma)'\underline{\Omega}^{-1}(\Gamma - \underline{\mu}_\Gamma) + \frac{1}{\nu_d}(\Gamma_d - \underline{\mu}_d)'(\Gamma_d - \underline{\mu}_d), q_\Sigma + nk + l + T)$,

2. $\Gamma|., Y \sim mN(\overline{\mu}_\Gamma, \Sigma, \overline{\Omega})$, where $\overline{\Omega} = (\frac{1}{\nu_\Gamma}\underline{\Omega}^{-1} + X'X)^{-1}$, $\overline{\mu}_\Gamma = \overline{\Omega}(\frac{1}{\nu_\Gamma}\underline{\Omega}^{-1}\underline{\mu}_\Gamma + X'(Y - Z\Gamma_d))$,

3. $\Gamma_d|., Y \sim mN(\overline{\mu}_d, \Sigma, \overline{\Omega}_d)$, where $\overline{\Omega}_d = (\frac{1}{\nu_d}I_l + Z'Z)^{-1}$, $\overline{\mu}_d = \overline{\Omega}_d(\frac{1}{\nu_d}\underline{\mu}_d + Z'(Y - X\Gamma))$,

   and, in the case of ν_Γ and ν_d being estimated:

4. $\nu_\Gamma|., Y \sim iG(s_\Gamma + \frac{1}{2}tr(\Sigma^{-1}(\Gamma - \underline{\mu}_\Gamma)'\underline{\Omega}^{-1}(\Gamma - \underline{\mu}_\Gamma)), n_\Gamma + \frac{n^2 k}{2})$,

5. $\nu_d|., Y \sim iG(s_d + \frac{1}{2}tr(\Sigma^{-1}(\Gamma_d - \underline{\mu}_d)'(\Gamma_d - \underline{\mu}_d)), n_d + \frac{nl}{2})$.

The usefulness of VARs in the analysis of multivariate time series is non-questionable. It is known that analysis of $CI(1,1)$ cointegrated series leads to reduced rank restrictions for the VARs' parameters. Cointegration is notion that stochastically trending series move together in the long-run, however one can ask whether stationary (possibly de-trended) series can move together in the short-run. It is of course possible and including such possibility into the model may improve e.g. structural analysis and forecast, especially in the short-run horizon. Such behaviour can also be modeled in the framework of VAR. Kozicki and Engle (1993) termed it common serial correlation. Both cointegration and common serial correlation are examples of a co-feature idea (for the introduction see e.g. Kozicki, Engle 1993 and Vahid, Engle 1993, for the survey see e.g. Centoni, Cubadda 2011).

As shown by Engle and Kozicki (1993), modeling common serial correlation in the framework of VAR leads to reduced rank of the matrix parameters of the process (14). If this co-correlation is contemporaneous the rank reduction restriction applies to the matrices A_1 through A_k, otherwise to the matrices A_i through A_k for some $i \geq 2$ and $i \leq k$.

In this research the possibility of only contemporaneous common serial correlation among the analysed series is tested, so the model of the following form is considered:

$$y_t = \gamma\delta_1 y_{t-1} + \gamma\delta_2 y_{t-2} + \cdots + \gamma\delta_k y_{t-k} + \Phi D_t + \varepsilon_t,\ \{\varepsilon_t\} \sim iiN(0, \Sigma),\ t = 1, 2, \ldots, T, \tag{16}$$

where meaning of symbols Σ, $\{\varepsilon_t\}_{t=1}^{T}$, D_t, Φ is left the same as in the process (14). The matrix form of the process (16) reads as follows:

$$Y = X\gamma\delta' + Z\Gamma_d + E, \tag{17}$$

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 (15)). 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 (15) is of reduced rank equal to $n-s$ ($\Gamma = \gamma\delta'$).

Imposing the rank restrictions leads to more parsimonious model, so much on the estimation efficiency can be gained, but the researcher have to deal with the problem of the parameters non-identification, which is similar to that one known from the error correction modeling, so we the solutions known from VEC models can be adopted.

Specifically, it is obvious that for any non-singular matrix M of order $n-s$ products $\gamma\delta'$ and $\gamma MM^{-1}\delta'$ 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, León-González and Strachan (2010) for the VEC models, also employed for the parameters of models with cointegration and so called weak form (polynomial) common cyclical features by Wróblewska (2011, 2012).

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

$$\gamma\delta' = \gamma O_{\Gamma}^{-1} O_{\Gamma}\delta' \equiv GD', \tag{18}$$

where O_Γ is an $n-s \times n-s$ symmetric positive definite matrix. In the left-hand of (18) 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 $D \in \mathbb{R}^{nk(n-s)}$. For G and D we settle matrix normal priors of the following form:

1. $D|\nu_G \sim mN(0, cI_{n-s}, P_D)$, where c is a positive constant; through the matrix P_D the researcher may incorporate prior information (for the details see Koop et al. 2010 and Strachan, Inder 2004),

2. $G|\nu_G \sim mN(0, \nu_G I_{n-s}, \Sigma)$.

The above-settled prior for D leads to a matrix angular central Gaussian distribution for $D(D'D)^{-\frac{1}{2}}$, which is a matrix with orthonormal columns $(D(D'D)^{-\frac{1}{2}} \sim MACG(P_D)$, for the details see Chikuse 2002). By assuming $\delta = D(D'D)^{-\frac{1}{2}}$ and $\gamma = G(D'D)^{\frac{1}{2}}$ we fulfill one of our assumptions. Additionally, according to our second assumption, $theMACG(P_D)$ prior is truncated, by taking into account only matrices with positive elements in the first row. For the remaining parameters we impose the same priors as in the VAR model and. 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. $\Sigma|., Y \sim iW(S + E'E + \frac{1}{\nu_G}GG' + \frac{1}{\nu_d}(\Gamma_d - \underline{\mu}_d)'(\Gamma_d - \underline{\mu}_d), q_\Sigma + n - s + l + T)$,

2. $G|., Y \sim mN(\overline{\mu}_G, \overline{\Omega}_G, \Sigma)$, 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 $\overline{\Omega}_{vD} = (((G'\Sigma^{-1}G) \otimes (X'X)) + (\frac{1}{c}I_{n-s} \otimes P_D^{-1}))^{-1}$, $\overline{\mu}_{vD} = \overline{\Omega}_{vD}vec(X'(Y - Z\Gamma_d)\Sigma^{-1}G)$ and $vec(M)$ denotes the vectorisation of a matrix M,

4. $\Gamma_d|., Y \sim mN(\overline{\mu}_d, \Sigma, \overline{\Omega}_d)$, where $\overline{\Omega}_d = (\frac{1}{\nu_d}I_l + Z'Z)^{-1}$, $\overline{\mu}_d = \overline{\Omega}_d(\frac{1}{\nu_d}\underline{\mu}_d + Z'(Y - XDG'))$,

   and, in the case of ν_G and ν_d being estimated:

5. $\nu_G|., Y \sim iG(s_G + \frac{1}{2}tr(\Sigma^{-1}GG'), n_G + \frac{n(n-s)}{2})$,

6. $\nu_d|., Y \sim iG(s_d + \frac{1}{2}tr(\Sigma^{-1}(\Gamma_d - \underline{\mu}_d)'(\Gamma_d - \underline{\mu}_d)), n_d + \frac{nl}{2})$.

Samples from the posterior distributions of δ and γ can be obtained by using transformations: $\delta = D(D'D)^{-\frac{1}{2}}$ and $\gamma = G(D'D)^{\frac{1}{2}}$.

As was previously mentioned, two groups of SVAR models are considered. In the first group only zero restrictions are imposed and Blanchard-Quah method is employed. This method is commonly known, so we are not going to describe it here in details (the interested readers are referred to e.g. Blanchard, Quah 1989, Clarida, Galí 1994, Lütkepohl 2006 pp. 367-368).

The identifying method applied in the group combining zero and sign restrictions is probably less known, so we outline it in the Technical Appendix.

4.2 Prior specifications used in the empirical analysis

In each considered group we impose priors of the following form:

1. for the specific parameters of VARs without common serial correlation:

$$Y = X\Gamma + Z\Gamma_d + E, \quad E \sim mN(0, \Sigma, I_T)$$

   (a) $\Gamma|\Sigma, \nu_\Gamma \sim mN(0, \Sigma, \frac{\nu_\Gamma}{nk} I_{nk})$,

   (b) $\nu_\Gamma \sim iG(200, 3)$, therefore $E(\nu_\Gamma) = 100$, $D(\nu_\Gamma) = 100$,

2. for the specific parameters of models with common serial correlation:

$$Y = XDG' + Z\Gamma_d + E, \quad E \sim mN(0, \Sigma, I_T)$$

   (a) $D \sim mN(0, \frac{1}{nk} I_q, I_{nk})$,

   (b) $G|\Sigma, \nu_G \sim mN(0, \nu_G I_q, \Sigma)$,

   (c) $\nu_G \sim iG(200, 3)$, therefore $E(\nu_G) = 100$, $D(\nu_G) = 100$,

3. for the common parameters:

   (a) $\Sigma \sim iW(0.01 I_n, n+1)$, where n denotes the number of variables and I_n - an identity matrix of order n,

   (b) $\Gamma_d|\Sigma, \nu_d \sim mN(0, \Sigma, \nu_d I_l)$, where l stands for the number of non-stochastic variables,

   (c) $\nu_d \sim iG(200, 3)$, therefore $E(\nu_d) = 100$, $D(\nu_d) = 100$.

4.3 Data

All data are from the Eurostat database. We use quarterly data spanning from 1998:1 to 2013:4. Real GDP is used as a measure of output. Real interest rate is calculated as a difference between 3-month money market nominal interest rate and actual inflation. Real exchange rate is based on 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. Price level is measured with a harmonised index of consumer prices (the same measure is used for inflation). Relative output and relative price level are constructed as the log-differences between domestic and foreign (euro area) variables. Real interest rate differential is the difference between domestic and foreign rates.

5 Empirical results

In all cases the empirical analysis starts with the Bayesian model comparison and we assume equal prior probability of each specification. There are seasonal dummies and a constant in each model. Additionally, we have decided to include 0-1 dummy in models for Slovakia to account for the participation in the ERM II and the euro area.

The Savage-Dickey density ratio (SDDR, see e.g. Verdinelli and Wasserman 1995) has been employed to compute posterior probability of each model specification (Bayes factors comparing each model to a model with only non-stochastic variables on the right side have been calculated).

The results are obtained with the advantage of the Bayesian model averaging technique employed in the set of models with the highest posterior probability.

Visual inspection of the CUMSUM plots (not presented in the paper) implies that the number of burnt-in cycles needed to attain the convergence equals 500,000. The results reported are based on 100,000 accepted Gibbs draws.

5.1 The case against the exchange rate as a shock absorber

Following Canzoneri et al. (1996) we start empirical analysis with 'the most parsimonious VAR possible' that includes only two variables, the relative output and the nominal exchange rate (both are in the log differences for stationarity). The approach allows us to identify two structural shocks, i.e. non-neutral (or permanent), u_t^p, and neutral, u_t^n, depending on whether they have a long-run impact on the relative output. The model is

$$\begin{bmatrix} \Delta y_t \\ \Delta s_t \end{bmatrix} = \begin{bmatrix} C_{11}(L) & C_{12}(L) \\ C_{21}(L) & C_{22}(L) \end{bmatrix} \begin{bmatrix} u_t^p \\ u_t^n \end{bmatrix} \tag{19}$$

where $C(L)$'s are lag polynomials and the identifying restriction is that $C_{12}(1) = 0$, i.e. neutral shocks have no long-run impact on output. The normalizing assumptions are that $C_{11}(1)$ and $C_{22}(1)$ are positive.

The set of compared models consists of 10 non-nested specifications which may differ in the lag order ($k \in \{5, 6, \ldots, 9\}$). For each settled k we consider two cases: without and with one co-feature vector. Among the most probably models there are only those with common serial correlation (see Table 4).

TABLE 4

The results of the forecast error variance (FEV) decomposition for both variables (in levels) are depicted in Table 5. In both countries relative output remained under an overwhelming influence of non-neutral shocks: they accounted for more than 90 per cent of the posterior median of FEV at all horizons. Even if one takes into account uncertainty related with this estimates, which is measured with a difference between the 84th and 16th quantiles of the posterior distribution, the predominance of permanent shocks is unquestionable.[8] The opposite is true for the variability of the nominal exchange rate – they are driven by neutral shocks although non-neutral shocks account for 6-7 per cent of median FEV at all forecast horizons.

TABLE 5

The results are in line with those obtained in other studies that used a two-variable VARs. Borghijs and Kuijs (2004) analysed five CEE countries (the Czech Rep., Hungary, Poland, Slovakia and Slovenia) and using data for 1995-2003 period found that 'the nominal exchange rate does not respond to the shocks that seem to cause the bulk of fluctuations in output' and that 'the results cast doubt on the usefulness of the exchange rate as shock absorber.' For more recent data, 1999-2013, Shevchuk (2014) obtained similar results.

The problem, however, is that the way shocks are defined seems to be rather questionable: demand, financial and monetary shocks are in fact put together into the neutral shock. It neglects the warning put forward by Enders (2004, p. 310) that it is wise to avoid a decomposition with only two types of disturbances when the presence of three or more important disturbances is suspected. Both Borghijs and Kuijs (2004) and Canzoneri et al. (1996) admit that the interpretation of the neutral shock is ambiguous and '[r]eluctantly... conclude that the parsimonious 2-variable VARs will not suffice' (Canzoneri et al. 1996, p. 17).

5.2 The case in favour of the exchange rate as a shock absorber

In order to disentangle real demand shocks from monetary shocks the three-variable VAR can be used. Following Clarida and Galí (1994) we built the model with the relative output, the real exchange rate and the relative price level (all are in the log differences for stationarity). Three structural shocks can be identified: supply, u_t^s, demand, u_t^d, and monetary, u_t^m. The model is

$$\begin{bmatrix} \Delta y_t \\ \Delta q_t \\ \Delta p_t \end{bmatrix} = \begin{bmatrix} C_{11}(L) & C_{12}(L) & C_{13}(L) \\ C_{21}(L) & C_{22}(L) & C_{23}(L) \\ C_{31}(L) & C_{32}(L) & C_{33}(L) \end{bmatrix} \begin{bmatrix} u_t^s \\ u_t^d \\ u_t^m \end{bmatrix} \tag{20}$$

where the identifying restrictions are that $C_{12}(1) = C_{13}(1) = C_{23}(1) = 0$, i.e. demand shocks do not affect output in the long-run and monetary shocks are neutral both for output and real exchange rate. The normalizing assumptions based on the diagonal signs are that $C_{11}(1), C_{22}(1), C_{33}(1)$ are positive (see Table 3).

The set of models compared consist of 15 different specifications, i.e. VAR with 5 throughout 9 lags ($k \in \{5, 6, \ldots, 9\}$) without common serial correlation or with

[8]The quantile difference used in the text corresponds to 0.68 probability mass of posterior distribution.

one or two co-features ($s \in \{0, 1, 2\}$). Models with posterior probabilities higher than the assumed prior probability are displayed in Table 6. Interestingly, among the models listed in Table 6 are only those with reduced rank structures, so the data again strongly support the hypothesis of the existence of common correlation among the analysed variables. The results indicate that the series share two features.

TABLE 6

Table 7 displays the results of the (median) variance decompositions at one-year forecast horizon for both economies.[9] Uncertainty related with these estimates is measured with the difference between the 84th and 16th quantile of the posterior distribution. It is reported in parentheses.

TABLE 7

Several interesting observations can be made. First, output variability has been predominantly driven by supply shocks in both economies: more than 90 per cent of variation in the relative output is accounted for by supply shocks. The other shocks prove to be unimportant, especially if one takes into account uncertainty related to the estimates. Second, the real exchange rates have remained under the overwhelming influence of demand shocks: these account for more than 90 per cent of variation in the real exchange rates whereas the (median) contribution of monetary shocks has been negligible. Interestingly the contribution of supply shocks to real exchange rate variability is twice as large in Poland (8%) as in Slovakia (3-4%). Finally, the real shocks account for more than 12 per cent of the FEV of the relative price level in Slovakia and only less than 3 per cent in Poland.

These findings cast even more doubts on two-variable VARs: it turns out that one should not interpret the neutral shock as a monetary (nominal) shock. It is rather composed of real shocks to aggregate demand. The argument, of course, could be that we focus on *real* and not on *nominal* exchange rate. Thus, in order to make the comparison clear-cut, the VAR with the nominal exchange rate instead of the relative price level has been estimated.[10] The forecast error variance decompositions for the relative output and the real exchange rate are almost unchanged. It turned out that the variability in the nominal exchange rate was driven by demand shocks in Poland with their median contribution to the FEV (at four quarter forecast horizon) equal 74.9 per cent and by monetary shocks in Slovakia with median contribution at 64.3 per cent.

Our results for Poland are consistent with findings of Stążka-Gawrysiak (2009) and Dąbrowski (2012b) who demonstrated that real demand shocks were the most important in explaining variation of the zloty's exchange rate. The results for the

[9] Results for other forecast horizons are quite similar and are available from the authors upon request.

[10] The results are not reported in tables though they are available from the authors upon request.

exchange rate of the Slovak koruna are in line with those of Canzoneri et al. (1996) who uncovered that supply and demand shocks accounted for 19-33 per cent of the variation in exchange rates (after one year) of five European Union Member States. This similarity is reasonable, because Canzoneri et al. (1996) focused on currencies whose exchange rates – like the one of the Slovak koruna – were stabilized within the European Monetary System (and its predecessor, the European currency snake).

Overall, our interpretation of the results obtained is that the higher exchange rate flexibility in Poland than in Slovakia can hardly be seen as a factor contributing to the propagation of monetary shocks. It seems that the opposite is true: the exchange rate of Polish zloty indeed acted as a shock absorber. The shock absorbing property of the exchange rate has been dampened in Slovakia by the policy oriented on exchange rate stability and the burden of the adjustment to shocks has been shifted on the relative price level.

Though the three-variable VARs give more insight into the shocks behind the exchange rate fluctuations, they do not distinguish between monetary and financial shocks. This is a shortcoming of the approach because it introduces an inconsistency between theoretical model and empirical analysis. If one allows for the time-varying risk premium driven by financial shocks in the uncovered interest rate parity then it can be demonstrated that the demand and monetary shocks are no longer uncorrelated and both contain an additive of financial shocks (see Dąbrowski, 2012, APE). This in turn undermines the assumption conventionally adopted in the empirical analysis that structural shocks are uncorrelated. Thus, the link between theory and empirics is broken: what is estimated cannot be legitimately called demand and monetary shocks. Fortunately, the problem can be solved by introducing a financial shock and application of four-variable VAR model.

5.3 A more balanced view on the exchange rate as a shock absorber

In order to treat a financial shock as a separate shock, the VAR has to be extended. Thus, we built the model with the relative output, the real interest rate differential, the real exchange rate and the relative price level (all are in the log differences for stationarity except interest rates which are in the differences). This enables us to identify four structural shocks: supply, u_t^s, demand, u_t^d, financial, u_t^f and monetary, u_t^m. The model is

$$\begin{bmatrix} \Delta y_t \\ \Delta r_t \\ \Delta q_t \\ \Delta p_t \end{bmatrix} = \begin{bmatrix} C_{11}(L) & C_{12}(L) & C_{13}(L) & C_{14}(L) \\ C_{21}(L) & C_{22}(L) & C_{23}(L) & C_{24}(L) \\ C_{31}(L) & C_{32}(L) & C_{33}(L) & C_{34}(L) \\ C_{41}(L) & C_{42}(L) & C_{43}(L) & C_{44}(L) \end{bmatrix} \begin{bmatrix} u_t^s \\ u_t^d \\ u_t^f \\ u_t^m \end{bmatrix} \tag{21}$$

where the identifying restrictions are that $C_{12}(1) = C_{13}(1) = C_{14}(1) = C_{24}(1) = C_{34}(1) = 0$, i.e. demand and financial shocks do not affect output in the long-run and monetary shocks are neutral for output, real interest rate differential and real exchange rate. These restrictions are not sufficient to identify four shocks, particularly demand and financial shocks could not be distinguished one from another. Therefore, not only normalizing assumptions but also sign restrictions are imposed: $C_{11}(1), C_{22}(1), C_{23}(1), C_{32}(1), C_{42}(1), C_{44}(1)$ are positive and $C_{33}(1)$ is negative (see Table 3).

An explicit treatment of financial shocks makes the four-variable VAR the most reliable model. On the one hand, it distinguishes between demand and monetary shocks, so the risk of underestimation of real shocks' contribution into exchange rate variability is substantially reduced as proved by the comparison of two- and three-variable VARs. On the other hand, the risk of overestimating the shock-absorbing property of the flexible exchange rate is rather small since financial shocks are separated from demand shocks. Thus, one cannot legitimately argue that the relatively high contribution of demand shocks to exchange rate variability – the common result from the three-variable VAR – is due to the fact that demand and financial shocks are tangled. Moreover, one can claim that with the global financial crisis in the sample financial shocks 'deserve' to be treated separately.

The set of 20 non-nested VAR specifications for each country has been considered ($k \in \{5, 6, \ldots, 9\}$ and $s \in \{0, 1, 2, 3\}$). Models with posterior probabilities higher than assumed prior probability are displayed in Table 8. Similarly to the three-variable case, in the wider system models allowing for common serial correlation gathered almost all posterior probability. The hypothesis of the existence of three co-features among the analysed series is strongly supported.

TABLE 8

Accounting for the common serial correlation leads to more diffused posterior probability of the VAR order. In our previous study based on the same set of data, but without common serial correlation feature, VAR of order 5 gathered almost all model posterior probabilities, so the results therein were based on this model only (see Wróblewska and Dąbrowski, 2014).

Since the four-variable VAR is more reliable than the previous two models, we present the detailed results obtained from it below. Before we proceed to the FEV decomposition, the residuals from the VAR are examined. These are interesting because they measure the magnitude and the (on-impact) response of an economy to a bundle of structural shocks. Thus, for example, the higher the variance of residuals from the equation for the relative output, the more vulnerable to asymmetric shocks an economy is. From Table 9, in which variances of residuals are presented (with ∓ 2 standard deviations), one can read that both economies have been subjected to asymmetric shocks to a comparable extent. The sharp difference between Poland and Slovakia is in the behaviour of real exchange rate: with more flexibility the Polish zloty was more responsive to shocks than the Slovak koruna. Since the relative price level vulnerability to shocks was similar across countries the stronger responsiveness of real exchange rate in Poland than in Slovakia was due to the behaviour of nominal exchange rate.

TABLE 9

These results are hardly a surprise: they match intuition well. The interesting question is whether the real (and nominal) exchange rate was under similar shocks as the relative output. If variables are under the influence of different shocks, then the correlation between residuals should be close to zero. In other words, variables are disconnected one from another. If, in turn, they are driven by similar shocks the correlation should be different from zero. Shocks behind changes in the output and real exchange rate in Poland were quite similar: the correlation between respective residuals is strongly negative (-0.29), whereas in Slovakia it is close to zero (-0.03). It should be emphasized that it is not the relative price level that is behind this correlation: in Poland the real exchange rate and the relative price level are only weakly correlated (0.12) and in Slovakia the opposite is true (0.35).

Interestingly, the real interest rate differential was more vulnerable to shocks in Slovakia than in Poland. It seems, therefore, that the greater stability of the Slovak koruna was paid for with higher variability of real interest rate differential. This observation is confirmed by similarity of shocks behind the real interest rate differential and the relative price level (and real exchange rate): the correlation between respective residuals was strongly negative (-0.36).

The importance of shocks is illustrated with the forecast error variance decompositions which are reported in Table 7. The FEV of the relative output is dominated by supply shocks that account for more than 90 per cent of variance in both countries. The real exchange rates are still under a strong influence of demand shocks, although their contribution in the FEV is not as large as in the three-variable VAR: 47.4% vs. 91.2% in Poland and 58.1% vs. 96.0% in Slovakia. The inclusion of financial shocks into the analysis seems to be a right move: they account for more than one-third of FEV of real exchange rate. The difference between Poland and Slovakia in this respect does not seems to be very large, especially if one takes into account uncertainty associated with estimates. Moreover, it seems that a slightly higher resilience of real exchange rate to financial shocks in Slovakia was possible because the real interest rate differential was exposed to these shocks to a greater extent than in Poland (86.0% vs. 47.5%). Interestingly, the contribution of real shocks to FEV of the relative price level in both economies was – unlike in three-variable VAR – quite similar. What was different was the importance of financial shocks: in Slovakia much higher than in Poland.

These findings are in line with the implications of the impossible macroeconomic trinity: under unfettered capital flows exchange rate stability can be attained at the cost of monetary autonomy. Since the exchange rate of the koruna has been stabilized, financial shocks have to be accommodated with changes in interest rate. It is also consistent with the analysis of residuals which made it clear that real interest rate differential in Slovakia was more volatile than in Poland.

Overall, even if financial shocks are explicitly included in the model the flexible exchange rate still does not seem to be the source of unnecessary volatility. In this respect our results are in line with these of Stążka-Gawrysiak (2009), however we find that financial shocks are much more important source of variation in the real exchange rate of the zloty, i.e. 30-40% and not 3-4% found by Stążka-Gawrysiak. Moreover, financial shocks cannot be 'switched off' by introducing more exchange rate stability. There is no free lunch: as the case of Slovakia shows financial shocks find their way to an economy through the real interest rate differential and the relative price level.

5.4 Exchange rates and the global financial crisis

The exchange rate behaviour in the pre-crisis and crisis periods can cast more light on the usefulness of flexible exchange rate regime. The initial observation is that the koruna and the zloty appreciated in tandem in the pre-crisis period. The average annual rate of real appreciation between the entry of the Slovak koruna into the ERM II (2005q4) and the beginning of the global financial crisis (2008q3) was 9.0% in Slovakia and 6.6% in Poland. Thus, the participation in the ERM II does not seem to be superior to exchange rate flexibility in holding back large, possibly excessive real appreciation.

In order to check whether the pre-crisis appreciation was excessive we examine its sources using historical simulations based on estimated VAR models. First, we compare the actual path of the real exchange rate with the path without any shocks. Simulations start in 2002q4. As it is clear from Figure 1 the appreciation of the koruna and the zloty in 2005-2006 was in line with the free-of-shocks paths but then in 2007 and 2008 both currencies surged much above these paths. When the financial crisis spread on the world economy at the end of 2008, the zloty depreciated substantially, overshooting its level without any shocks. The case of the koruna is different: it remained strong and the deviation subsided only gradually from more than 18 per cent in 2008q4 to around 1 per cent at the end of 2013.

FIGURE 1

To gain more insight into the nature of these deviations the importance of structural shocks is illustrated in Figures 2 and 3. The solid line corresponds to the overall deviation and dashed and dotted lines to the contribution of structural shocks to the deviation. For instance, in the left panel of Figure 2 the deviation of the actual real exchange rate of the zloty from the path without any shocks is illustrated with a solid line and the dashed and dotted lines correspond to the importance of supply and demand shocks, respectively.

FIGURE 2 and 3

In both economies demand and financial shocks played a crucial role in exchange rate fluctuations. For example, they contributed a lot to the large depreciation of the Polish zloty in the wake of 2001-2002 recession. Monetary shocks were unimportant: their contribution has been virtually nil. The importance of supply shocks increased during the global financial crisis, especially in Poland.

The pre-crisis real appreciation in Poland was driven mainly by financial shocks with a positive contribution from demand shocks only in the first half of 2008. In Slovakia contribution of demand shocks to appreciation was not smaller than that of financial shock. It is interesting to observe that in Poland the pre-crisis appreciation was reversed at the beginning of 2009 because financial shocks pushing for the appreciation dissipated and real shocks put a strong downward pressure on the zloty. After a rebound in 2009 the exchange rate remained under the influence of supply and demand shocks. Since the debt crisis in the euro area in 2011 not only real but also financial shocks pushed the real exchange rate down.

In contrast to Poland, the pre-crisis appreciation was not reversed in Slovakia. The real exchange rate stabilised since the euro adoption. Unfortunately, its level seems to be overvalued: by 11.6 per cent in 2008q4 if the path without demand shocks is treated as a benchmark or by 5.7 per cent for a baseline without financial shocks (other shocks are negligible). Our results are in line with other studies. For example, Sivák and Péliová (2012), using a different methodology, argued that the conversion rate of the koruna was set at an overvalued level. Interestingly, the misalignment implied by the benchmark without financial shocks is almost the same as the one identified by Dąbrowski (2009). He estimated the equilibrium exchange rate (based on the purchasing power parity adjusted for the Balassa-Samuelson effect) and found the conversion rate overvalued the koruna by 5-6%.

6 Conclusion

The general question that has motivated our analysis is whether the exchange rate acted as a shock absorber or rather a shock-propagating mechanism in Poland and Slovakia. We take the advantage of the opportunity to compare the results for relatively similar economies with respect to institutional framework, level of economic development and history but different strategies of monetary integration with the euro area. Our findings can be summarised in three main points.

First, building on the point raised by Artis and Ehrmann (2006) that the results obtained in any structural VAR exercise are only as good as the identification scheme that is employed to 'make sense' of the residuals, we examine three broad classes of structural VARs (two-, three- and four-variable models) with two types of restrictions (zero and sign restrictions). We find that the empirical support for both extreme hypotheses about the flexible exchange rate regime, i.e. that it is chiefly a shock-propagating mechanism and that it mainly acts as a shock absorber, rest on the imperfect identification of shocks. In a two-variable VAR neutral shocks are a far too heterogeneous category as they include a mixture of demand, financial and monetary shocks and in a three-variable VAR demand and/or monetary shocks contain an additive of financial shocks. Thus, our preferred model is the four-variable VAR.

Second, the evidence is found that the relative output and the exchange rate in Poland were under similar shocks (the correlation was strong and significant) whereas in Slovakia shocks behind fluctuations of the relative output seemed to be unrelated to those driving the exchange rate. This is also confirmed by the forecast error decompositions: the relative output is almost exclusively driven by supply shocks in both economies and a contribution of supply shocks to real exchange rate variability has been twice as large in Poland as in Slovakia (8% vs. 4%). Moreover, a considerable part of exchange rate variability (around 50%) is accounted for by demand shocks in both economies even if financial shocks are clearly separated from them.

Third, the previous point still leaves some room for the importance of financial shocks. Using historical simulations we demonstrate that they indeed contributed to the strong appreciation of the zloty in the run-up to the crisis. Thus, the argument that the flexible exchange rate regime resulted in an excessive appreciation is not groundless. There are, however, two important counterarguments that lend support to the viability of the exchange rate flexibility: Firstly, though financial shocks had weaker influence on the real exchange rate in Slovakia, it was not negligible and the koruna appreciated even more than the zloty. The participation in the ERM II did not protected against the strong appreciation. Secondly, when the crisis hit the flexible exchange rate proved to be useful: the depreciation of the zloty was helpful in mitigating the adverse supply and demand shocks. Since Slovakia entered the euro area in January 2009 the exchange rate misalignment could not be smoothed out with the depreciation of the koruna. It took five years to eliminate the real overvaluation of the koruna.

Let us close with a reservation about policy implications: We are far from arguing that the euro adoption is a bad option for an economy like Poland or Slovakia. Our analysis is focused on one aspect of monetary integration only, namely the usefulness of exchange rate flexibility. Thus, potential benefits of the euro adoption, e.g. lower transaction costs or greater price transparency, are not taken into account. Our point is rather that the flexible exchange rate has proved to be a useful mechanism of shock absorption in Poland, particularly during the global financial crisis whereas in Slovakia the burden of adjustment was on the real economy. Moreover, building on the Slovak experience the conversion rate should be chosen not so much with an eye to the current exchange rate but set at the level that is not distorted by monetary, financial and demand shocks.

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Table 1: Exchange rate regime and monetary policy framework in Poland and Slovakia

Period	Regime	Notes
De facto exchange rate regime in Poland:		
2000-2008	independently floating	till Apr 12, 2000 crawling band
2008-	free floating	reclassification reflects a methodological modification introduced by the IMF Sept 23, 2011 – Dec 31, 2011 'floating' due to interventions in the foreign exchange market
Monetary policy framework in Poland:		
1999-	inflation targeting framework	target: inflation below 4 per cent by the year 2003 and then 2.5 per cent
De facto exchange rate regime in the Slovak Republic:		
1998-2005	managed floating with no predetermined path for the exchange rate	The NBS does not support the koruna's exchange rate but intervenes primarily to smooth large fluctuations in the exchange rate and when the exchange rate moves to an unacceptable level
2005-2008	pegged exchange rate within horizontal bands	The Slovak Rep. entered the ERM II on Nov 25, 2005 with the central rate of the koruna set at 38.4550 SKK per 1 EUR and the standard fluctuation band of ±15% around the central parity Effective Mar 19, 2007 the central parity changed to 35.4424 SKK per 1 EUR and effective May 28, 2008 to 30.1260 SKK per 1 EUR.
2009-	free floating	The Slovak Rep. is a member of the euro area since Jan 1, 2009
Monetary policy framework in the Slovak Republic:		
2001-2005	other monetary policy framework	The country has no explicitly stated nominal anchor, but rather monitors various indicators in conducting monetary policy
2005-2007	inflation targeting and/or exchange rate anchor	The country adopts more than one nominal anchor in conducting monetary policy
2008	exchange rate anchor	
2009-	other monetary policy framework	The Slovak Rep. is a member of the euro area since Jan 1, 2009

Source: Data from the IMF AREAER, various issues.

Table 2: GDP growth and exchange rate variability

Variable / Period	2000-02	2003-05	2006-08	2009-11	2012-13
Poland [percentage points]					
GDP growth, average	2.4	4.2	5.9	3.2	1.8
GDP gap, average	-2.1	3.2	2.5	1.7	-0.4
GDP growth, variability	1.9	1.2	1.3	1.4	1.1
Real ex. rate, variability	5.7	6.1	5.1	5.6	1.2
Nominal ex. rate, variability	5.5	5.7	5.2	5.6	1.6
Slovakia [percentage points]					
GDP growth, average	3.0	5.3	7.9	0.8	1.4
GDP gap, average	-2.8	-0.6	6.2	-0.9	-2.7
GDP growth, variability	2.6	1.6	2.7	4.4	0.9
Real ex. rate, variability	3.2	2.1	3.0	2.2	0.6
Nominal ex. rate, variability	1.4	1.2	3.3	1.8	0.2

Table 3: Long-run identifying restrictions derived from the model.

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

Table 4: The most probable models in the two-variable systems

Poland		Slovakia	
model	$p(M_{(k,s)}\|Y)$	model	$p(M_{(k,s)}\|Y)$
(9,1)	0.215	(8,1)	0.281
(5,1)	0.211	(7,1)	0.261
(8,1)	0.201	(6,1)	0.200
(6,1)	0.186	(9,1)	0.177
(7,1)	0.182	(5,1)	0.081

Note: Prior probability of each specification: $p(M_{(k,s)}) = \frac{1}{10}$.

Table 5: Forecast error variance decomposition, two-variable VAR, one-year forecast horizon.

		Non-neutral			Neutral		
Variable	Model	50th	16th	84th	50th	16th	84th
Poland							
Relative output	2 vars	99.5	96.6	99.8	0.5	0.0	3.4
Nominal exchange rate	2 vars	6.8	1.0	17.9	93.2	82.1	99.0
Slovakia							
Relative output	2 vars	96.2	87.8	99.3	3.8	0.7	12.2
Nominal exchange rate	2 vars	5.9	0.0	22.0	94.1	78.0	99.6

Note: '2 vars' stands for two-variable model.

Table 6: The most probable models in the three-variable systems

Poland		Slovakia	
model	$p(M_{(k,s)}\|Y)$	model	$p(M_{(k,s)}\|Y)$
(5,2)	0.215	(8,2)	0.285
(9,2)	0.184	(9,2)	0.161
(8,2)	0.180	(7,2)	0.149
(6,2)	0.170	(6,2)	0.137
(7,2)	0.169	(5,2)	0.102

Note: Prior probability of each specification: $p(M_{(k,s)}) = \frac{1}{15}$.

Table 7: Forecast error variance decomposition, three-variable and four-variable VARs, one-year forecast horizon.

		Supply			Demand			Financial			Monetary		
Variable	Model	50th	16th	84th	50th	16th	84th	50th	16th	84th	50th	16th	84th
Poland													
Relative output	3 vars	99.2	95.9	99.9	0.5	0.0	2.9				0.0	0.0	0.8
	4 vars	99.4	96.5	99.8	0.0	0.0	1.3	0.0	0.0	1.1	0.0	0.0	0.5
Real interest rate diff.	3 vars												
	4 vars	1.1	0.0	4.5	49.5	9.5	89.5	47.5	7.5	87.5	0.0	0.0	0.4
Real exchange rate	3 vars	8.5	1.6	20.5	91.2	79.3	98.1				0.0	0.0	0.3
	4 vars	8.2	1.9	19.1	47.4	9.5	82.4	40.4	6.5	79.3	0.0	0.0	0.4
Relative price level	3 vars	1.4	0.2	5.7	1.2	0.1	4.8				96.2	90.3	98.9
	4 vars	1.2	0.0	5.7	1.6	0.0	5.3	0.7	0.0	4.0	94.1	87.4	97.8
Slovakia													
Relative output	3 vars	95.0	86.1	98.5	2.4	0.6	9.2				1.3	0.3	5.1
	4 vars	93.7	85.0	98.3	1.5	0.0	6.2	1.6	0.0	5.9	1.0	0.0	3.8
Real interest rate diff.	3 vars												
	4 vars	3.3	0.0	14.4	6.0	0.0	26.3	86.0	63.7	96.6	0.0	0.0	0.7
Real exchange rate	3 vars	3.7	0.4	15.0	96.0	84.9	99.4				0.0	0.0	0.3
	4 vars	3.7	0.0	13.4	58.1	30.3	78.0	34.5	14.6	62.4	0.0	0.0	0.5
Relative price level	3 vars	3.1	0.4	12.9	9.8	3.2	19.0				84.1	72.8	92.4
	4 vars	3.4	0.0	12.7	1.7	0.0	5.7	16.5	7.1	28.0	73.9	61.7	84.6

Note: '3 vars' and '4 vars' stand for three-variable and four-variable models, respectively.

Table 8: The most probable models in the four-variable systems with zero and sign restrictions

Poland		Slovakia	
model	$p(M_{(k,s)}\|Y)$	model	$p(M_{(k,s)}\|Y)$
(5,3)	0.218	(8,3)	0.226
(6,3)	0.204	(7,3)	0.169
(7,3)	0.189	(9,3)	0.157
(8,3)	0.187	(6,3)	0.150
(9,3)	0.186	(5,3)	0.089
		(8,2)	0.049
		(9,2)	0.045

Note: Prior probability of each specification: $p(M_{(k,s)}) = \frac{1}{20}$.

Table 9: Residuals from the four-variable VAR: variances and correlations.

	Poland			Slovakia		
	Average	Lower	Upper	Average	Lower	Upper
Variance of residuals from the equation for:						
Relative output, ϵ_y	0.0004	0.0003	0.0006	0.0006	0.0003	0.0008
Real interest rate diff., ϵ_r	0.0002	0.0001	0.0003	0.0005	0.0003	0.0006
Real exchange rate, ϵ_q	0.0022	0.0014	0.0031	0.0006	0.0003	0.0008
Relative price level, ϵ_p	0.0002	0.0001	0.0003	0.0003	0.0002	0.0004
Correlation of residuals:						
(ϵ_y, ϵ_r)	-0.0333	-0.2890	0.2224	-0.0110	-0.3158	0.2939
(ϵ_y, ϵ_q)	-0.2903	-0.5261	-0.0546	-0.0288	-0.3166	0.2590
(ϵ_y, ϵ_p)	-0.0773	-0.3303	0.1756	-0.0613	-0.3302	0.2076
(ϵ_r, ϵ_q)	0.0402	-0.2082	0.2887	-0.3598	-0.5897	-0.1299
(ϵ_r, ϵ_p)	0.0579	-0.1723	0.2882	-0.3623	-0.5919	-0.1327
(ϵ_q, ϵ_p)	0.1162	-0.1131	0.3456	0.3484	0.1419	0.5548

Note: Lower and upper bands are minus and plus 2 standard deviations, respectively.

Poland Slovakia

Figure 1: Real exchange rates in Poland and Slovakia.

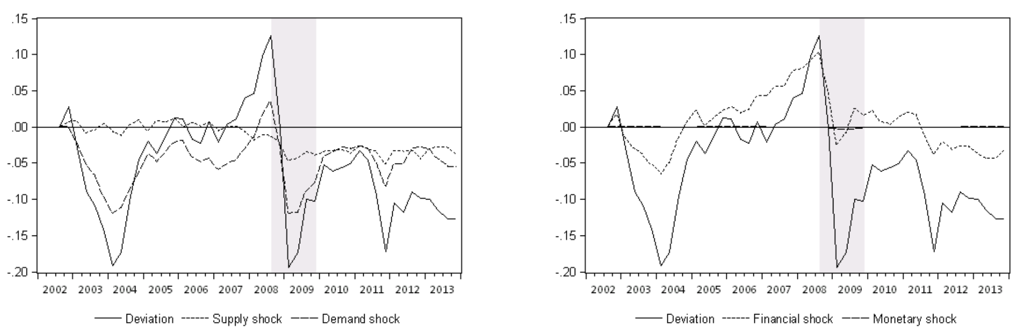

Figure 2: Real exchange rate fluctuations and shocks in Poland.

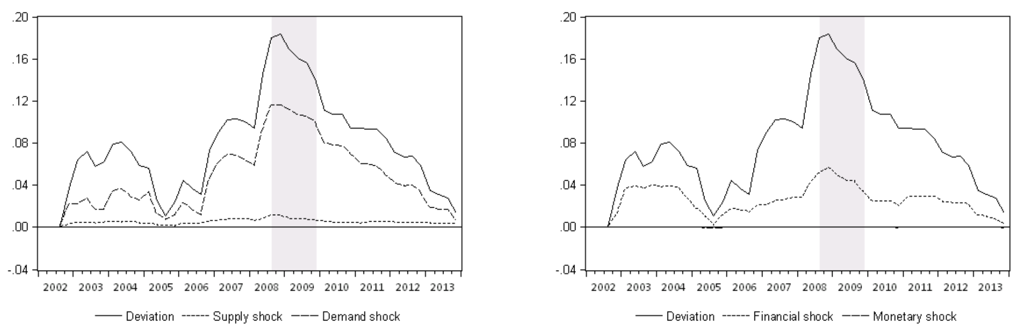

Figure 3: Real exchange rate fluctuations and shocks in Slovakia.

Supplementary material (for on-line publication only)

Appendix to Exchange rate as a shock absorber or a shock propagator in Poland and Slovakia – an approach based on Bayesian SVAR models with common serial correlation

Arias – Rubio-Ramírez – Waggoner identifying procedure

The procedure constructed by Arias, Rubio-Ramírez and Waggoner (2014) enables to impose both sign and zero restrictions on impulse response functions of any horizon(s). We stack them in a matrix $F(.,.)_{hn\times n}$, where h denotes the number of horizons we impose restrictions on. The matrix $F(.,.)$ is a function of the structural-form parameters. Describing the adopted algorithm we use the notation introduced by Arias et. al. (2014). The zero restrictions imposed on the $j-$th structural shock are represented by matrices Z_j, $1 \leq j \leq n$ whereas sign restrictions - by matrices S_j, $1 \leq j \leq n$. The number of columns in Z_j and S_j equals the number of rows in F. The rank of S_j is s_j and the rank of Z_j equals z_j, so $s_j + z_j$ is the number of sign and zero restrictions placed on the responses to the $j-$th structural shock. The structural parameters satisfy the sign restrictions if $S_j F(.,.)e_j > 0$ and the zeros restrictions if $Z_j F(.,.)e_j = 0$ for $1 \leq j \leq n$, e_j denotes the j-th column of the identity matrix of order n (I_n).

The algorithm consists of the following steps (see Arias et. al. (2014) for more detailed description and explanation):

1. Draw Γ and Σ from the posterior distribution of the reduced-form parameters.

2. Perform the Cholesky decomposition of Σ: $\Sigma = T'T$, where T is upper triangular with positive elements on the main diagonal. Matrices T^{-1} and ΓT^{-1} are the recursive-form parameters.

3. Draw an orthogonal matrix Q from the uniform distribution with respect to the probability measure defined on the set of orthonormal matrices conditional on the zero restrictions, i.e. such that $(T^{-1}Q, \Gamma T^{-1}Q)$ satisfies the zero restrictions. The proposed algorithm is as follows (see Theorem 3 in Arias et. al., 2014): Set $j = 1$ and do the subsequent steps as long as $j < n$:

   (a) Find a matrix N_{j-1} forming an orthonormal basis for the null space of $\begin{pmatrix} Z_j F(T^{-1}, \Gamma T^{-1}) \\ Q'_{j-1} \end{pmatrix}$, with the assumption that Q_0 is an empty matrix.

   (b) Draw y_j form the standard normal distribution on $\mathbb{R}^{n_j}$, where $n_j > 0$ is the number of columns in N_{j-1}.

(c) Let $q_j = N_{j-1}y_j/||y_j||$.

(d) Let $j = j + 1$ and if $j < n$ return to step (a), otherwise stop.

4. Keep the draw if $S_j F(T^{-1}Q, \Gamma T^{-1}Q)e_j > 0$ are satisfied for $1 \leq j \leq n$.

5. Return to step 1 until the required number of posterior draws satisfying both the sign and zero restrictions has been obtained.

In the four-variable systems to place the restrictions we employ the following matrices: $Z_1 = \begin{pmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \end{pmatrix}$, $Z_2 = \begin{pmatrix} 1 & 0 & 0 & 0 \end{pmatrix}$, $Z_3 = \begin{pmatrix} 1 & 0 & 0 & 0 \end{pmatrix}$ for five excluding restrictions and $S_1 = \begin{pmatrix} 0 & 0 & 0 & 1 \end{pmatrix}$, $S_2 = \begin{pmatrix} 0 & 1 & 0 & 0 \\ 0 & 0 & -1 & 0 \end{pmatrix}$, $S_3 = \begin{pmatrix} 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{pmatrix}$, $S_4 = \begin{pmatrix} 1 & 0 & 0 & 0 \end{pmatrix}$ for sign restrictions (the shocks' order is as follows: monetary, financial, demand, supply).