Sources of real exchange rate variability in Poland – Evidence from a Bayesian SVAR model with Markov Switching Heteroscedasticity

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Abstract

This paper investigates the sources of real exchange rate fluctuations in Poland in 2000-2016. The objective is to assess the relative importance of cost, demand and monetary shocks in driving the exchange rate in Poland. A two-country and two-good New Keynesian open economy model as developed by Engel and West (2006) is used as a theoretical framework. A Bayesian SVAR model with Markov switching heteroscedasticity is used in empirical part. The structural shocks are identified on the basis of the changes in volatility and named with reference to the sign restrictions derived from the economic model. We identify two regimes/states: one with high volatility and the other featuring low volatility. Estimated impulse response functions are in line with the theoretical model though uncertainty is rather large. The main finding is that the contribution of cost and demand shocks to exchange rate variability is about 50 percent in normal times implying that the flexible exchange rate acts as a shock absorber. In turbulent times monetary shocks dominate the exchange rate variability, but their impact on a real economy is short-lived. This undermines the claim that the exchange rate flexibility is dangerous due a shock-propagating nature of the exchange rate.

Keywords: *real exchange rate, open economy macroeconomics, Bayesian MS-VAR models*

JEL Classification: F31, F41; C11

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1 Introduction

Poland is a small open economy with flexible exchange rate regime. It is open both to trade and financial flows: openness to trade as measured by a ratio of the sum of exports and imports to GDP was more than 100 percent in 2016 and the capital openness was above the world median in 2014 (as measured by the Chinn-Ito index; see Chinn and Ito, 2008). According to the IMF's exchange-rate-regime classification, the Polish currency freely floats since April 12, 2000.

The objective of this research is to examine the relative importance of cost, demand and monetary shocks in driving the exchange rate in Poland, and thus to provide evidence that is useful in assessing the desirability of the exchange rate flexibility. Existing evidence is far

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from unambiguous. In some studies nominal shocks are found to be the main driver of exchange rate flexibility (see, e.g., Shevchuk, 2015), but in others real shocks are found to be the primary source of exchange rate variability (see, e.g., Arratibel and Michaelis, 2014; Dąbrowski and Wróblewska, 2016). Yet another view was expressed by Alexius and Post (2008) who found that ‘exchange rates display some stabilizing properties but can mainly be characterized as disconnected from the rest of the economy.’[4] We use a Bayesian Markov Switching Heteroskedastic VAR model to identify shocks hitting an economy. Overidentifying sign restrictions derived from a New Keynesian small open economy model are imposed to obtain economically interpretable shocks.

2 Theoretical framework and data

A small open economy model is used as a theoretical model. It is a two-country and two-good New Keynesian model developed by Engel and West (2006). The model consists of four main equations: interest rate rule, Phillips curve, IS relation and uncovered interest rate parity condition. It is assumed that the domestic central bank cares about the exchange rate stability, so the exchange rate term enters the interest rate rule. There are three exogenous stationary AR(1) disturbances: demand, cost and monetary that are driven by respective structural shocks. Each variable is defined as a difference between a domestic variable and a foreign one. This allows us to focus on asymmetric shocks, i.e. the ones that matter when the usefulness of the exchange rate flexibility is to be assessed (when shocks are symmetric the flexible exchange rate is of limited use). As pointed out by Artis and Ehrmann (2006), the drawback of such an approach is that it ‘yields no information on the comparative frequency of symmetric and asymmetric shocks’ and only the latter ‘necessitate exchange rate adjustments.’ The approach we follow, however, is much more parsimonious, well-established in the literature and corresponds to the nature of the exchange rate which is itself a relative variable.[5]

A system of stochastic difference equations yields the solution that allows us to determine the direction of the short-term reactions of three main endogenous variables to structural shocks. The relative output gap reacts positively to a demand shock and negatively to cost and monetary shocks. All shocks have a positive impact on the real exchange rate (an increase

[4] The result, however, was obtained for advanced small open economies, Canada and Sweden.
[5] Moreover, when a shock is symmetric, a flexible exchange rate can be useful if domestic and foreign preferences with respect to its accommodation are different.

corresponds to an appreciation of domestic currency). The relative inflation gap's response to both demand and cost shocks is positive, but it is negative to a monetary shock.

We use quarterly data for Poland and the euro area (EA-12) spanning the period 1995q1-2017q2. The data for the underlying variables were retrieved from the Eurostat database and include: real GDP, implicit GDP deflator, average euro/ECU exchange rate and harmonised index of consumer prices (HICP). These data are used to construct three main endogenous variables. The relative output gap is a difference between domestic and foreign output gap and each gap is a cyclical component of the log of relevant GDP identified with the HP filter. The real exchange rate (RER) is constructed on the basis of the (index of the) nominal exchange rate and (seasonally adjusted) HICP indices for Poland and the euro area and then a cyclical component of the log of RER is extracted with the HP filter. The relative inflation gap is a difference between domestic and foreign inflation gaps and each gap is a cyclical component of relevant QoQ inflation (based on the implicit GDP deflator) identified with the HP filter.

3 Methodology

We base our empirical analysis upon Bayesian VAR(2) model with Markov switching heteroscedasticity (henceforth MSH-VAR(2)). As we also take into account the possibility of cointegration, the basic model is presented in the VEC form:

$$\Delta y_t = \alpha\beta' y_{t-1} + \Gamma_1 \Delta y_{t-1} + \nu + \varepsilon_t, \quad \varepsilon_t | S_t \sim iiN(0, \Sigma_{S_t}), \quad t = 1,2,\dots,T, \tag{1}$$

where $y_t = [Output\_GAP_t \quad Ex\_Rate_t \quad Inf\_GAP_t\ ]'$ represents the column vector of the endogenous variables introduced in the previous section, and $\{S_t,\ t \in \mathbb{Z}\}$, $S_t \in \{1, 2\}$, is a two-state homogenous and ergodic Markov chain with transition probabilities denoted by p_{ij} ($\forall_{t\in\mathbb{Z}}\, P(S_t = j | S_{t-1} = i) = p_{ij},\ p_{i1} + p_{i2} = 1,\ p_{ij} \in (0,1),\ i,j \in \{1, 2\}$). Matrices α and β collect the adjustment coefficients and the cointegration vectors, respectively.

For the covariance matrices we impose the inverted Wishart priors $\Sigma_1 \sim iW(10S_\Sigma, n + 11)$, $\Sigma_2 \sim iW(5S_\Sigma, n + 11)$, where hyperparameter S_Σ is obtained fa training sample covering the period 1996q2-1998q3. The difference between the priors assumed for Σ_1 and Σ_2 reflects our prior belief according to which the second-state variance is lower than in the first state. The joint prior for α, Γ_1 and ν is matrix normal:

$$(\alpha, \Gamma_1, \nu)' \sim mN\left(0, I_3, \begin{pmatrix} 0.01I_3 & 0_{3\times4} \\ 0_{4\times3} & 0.5I_4 \end{pmatrix}\right).$$

Finally, for the transition probabilities we impose uniform distribution $p_{ii} \sim U(0,1) = Beta(1,1), i = 1,2$. It has turned out that the analysed data may be treated as a realisation of covariance stationary three-dimensional VAR(2) process, so we assume $\beta = I_3$. The joint prior distribution is truncated by the stability condition. The states are identified by the assumption that the conditional volatility of exchange rate in the first state is higher than in the second one.

The reduced-form errors (ε_t) are linear combinations of the structural shocks (denoted as u_t), i.e. $\varepsilon_t = Bu_t$. To identify the structural shocks we employ the method proposed by Lanne et al. (2010), in which we make use of the changes in covariance matrices (Σ_1, Σ_2) of the reduced-form shocks. In the two-state case there always exist matrices B and Λ_2 such that

$$\Sigma_1 = BB', \qquad \Sigma_2 = B\Lambda_2 B', \qquad \Lambda_2 = diag(\lambda_{21}, \lambda_{22}, \lambda_{23}). \tag{2}$$

Lanne et al. (2010) show that this decomposition is locally unique, i.e. up to changing the signs and ordering of B's columns, if the diagonal elements of Λ_2 are all distinct. Using these results the structural shocks defined as $u_t = B^{-1}\varepsilon_t$ are locally identified, so to obtain global identification we additionally impose the shocks' order and their signs by assuming the descending order for $\lambda_{2i}, i = 1, 2, 3$, and $B_{11} < 0$, $B_{22} > 0$, $B_{33} < 0$, with B_{ii} (i = 1, 2, 3) denoting the i-th diagonal element of B. The three latter inequalities, constituting normalisation restrictions, are set according to the economic model underlying this study, and are also further employed in another MSH-VAR(2) model, which we discuss below. It is worth emphasising that structural shocks obtained in this framework feature switching (therefore, time-varying) covariance matrices (I_3 in the first state and Λ_2 in the second one), but their impacts remain unchanged. Also, since Λ_2 is diagonal, the structural shocks are orthogonal also in the second state, with $\lambda_{2i}, i = 1, 2, 3$, being their variances. Finally, since $Var(u_{ti}|S_t = 1, \theta) = 1$ and $Var(u_{ti}|S_t = 2, \theta) = \lambda_{2i}$, each λ_{2i} may be regarded a relative change of u_{ti}'s variance in the second state with respect to its (unit) value in the first regime.As mentioned above the structural shocks can be statistically identified if their variances in the second state are all distinct. Therefore, in Table 1 we inspect the posterior quantiles for $\lambda_{2i}, i = 1, 2, 3$, and their differences.

It can be noticed that the 90%-credible intervals of variances overlap and the posterior marginal densities of their differences are not well-separated from zero, so the conditions for local identification appear not fulfilled. For this reason we have decided to facilitate the identification of structural shocks in our study by introducing sign restrictions (posited by the economic model) for some other (rather than only diagonal) elements of *B*. Such an approach, formally resulting in a different Bayesian model (further referred to as the second model), combines the purely statistical procedure delivered by the model we discussed hitherto, with one of the most common approach to shock identification by sign restrictions (in accordance with the economic model). We refer the reader to Herwartz and Lütkepohl (2014) for a broader discussion on combining identification by heteroscedasticity with typical short-run restrictions, though only in the frequentist (rather than Bayesian) setting.

Table 1. Posterior quantiles of the structural shocks' variances in the second state.

probability	**0.05**	**0.16**	**0.5**	**0.84**	**0.95**
λ_{21}	0.408	0.507	0.716	1.030	1.325
λ_{22}	0.131	0.177	0.283	0.429	0.547
λ_{23}	0.047	0.060	0.086	0.124	0.157
$\lambda_{21} - \lambda_{22}$	0.110	0.203	0.415	0.727	1.009
$\lambda_{22} - \lambda_{23}$	0.047	0.087	0.188	0.336	0.456
$\lambda_{21} - \lambda_{23}$	0.310	0.412	0.625	0.941	1.237

As regards our particular choice of sign restrictions for the elements of *B* (i.e. instantaneous effects of the shocks upon the endogenous variables), we confronted their posterior distributions in the model discussed above (results available upon request) with reactions predicted by the economic theory, and it enabled us to name the shocks and then to combine both types of identification restrictions. Ultimately, we constrained only (immediate) reactions to the cost and monetary shocks, corresponding to the first and third column of *B* (with imposing the same signs upon B_{11}, B_{22} and B_{33} as previously). Fig. 1 displays the priors and posteriors of *B*'s elements in the resulting (second) model, with sign restrictions (postulated by the theory) indicated in parentheses (in cases when the sign is not actually imposed in the model) and brackets (when the sign is imposed in the model). It is also worth mentioning that the posterior distributions of *B*'s elements in the previous model, i.e. the one without additional sign constraints, tend to feature slight bimodality, with, and most importantly, the prominent mode in each distribution being supported by the economic model (results available upon request). Introducing additional restrictions removed the bimodalities almost completely, perhaps with a sole exception of B_{12} (see Fig. 1).

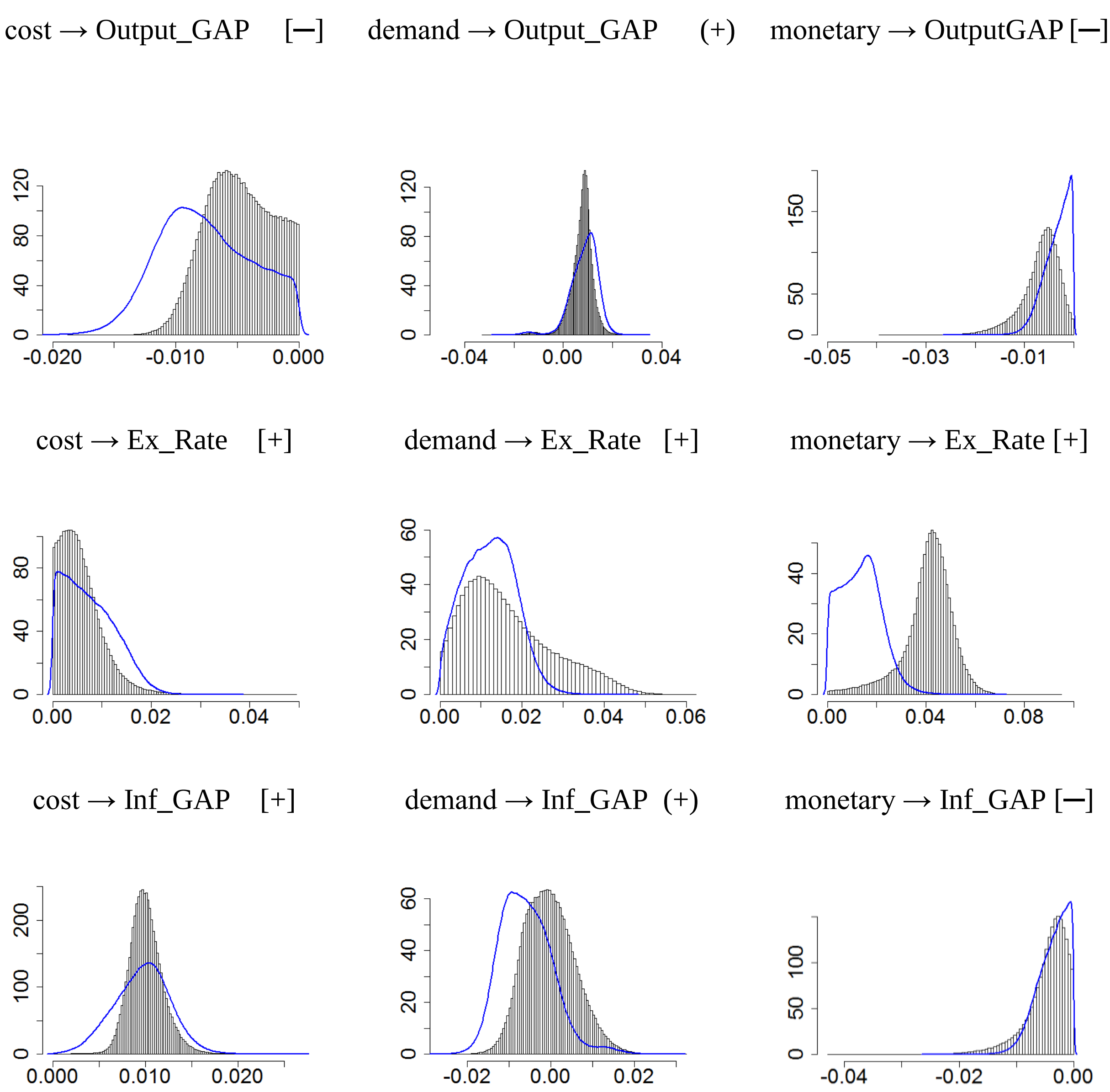

Fig. 1. Marginal posteriors (histograms) and priors (lines) of the elements of the instantaneous reactions matrix (B) obtained within the second model.

4 Empirical results

In this section we focus only on the results obtained within the second model, i.e. the one with additional sign restrictions imposed. Conventional tools of structural vector autoregressive analysis were used, i.e. structural impulse responses and forecast error variance decompositions. Our research question is whether fluctuations in the exchange rate in Poland were related to changes in the real economy or were driven by volatility in financial markets. In the former case the exchange rate flexibility can be considered a useful adjustment mechanism, whereas and in the latter a factor that contributes to the transmission of undesired financial volatility into the real economy.

In Fig. 2 the impulse response functions of the real exchange rate to identified shocks over five-year period are illustrated (results for other variables are available upon request). Solid lines display the posterior mean reactions, whereas and broken lines represent the intervals of plus/minus two standard deviations. All reactions are positive and rather smooth: after a shock the output gap, real exchange rate and inflation gap gradually return towards their long-run paths. In principle, the IRFs obtained are consistent with the model of a small open economy, although the degree of uncertainty if rather large. It is worth emphasizing that sign restrictions are imposed only on the instantaneous reactions (see Fig. 1) and all other (i.e. the lagged) reactions are left unrestricted, so the trajectories of the IRFs are not determined a priori. The IRFs to demand and cost shocks are in line with the shock absorbing property of the flexible exchange rate. In the wake of a positive demand shock the aggregate demand and output increase, but their reaction is mitigated by a real appreciation of the domestic currency. Interestingly, the inflation gap remains almost unchanged, which means that the adjustment process is through the exchange rate changes. This could be an important advantage if prices are sticky. A positive cost shock raises inflation gap and thus induces a real appreciation of the domestic currency. These changes, together with a rise in the real interest rate due to the central bank’s focus on price stability, contribute to a decrease in the aggregate demand and output. Since the shock is temporary, the change in the real exchange rate will fade away in the long term (see also Fig. 2). This, however, means that the domestic currency needs to depreciate in nominal terms both in the short and long run. Without the exchange rate flexibility the nominal depreciation would not be possible, and for the real exchange rate to return to its long-run path a period of low domestic inflation or even deflation would be required.[6] Interestingly, the real exchange rate is found to be back to its long-term path faster than both the output and inflation. The implication is that the *nominal* exchange rate flexibility remains important beyond the horizon of the adjustment of the *real* exchange rate. A positive monetary policy shock results in a strong real appreciation of the domestic currency which strengthens an adverse impact of the shock on the output. Moreover, the reaction of inflation is almost non-existent, so the induced appreciation does not seem to be conducive to a reduction of inflation (through the exchange rate channel of monetary

[6] More precisely, the domestic inflation would have to be lower than the foreign inflation. This, however, given that the inflation abroad were low, would indeed require a domestic *deflation*.

transmission). Thus, the IRFs to a monetary shock are consistent with the claim that the exchange rate propagates a shock to the real economy.

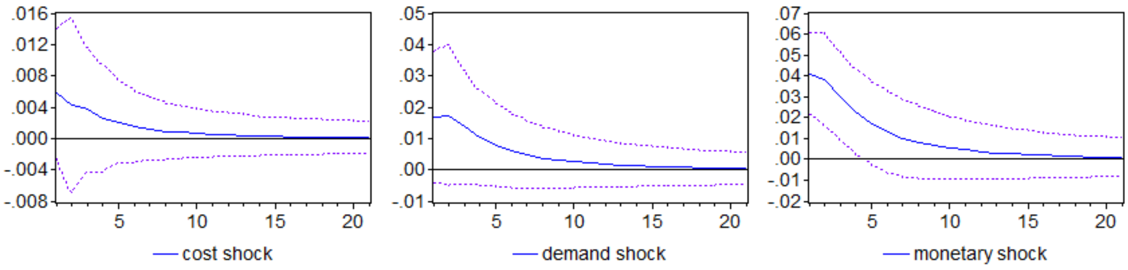

Fig. 2. Impulse response functions of the real exchange rate in Poland.

Table 2. Posterior means of forecast error variance decomposition of the relative output and real exchange rate and relative inflation in Poland.

Forecast	**Shocks in turbulent times**			**Shocks in normal times**		
horizon	**cost**	**demand**	**monetary**	**cost**	**demand**	**monetary**
			Relative output gap			
0	17.5	49.0	33.5	35.7	52.0	12.3
20	12.3	36.1	51.6	32.0	43.7	24.3
			Real exchange rate			
0	2.4	18.9	78.7	11.7	31.5	56.8
20	2.5	20.8	76.7	12.4	33.9	53.7
			Relative inflation gap			
0	69.4	16.1	14.5	87.6	9.5	2.9
20	33.8	22.6	43.6	64.7	20.0	15.3

Before the importance of shocks is assessed, the two volatility regimes need to be discussed. Our empirical approach enables us to identify two distinct states: one with high volatility and the other with low volatility. The former, which we refer to as ‘turbulent times,’ comprises (the posterior probability of the regime above 0.5): (1) the EU pre-accession period (2000q1-2005q2), (2) global financial and economic crisis (2007q4-2010q1), and (3) the debt crisis in the euro area (2011q3–2012q2). Interestingly, our findings are consistent with the evolution of the common measure of volatility, i.e. the VIX index (results available upon request). The second regime encompasses the remaining periods and is henceforth labelled as ‘normal times.’

In order to assess the contribution of shocks to the variability of output, real exchange rate and inflation, we analyse the forecast error variance decomposition. Table 2 reports the results obtained for two forecast horizons: zero quarters (the short run) and twenty quarters (the long run). The results depend on the volatility regime. During the normal times the relative output gap is driven mainly by demand shocks, the relative inflation gap by cost shocks, and the real exchange rate by monetary shocks. In turbulent times a contribution of monetary shocks is much larger than in tranquil times, whereas that of cost shocks decreases substantially. Two more important observations can be made with reference to the usefulness of exchange rate flexibility. First, even though the variability of the real exchange rate is mainly driven by monetary shocks, the contribution of the two other shocks is close to 50 percent in normal times. Bearing in mind that these shocks account for 75-90 percent of output variability, one can argue that the flexible exchange rate acts as a shock absorber. Second, in turbulent times, when monetary shocks become more prevalent, the exchange rate acts more like a shock propagator. The cost and, in particular, demand shocks remain, however, important sources of both exchange rate variability (20-25 percent) and output variability (50-65 percent). Therefore, in turbulent times, the shock-absorbing property of the flexible exchange rate is limited rather than non-existent. Moreover, the reaction of the output to a monetary shock, as measured by the IRFs, is rather short-lived (after six quarters it peters out), so the shock-propagating property of the flexible exchange rate should not be overstated.

Conclusions

Our main findings can be summarised as follows. First, monetary shocks account for about 50 percent of exchange rate variability, but the contribution of cost and demand shocks is large, especially in normal times. Together with the theory-consistent trajectories of impulse response functions, the result lends support to the hypothesis that the flexible exchange rate in Poland acted as a shock absorber. Second, in turbulent times the relative importance of monetary shocks is much higher than in tranquil times. We abstain, however, from interpreting this finding as an argument in favour of the hypothesis that the exchange rate acted as a shock propagator, because the impact of a monetary shock on the real economy was found to be rather short-lived, unlike that of either cost or demand shock. We see two things that can be done to strengthen the results. First, our approach needs to be extended to allow for identification of financial shocks that are presumably hidden behind monetary shocks.

Second, in order to establish more convincing insights into the role of the exchange rate, we need to compare countries with fixed and flexible exchange rate. We leave these issues for further research.

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