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A small New Keynesian model to analyze business cycle dynamics in Poland and Romania

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ABSTRACT

In this paper, we derive a small textbook New Keynesian DSGE model to evaluate Polish and Romanian business cycles during the 2003-2014 period. Given the similarities between the two economies, we use an identical calibration procedure for certain coefficients and marginal prior distributions for the others, rendering the resulting cross-country differences as primarily data-driven. The estimated structural coefficients for the two countries have comparable values, implying similar qualitative macroeconomic transmission mechanisms. However, the Romanian shocks display much more variability, and the impulse response functions have similar shapes but deeper trajectories. The model-simulated theoretical moments for the output growth, the inflation rate and the nominal interest rate (means, standard deviations and cross-correlations) are close to their actual data counterparts, demonstrating the models' ability to match and replicate statistical properties of the observed variables. Shock decompositions of the output and the inflation rate revealed the driving forces of the business cycles; demand shocks explain much of the GDP growth dynamics (persistent positive contributions before the crisis and negative thereafter), whereas prices were also driven by supply and monetary policy shocks, the latter being more important for Poland.

KEY WORDS: New-Keynesian model, DSGE, business cycle, Bayesian estimation, shock decomposition

JEL Classification: C11, C32, E32

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Introduction

In recent years, New Keynesian Dynamic Stochastic General Equilibrium (DSGE) models have become standard workhorses in analyzing business cycle fluctuations in both academic and institutional (particularly central bank) environments. These models' theoretical advantages originate from microeconom-

ic optimizations and the immunity to the Lucas critique (because of the assumed rational expectations behavior of the agents) and have been augmented by practical advances in Bayesian estimation and computer power.

The origins of the DSGE models are usually associated with the Real Business Cycle (RBC) literature, and more precisely, with the seminal contribution of Kydland and Prescott (1982). Given the assumption that the agents are characterized by rational expectations, agents re-optimize their decisions following any shock, rendering the model invulnerable to the Lucas

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(1976) critique. Calibrating most parameters and estimating those remaining, the Kydland and Prescott (1982) model-implied theoretical moments and cross-correlations were remarkably compatible with actual United States data.

A harsh hypothesis adopted by the RBC school was the impossibility of economic policymakers to affect real variables. The New Keynesian paradigm restored monetary non-neutrality by acknowledging the short-term capability of a central bank or a fiscal authority to influence the output, given the existence of temporary price rigidity. Mankiw and Romer (1991) provide an ample overview of the New Keynesian literature, covering both general and specific features of this economic theory stream.

Literature regarding modern DSGE models is composed of theoretical derivations and empirical estimations for different scale models. Ireland (2004) and An and Schorfheide (2007) derive and estimate small New Keynesian models with sticky prices using three observable variables (similar to the model employed in this paper). The next generation of DSGE models includes Smets and Wouters (2003) and Christiano, Eichenbaum and Evans (2005), whose models encompass additional nominal and real rigidities, such as consumption habit, capital depreciation, investment adjustment costs, price indexation, and sticky wages. The estimated models of the early 2000s associated the DSGE environment as a powerful modelling device, and stimulated the adoption of these tools by actual policymakers for real-life/real-time economic policy design and forecast. EAGLE of the European Central Bank (Gomes, Jacquinot, & Pisani, 2010) and Ramses of the Sveriges Riksbank (Adolfson et al., 2013) are a few examples.

The labor market block was enriched similar to Erceg, Henderson and Levin (2010) monopolistic labor supplying households and sticky wages, and Mortensen and Pissarides' (1994) search and matching framework. The financial frictions gained a reputation after the late 2000s crisis, given it was driven and propagated within domestic and international financial flows. However, the financial accelerator mechanism was introduced and formalized much earlier, in Bernanke, Gertler and Gilchrist (1999). In the Bernanke et al.'s (1999) framework, entrepreneurs borrow money proportionally to their net worth levels to finance capi-

tal acquisitions, but are subject to idiosyncratic risk shocks that affect their ex-post returns.

If the DSGE model is to be used for an emerging and/or small open economy, certain distinct features should be addressed. First, the roles of the exchange rate, the currency risk premium and trade flows are expected to be more important for the model's dynamics, as in Adolfson et al. (2007). Second, to the extent that a significant amount of domestic liabilities are denominated in a foreign currency (i.e., the dollarization phenomenon), the currency substitution effects may play an important role for both nominal and real variables' evolution, as in Castillo, Montoro and Tuesta (2006). Thus, the current strand is represented by large-scale models, featuring distinctive economic sectors, an open economy dimension, and financial and labor market frictions, which include a large variety of structural shocks, as in Christiano, Trabandt and Walentin (2011).

Two distinct methods are usually applied when evaluating a DSGE model. Calibration has been used at least since Kydland and Prescott (1982). Calibration implies fixing certain parameters to certain values, which are derived outside the model, but have meaningful interpretations (such as matching certain moments in the data or ensuring particular steady state values). Combined with the accelerated development of computer power in recent decades, estimation became the preferred approach to link data to the model equations. Among the estimation approaches, the maximum likelihood and Bayesian methods (which allow the inclusion of non-sample information via prior distributions for the parameters of interest) remain dominant, with the latter recently gaining increased popularity; refer to DeJong and Dave (2007) for extended reviews and technical details. However, the datasets of observable variables usually do not allow for proper identification of all the parameters; therefore, calibrating certain coefficients and estimating the others represents a common procedure to follow.

Despite the general consensus achieved in the literature regarding the usefulness of the DSGE models, certain limitations exist. The oversimplifying assumptions when modelling the real economy structure were partially resolved because of recent developments, which insert financial and labor markets into the

model (as noted above), but at the cost of increased complexity. Additionally, when compared to reduced-form models, such as Vector Autoregressions (VAR), DSGE models often lose in terms of data fit and forecasting accuracy. Pagan (2003) argues VAR models display an increased degree of empirical coherence (and match actual data well), whereas DSGE models inherit a higher degree of theoretical coherence (given rich parameter restriction structures), but are not very compatible with the data.

In this paper, we derive a small-scale New Keynesian DSGE model. The reduced dimension offers increased flexibility and tractability, in contrast to larger models with many structural shocks. Additionally, a simple model environment and stochastic structure facilitates estimation for short samples. As opposed to Ireland (2004) and An and Schorfheide (2007), we explicitly include the consumption habit in the households utility function. Additionally, we consider Calvo's (1983) mechanism for staggered price setting in a monopolistic environment instead of the ad-hoc price adjustment costs function as in Rotemberg (1992). The resulting hybrid aggregate demand and Phillips curves, completed by an interest rate smoothing reaction function as in Taylor (1993) for the central bank, determines a satisfactory degree of persistence. The stylized economy is perturbed by three structural innovations: a demand/consumption preferences shock, a supply/technology shock, and a monetary policy shock.

We estimate the model using Polish and Romanian data, with the following observed variables for 2003Q1 – 2014Q3 period: quarterly real GDP growth, the harmonized index of consumer prices (HICP) quarterly inflation rate, and the 3-month money market nominal interest rate. We consider identical calibration procedures and prior distributions for both economies, because we are interested in the marginal contribution of the data to the cross-country differences. Additionally, similarities of the Polish and Romanian economies support the approach we undertake. Both are small open economies with a floating exchange rate regime, similar degrees of openness (as measured by the imports plus exports to GDP ratios), high integration with the Euro area (in terms of goods and services trade, but also financial flows), with an inflation targeting monetary policy strategy (Poland

since 1998 and Romania since 2005; hence, the central bank's credibility level is likely to differ). In this respect, we proceed in accordance with Smets and Wouters (2005), who compare identical medium-scale DSGE models for the Euro area and the United States.

The record of estimated DSGE models for Poland and (particularly) Romania is relatively scarce. Caraiani (2007) uses a model similar to our model to reveal moderate price stickiness and strong central bank response to both inflation and output gap dynamics in Romania. However, we differ in considering the consumption habit and inflation indexation (rendering an increased inertia) and, more importantly, in specifying well-identified demand and supply shocks, whereas Caraiani (2007) includes two supply shocks (technology and inflation) and no demand shock. Caraiani (2008) adds the small open economy dimension to a textbook DSGE model to find that Euro area variables are important for domestic economic developments. Grigoraş (2010) estimates the Adolfson et al. (2007) model using Romanian data and concludes it produces a reasonable in-sample fit. Regarding Poland, Kolasa (2013) uses a DSGE model for a business cycle accounting procedure to identify the wedges that drive the output in five new European Union member states (including Poland). The central bank of Poland uses a large scale DSGE model for its policy analysis and decision making processes, described in Grabek, Kłos and Koloch (2011). Caraiani (2013) presents empirical evidence for both Romanian and Polish central banks' response to the exchange rate, in addition to inflation and the output gap.

The objectives of this paper are the following. First, we assess the economic transmission mechanisms implied by the estimated models (resulting from the impulse response functions) and comment on their differences and compatibility with the economic theory. Second, we compare the data-implied moments for the observed variables and their model-derived counterparts, discussing the ability of the DSGE model to fit the data and filter consistent unobserved variables. Third, we compare the sources of business cycle fluctuations during the last decade in the two economies using the shock and variance decompositions of the observed variables. Thus, the testing inventory is very diverse and is capable of revealing general and specific features of the models.

The model

The model we employ has a standard textbook small New Keynesian structure, similar to Clarida, Gali and Gertler (1999), Ireland (2004) or An and Schorfheide (2007). In spite of its simple architecture and reduced dimensions, the model embeds certain features and rigidities present in medium- and large-scale seminal DSGE models, such as Smets and Wouters (2003), Christiano et al. (2005), Adolfson et al. (2007) or Christiano et al. (2011).

The stylized economy is populated by a large number of identical households. The homogeneity of these is guaranteed by the assumption of certain perfect consumption insurance that can be traded between the households, making the application of the representative agent framework possible. The issue each household encounters is represented by the following utility maximization problem:

$$E_0 \max \sum_{t=0}^{\infty} \beta^t [U(C_t, C_{t-1}) - h_t] a_t \tag{1}$$

where $U(C_t, C_{t-1})$ is the utility function of current and past consumption good C_t purchased from a final good producer, h_t measures the labor induced disutility, and a_t is a consumption preference shock that affects the intertemporal consumption allocation. We assume the following functional forms for $U(C_t, C_{t-1})$ and a_t :

$$U(C_t, C_{t-1}) = \frac{(C_t - bC_{t-1})^{1-\sigma_C}}{1-\sigma_C} \tag{2}$$

$$\ln a_t = \rho_a \ln a_{t-1} + \varepsilon_{at}, \quad \varepsilon_{at} \sim N(0, \sigma_a) \tag{3}$$

with b representing the consumption habit coefficient and σ_C the relative risk aversion or (inverse) elasticity of intertemporal substitution. The consumption preference shock follows an AR(1) process with autoregressive parameter ρ_a and normally distributed innovations ε_{at} .

The household maximizes (1) subject to the budget constraint (4):

$$T_t + D_t + B_{t-1} + W_t h_t \geq P_t C_t + \frac{B_t}{i_t} \tag{4}$$

with the following notations: T_t : lump-sum government transfers, D_t : profits accruing from the monopolistic intermediate goods producers ownership (e.g., dividends), B_{t-1} : revenue originating from purchas-

ing nominal bonds in period $t - 1$, W_t : nominal wage, P_t : consumption good price, and i_t : nominal gross interest rate such that $\frac{B_t}{i_t}$ represents expenditures for

purchasing a quantity of bonds that will earn a non-contingent revenue B_t in the next period.

The first order conditions with respect to C_t , B_t and h_t are:

$$a_t \frac{\partial U(C_t, C_{t-1})}{\partial C_t} + \beta E_t \left[a_{t+1} \frac{\partial U(C_{t+1}, C_t)}{\partial C_{t+1}} \right] = P_t \lambda_t \tag{5}$$

$$\beta E_t \left[\frac{\lambda_{t+1}}{\lambda_t} \right] i_t = 1 \tag{6}$$

$$\lambda_t W_t = a_t \tag{7}$$

where λ_t is the Lagrange multiplier associated to the budget constraint, which also represents the welfare's marginal value to the household. Note that a combination of (5) and (6) results in a common aggregate demand or investment-savings curve, according to which current consumption depends on past and future consumptions and the real interest rate $E_t(i_t / \pi_{t+1})$, where $E_t(\pi_{t+1}) = E_t(P_{t+1} / P_t)$ is the expected gross inflation rate.

The production of final good Y_t is performed by a representative retail firm that operates in a perfectly competitive environment and uses the following Dixit-Stiglitz aggregator production function:

$$\left(\int_0^1 Y_t(i)^{\frac{\theta-1}{\theta}} di \right)^{\frac{\theta}{\theta-1}} \geq Y_t \tag{8}$$

where θ measures the elasticity of substitution between the intermediate goods $Y_t(i)$, $i \in [0,1]$. Cost minimization implies the demand schedule for intermediate goods (9) and the aggregate price index (10):

$$Y_t(i) = \left(\frac{P_t(i)}{P_t} \right)^{-\theta} Y_t \tag{9}$$

$$P_t = \left(\int_0^1 P_t(i)^{1-\theta} di \right)^{\frac{1}{1-\theta}} \tag{10}$$

Next, we focus on the description of the production process in the intermediate goods sector. These firms are characterized by monopolistic competition, and we as-

sume a linear constant return to scale production function using labor services $h_t(i)$ supplied by the households and a common stationary technology z_t as inputs:

$$Y_t(i) = z_t h_t(i) \tag{11}$$

$$\ln z_t = \rho_z \ln z_{t-1} + \varepsilon_{zt}, \quad \varepsilon_{zt} \sim N(0, \sigma_z) \tag{12}$$

As previously noted, because the households own these firms, any profits obtained (D_t) are channeled to the former and appear in the budget constraint (4). Solving the cost minimization problem subject to the production function (11), the familiar relation between the nominal wages and nominal marginal cost (NMC_t) results: $W_t = NMC_t z_t$, which can be rewritten in terms of real wage (W_t/P_t) and real marginal cost ($RMC_t = NMC_t/P_t$) as

$$W_t/P_t = RMC_t z_t \tag{13}$$

Because the intermediate producer i is a monopolist, it has the power to independently set the price for his good, $P_t(i)$. In accordance with Calvo (1983), we assume that each period a random share $1-\gamma$ of the firms can optimally set their prices, whereas the remaining γ firms simply index their prices with past inflation, corrected with an indexation coefficient γ_P . The indexation is found to be important in Copaciu, Neagu and Braun-Erdei (2010), because Romanian firms use both forward- and backward-looking information when reviewing prices. Those who can reset the prices choose their optimal price P_t^* (which is equal for all the firms) by maximizing discounted future profits flows:

$$\max E_t \sum_{i=0}^{\infty} \gamma^i \beta^i \frac{\lambda_{t+i}}{\lambda_t} \left[\frac{P_t^*}{P_{t+i}} \left(\frac{P_{t+i-1}}{P_{t-1}} \right)^{\gamma_P} - P_{t+i} RMC_{t+i} \right] Y_{t+i}(i) \tag{14}$$

Solving the above problem by using demand function (9), the solution P_t^* satisfies

$$E_t \sum_{i=0}^{\infty} \gamma^i \beta^i \frac{\lambda_{t+i}}{\lambda_t} Y_{t+i} \left(\frac{P_{t+i-1}}{P_{t-1}} \right)^{-\gamma_P \theta} P_{t+i}^{\theta} \left[\frac{P_t^*}{P_{t+i}} \left(\frac{P_{t+i-1}}{P_{t-1}} \right)^{\gamma_P} + \frac{\theta}{\theta-1} P_{t+i} RMC_{t+i} \right] = 0 \tag{15}$$

Applying the law of large numbers to (10), the aggregate price index (gross) inflation rate can be expressed as:

$$\pi_t^{1-\theta} = (1-\gamma)(\pi_t^*)^{1-\theta} + \gamma(\pi_{t-1}^{\gamma_P})^{1-\theta} \tag{16}$$

where $\pi_t^* = P_t^*/P_{t-1}$. Relation (16) is a standard New Keynesian hybrid Phillips or aggregate supply curve, according to which the current inflation depends on past and future inflation and real marginal cost, as follows from the definition of P_t^* in (15).

Because we assume no other goods than the consumption good C_t , the aggregate resource constraint is simply

$$Y_t = C_t \tag{17}$$

Additionally, the bonds market clearing implies $B_t = B_{t-1} = 0$, whereas the profits are cancelled with the government transfers: $T_t + D_t = 0$.

To close the model, we assume the central bank uses a standard Taylor rule for the nominal interest rate dynamics:

$$\ln \left(\frac{i_t}{i_{SS}} \right) = \rho_i \ln \left(\frac{i_{t-1}}{i_{SS}} \right) + (1-\rho_i) \left[k_{\pi} \ln \left(\frac{\pi_t}{\pi_{SS}} \right) + k_Y \ln \left(\frac{Y_t}{Y_{SS}} \right) \right] + \varepsilon_{it} \tag{18}$$

where SS superscript indicates steady state values, and ε_{it} is a monetary policy shock with $\varepsilon_{it} \sim N(0, \sigma_i)$.

Summarizing the description of the model, the full list of endogenous variables includes $[a_t, C_t, \lambda_t P_t, i_t, \pi_t, W_t/P_t, Y_t, z_t, RMC_t, \pi_t^*]$, for which an equal number of equations is considered: (3), (5)-(7), (12)-(13), (15)-(18). Note that we use $\lambda_t P_t, W_t/P_t, \pi_t$ and π_t^* as final variables following certain convenient substitutions, because the price levels P_t and P_t^* are not identified in the model. The system is perturbed by three shock processes: demand (ε_{at}), supply (ε_{zt}) and monetary policy (ε_{it}). Because we consider three observable variables (as described next), we prefer to model supply shock as a technology innovation, instead of as a markup shock affecting the time-varying elasticity of substitution θ in (8), as in Ireland (2004). Estimation shows both shocks are not separately identified given the three observables. Although the alterna-

tive estimation in which we substitute the technology shock with the mark-up shock provides similar results, we favor the former.

Estimation methodology and data

For the Bayesian estimation of the model, we use three observable variables: the seasonally adjusted real GDP quarterly growth rates ($\Delta \ln(Y_t^{data})$), the quarterly inflation rate of harmonized index of consumer prices (π_t^{data}) and the 3-month money market interest rate (i_t^{data}). The source of the data is the Eurostat, and it covers the 2003Q1 – 2014Q3 period and is plotted in Figure 1 (blue lines). We did not consider observations prior to 2003 because those are not consistent with the more recent readings for Romania (quarterly inflation rates of more than 8% and interest rates of over 40% per annum before 2003). Additionally, because the central bank of Romania switched to inflation targeting in 2005, a shorter sample is less affected by the regime change (a regime switching approach would be more appropriate in this case, a strategy we leave for future analysis), rendering the comparison more consistent. Although the limited sample length allows for more homogenous data, it renders the posterior distributions very sensitive to the addition of new observations and to the results being period-specific (to the analyzed interval) rather than general. Additionally, a proper assessment of forecasting accuracy is difficult to perform in a short sample environment.

The measurement equations link the variables from the data to the model endogenous variables:

$$\Delta \ln(Y_t^{data}) = \ln(\mu \cdot Y_t / Y_{t-1}) \quad (19)$$

$$1 + \pi_t^{data} = \pi_t \quad (20)$$

$$1 + i_t^{data} / 4 = i_t \quad (21)$$

where μ is the steady state gross quarterly growth rate of real GDP. Similar to Smets and Wouters (2005), we use a deterministic trend for real variables, given by μ . However, because the samples are short, we prefer calibrating μ at the historical mean quarterly growth rate of the real GDP series instead of treating it as an unknown parameter.

The calibration of certain coefficients (refer to Table 1) is performed in accordance with the relevant literature, because not all the parameters are identified given the observables in (19)-(21). μ , i^{SS} and π^{SS} are calibrated to their corresponding sample averages. Regarding the inflation rate, usually the steady state is calibrated at the annual target value (currently 2.5% for both countries). However, given the noticeable disinflation trend observed for Romania (and declining annual inflation targets used before 2013), we prefer to use sample averages. Thus, at least for Romania, the model will be accurate in terms of *average* inflation rate fit, but less compatible with the current low inflation environment. The discount factor is set to the inverse real interest rate in steady state, i.e., $\beta = \pi^{SS} / i^{SS}$. Preliminary estimations showed the elasticity of substitution θ is not identified; therefore, we set it to 6, yielding a price mark-up of 20%, as calibrated in Christiano et al. (2011) and estimated in Adolfson et al. (2007).

We adopt a standard Bayesian approach for estimating the remaining parameters, according to which the data likelihood is combined with the information contained in prior distributions to yield parameters' posterior distribution. Marginal prior distributions are identical for the two economies and are standard for the seminal papers; these are presented in the first portion of Table 2. Formally, the risk aversion parameter's prior mean is set to 1.5 (to facilitate estimation, we specify the σ_C coefficient in (2) as $\sigma_C = \sigma^* + 1$, implementing a Gamma prior without bounds on σ^*), whereas the standard deviation is reasonably large. This value is in accordance with the posterior mean of 1.39 obtained in Smets and Wouters (2003) and with the prior mean of 2 in Benchimol (2014) for the Euro area. The Calvo price-rigidity parameter is specified such that, a priori, firms reset prices every three quarters, whereas the indexation coefficient has a mean of 1/3. The inflation and output gap coefficients in the Taylor rule are set to 1.5 and 0.5/4, with higher prior standard deviations compared with the reference papers. The autoregressive parameters are Beta distributed with 0.7 means. For the structural shocks' standard deviations, we consider non-informative Inverse Gamma distribution with 0.5 means.

The estimations were performed using the Dynare toolbox (version 4.4.3) and MATLAB (version

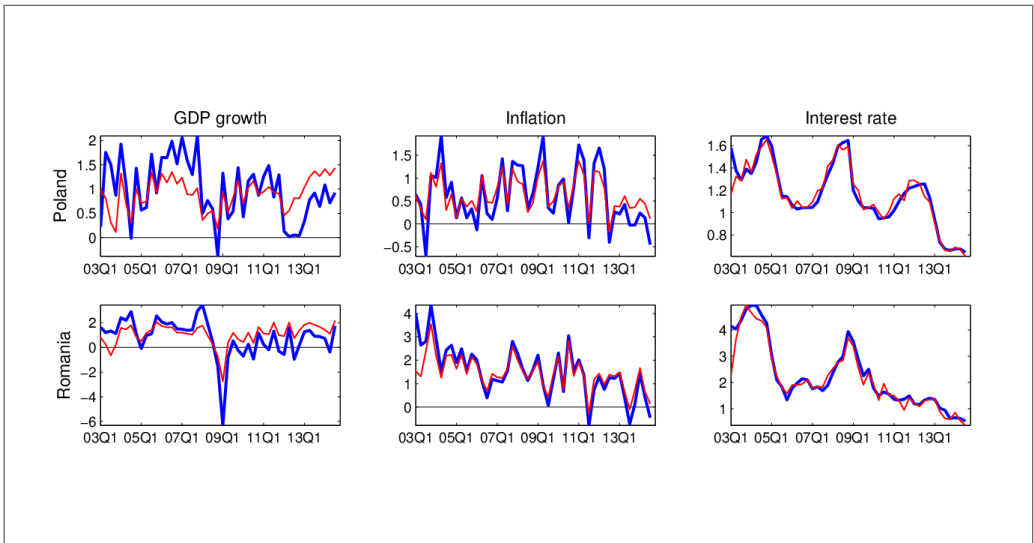


Figure 1. Observed data (blue) and one-sided Kalman filter estimates (red)

Table 1. Calibrated parameters

	Poland		Romania	
	Model calibration	Implied annual value (%)	Model calibration	Implied annual value (%)
μ	exp(0.0098)	3.94	exp(0.0082)	3.29
i^{SS}	exp(0.0117)	4.71	exp(0.0225)	9.10
π^{SS}	exp(0.0062)	2.49	exp(0.0153)	6.17
β	0.9945	–	0.9928	–
θ	6	–	6	–

R2013b). The results discussed in what follows are based on the means of parameters' posterior distributions, obtained from 750000 Metropolis-Hastings simulations with first 250000 burned and acceptance rates of approximately 25% (more precisely, 25.1% for Poland and 23.7% for Romania). The prior-posterior plots are available in an online appendix. In virtually all cases, the shapes of posterior

distributions are sharper than the corresponding prior distributions, suggesting the data contains meaningful information regarding the structural parameters. We do not find evidence of convergence issues, given that the posterior modes and the peaks of posterior distributions coincide; in addition, there are no posterior distributions with multiple modes.

Table 2. Prior distributions and posterior modes

	Prior distributions			Posterior distribution			
	Type	Mean	SD	Poland		Romania	
				Mode	SD	Mode	SD
σ^*_C	Gamma	0.5	0.2	0.4408	0.1780	0.4149	0.1701
b	Beta	0.5	0.2	0.5914	0.0768	0.6509	0.0727
γ	Beta	2/3	0.1	0.7560	0.0481	0.8197	0.0418
γ_P	Beta	1/3	0.1	0.1873	0.0675	0.2143	0.0754
ρ_t	Beta	0.7	0.1	0.9187	0.0172	0.8329	0.0332
k_π	Normal	1.5	0.2	1.4316	0.1824	1.5397	0.1649
k_Y	Normal	0.125	0.1	0.0924	0.0821	0.0889	0.0573
ρ_a	Beta	0.7	0.1	0.7976	0.0626	0.8047	0.0556
ρ_z	Beta	0.7	0.1	0.5601	0.1090	0.6537	0.1143
100 σ_t	Inv.Gamma	0.5	2	0.1272	0.0152	0.3988	0.0451
10 σ_a	Inv.Gamma	0.5	2	0.2811	0.0636	0.7219	0.1706
10 σ_z	Inv.Gamma	0.5	2	0.7425	0.4846	2.5994	1.4236

Note: "SD" indicates standard deviations; $\sigma^*_C = \sigma_C - 1$, to avoid imposing bounds.

Results

The results are discussed with respect to three dimensions: the estimated parameters and implied macroeconomic transmission mechanisms (i), the consistency of the variables simulated within the model with actual data (ii), and the decomposition of GDP growths and inflation rates dynamics into corresponding structural shocks (iii).

Estimated parameters and impulse response functions

First we comment on the posterior means estimates for the structural parameters in the second part of Table 2. Risk aversion, consumption habit and inflation indexation parameters are similar for the two economies. A risk aversion coefficient of approximately 1.4 is consistent with the corresponding value for the Euro area obtained in Benchimol (2014) and, particularly, Smets and Wouters (2003), suggesting similarities in consump-

tion patterns across the emerging and developed European countries. The Calvo rigidity parameters imply the prices are updated every 11 months in Poland and every 16 months in Romania, larger than in the previous estimates for the two countries; Grigoraş (2010) obtains $\gamma = 0.36$ for Romania, whereas Grabek et al. (2011) find $\gamma = 0.69$ for Poland. The differences are likely caused by the different samples and model structures employed. Interest rate smoothing is stronger in Poland; however, in Romania, the central bank is marginally more aggressive with respect to inflation. This behavior is again in opposition to Grigoraş (2010) and Grabek et al. (2011), who obtained $k_\pi = 1.3$ and $k_\pi = 1.8$ respectively. The persistence of both demand and supply shocks are very similar. Given that Romanian data are more volatile (refer to Figure 1), the structural shocks have (much) larger estimated standard deviations.

Impulse response functions (IRF) to one standard deviation in each shock are displayed in Figure 2. Be-

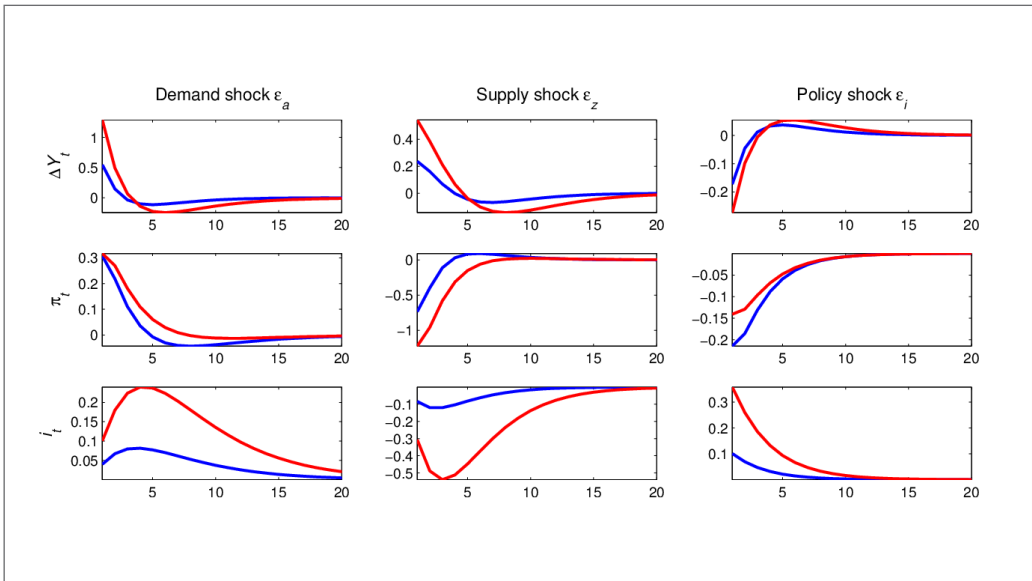


Figure 2. Impulse response functions (Poland: blue, Romania: red)

cause the estimated standard deviations have larger magnitudes for Romania, the corresponding responses are stronger. Otherwise, the economic transmission mechanisms are very similar. Output growth, inflation and interest rate rise following a demand shock, whereas the technology shock produces an increase in output and decreases in both the inflation and the interest rate. Typical monetary policy tightening is a 35 basis points nominal interest rate hike in Romania and a 10 basis points hike in Poland. These policy behaviors produce a similar relative reduction in GDP growth (0.3% and 0.15%, respectively on impact); however, the inflation rate reacts stronger in Poland (-0.2% vis-a-vis -0.15% on impact).

Model’s consistency with actual data

Within the Bayesian estimation of the model, we utilize the Kalman filter to recursively specify and maximize the likelihood function. A one-sided (i.e., using the information set before the current moment) Kalman filter estimates for the observed data are plotted in Figure 1 (red lines). The model becomes remarkably capable of replicating the inflation and interest rates paths, whereas the fit of the GDP growth is less

impressive. Additionally, after running the Kalman smoother (i.e., using all sample information set), the two-sided estimates of the endogenous variables can be extracted. In Figure 3, we present the DSGE model smoothed output deviation from the steady state along the Hodrick-Prescott (HP) filtered output gap (we use a standard value for the smoothing parameter, namely 1600). Although the Kalman filter gaps have larger amplitudes than the HP filtered ones, their shapes are similar for both Poland and Romania, rendering the DSGE model able to replicate certain statistical properties of the data. The Romanian gap is 2-to-3 times larger (in absolute values) than the Polish one, implying more considerable business cycle fluctuations. In addition, the current situation (as of 2014) clearly suggests a slack in economic activity, rendering negative gaps estimated within the DSGE models more appropriate than the closed or positive ones obtained with the Hodrick-Prescott filter.

Because we calibrate the steady states of output growth, the inflation rates and the interest rates to their sample averages, the model perfectly matches the means of these variables, as presented in Table 3. However, the corresponding 90% confidence bands

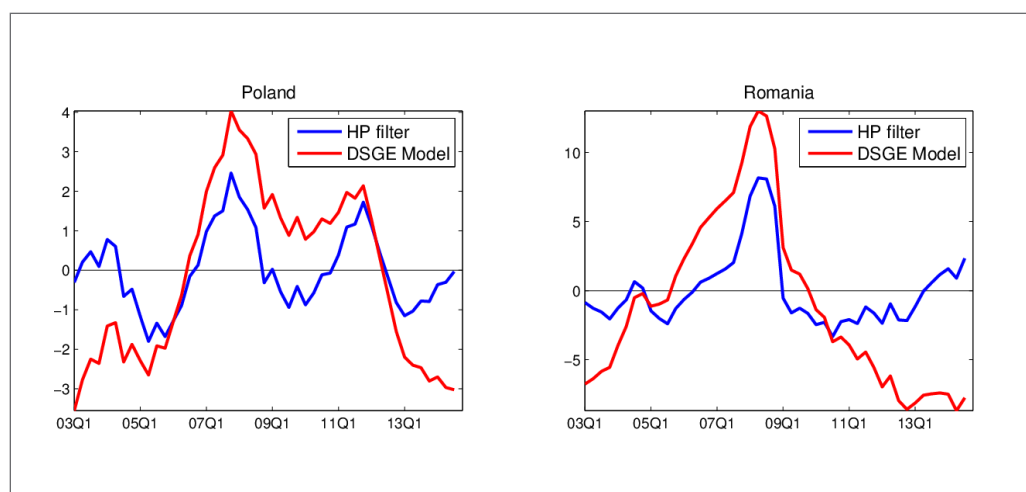


Figure 3. DSGE (Kalman filter) and Hodrick-Prescott (HP) filter output gaps

Table 3. Means and standard deviations matching (90% confidence bands in [])

	Poland				Romania			
	Mean		Standard deviation		Mean		Standard deviation	
	data	model	data	model	data	model	data	model
GDP growth	0.98	0.98 [0.89, 1.07]	0.61	0.72 [0.58, 0.87]	0.82	0.82 [0.56, 1.07]	1.55	1.71 [1.36, 2.10]
Inflation	0.62	0.62 [0.33, 0.91]	0.64	1.01 [0.79, 1.23]	1.53	1.53 [0.75, 2.32]	1.09	1.78 [1.29, 2.18]
Interest rate	1.17	1.17 [0.92, 1.41]	0.28	0.35 [0.20, 0.44]	2.25	2.25 [1.15, 3.36]	1.26	1.46 [0.79, 1.84]

simulated within the model (technically, we simulate 10000 independent series of length 47, i.e., the number of sample observations, from the models' implied distributions and compute corresponding statistics and associated percentiles) provide reasonably certain estimates. The actual standard deviations of Polish GDP growth and interest rate are nearly identical to the model-implied median estimates, whereas the inflation rate low volatility could not be

perfectly replicated within the model. For Romania, the DSGE model nearly perfectly matches the output growth volatility, whereas the standard deviation of interest rate is well inside the 90% confidence band provided by the simulations. Similar to Poland, the model-implied inflation is more volatile than its data counterpart.

To test the model's efficiency in replicating certain inter-relations among the macroeconomic variables

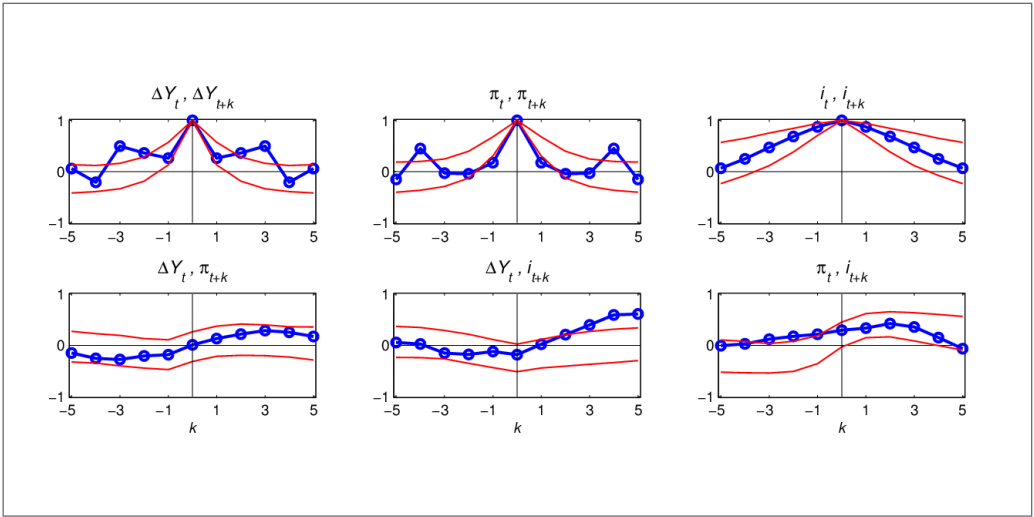


Figure 4. Poland data (blue dotted) and model 90% confidence band (red) cross-correlations

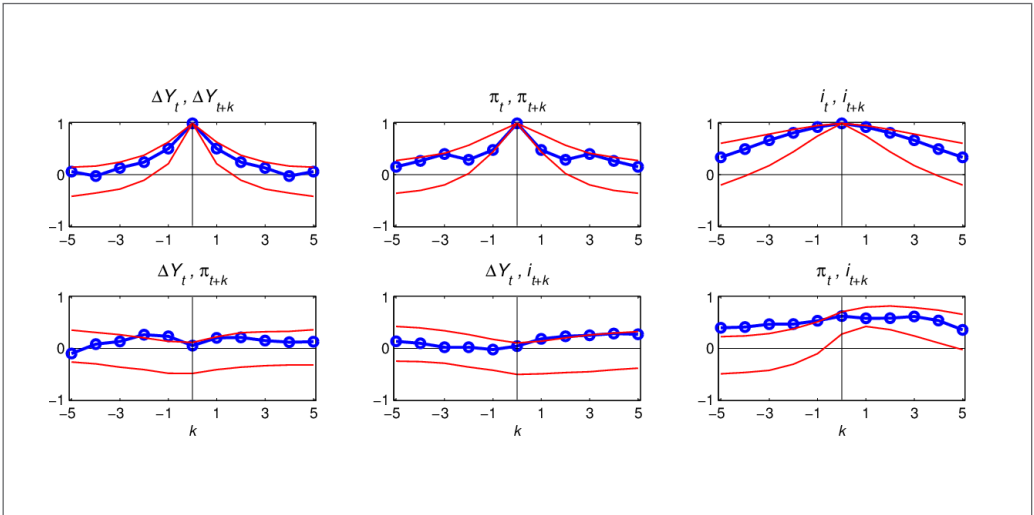


Figure 5. Romania data (blue dotted) and model 90% confidence band (red) cross-correlations

observed in the data, we plot actual cross-correlations (blue dotted lines) and model-implied 90% confidence band cross-correlations (red lines) in Figure 4 for Poland and in Figure 5 for Romania. The DSGE model is able to reproduce the autocorrelation coefficients for

the three variables at all 5 lags/leads for both economies (with a minor exception for Poland). The model has certain difficulties in imitating the relations between GDP growth and interest rate leads in Poland (Figure 4, lower central panel), and the inflation rate

and the lagged interest rate in Romania (Figure 5, lower right panel). Excluding these, the small New Keynesian model yields a reasonable approximation of data cross-correlations.

Structural shocks and historical decomposition

The smoothed shocks processes are shown in Figure 6. As in Christiano et al. (2011), we plot the demand and technology shocks as their autoregressive components (a_t and z_t), instead of the corresponding pure innovations (ε_{at} and ε_{zt}). Given that the data are less volatile in Poland compared to the Romanian time series, the corresponding structural shocks display much smaller fluctuations, particularly the supply/technology innovation. The estimated processes are well correlated across countries (suggesting a high degree of business cycle synchronization; using a business cycle accounting procedure, Kolasa (2013) finds that the new European Union members are becoming more synchronized among themselves, but also with the Euro area) but also with the actual economic events that occurred during the analyzed period. For example, both demand shocks display large positive variations in 2006-2008 during the advanced phases of the economic boom; thereafter, they fall sharply, simultaneously with the onset of the global economic and financial crisis. Because of their persistence, the demand shocks display large negative values in the recent past, indicating consistent slack remains in the two economies. The supply shocks closely follow the observed inflation rates' paths (as noted earlier, the technology shock is equivalent to a time-varying elasticity of substitution or mark-up shock; thus, it has a considerable impact on inflation rate data) but also replicate the economic boom (positive in 2006-2008), the crisis impact (drops in 2008-2009) and attempted subsequent recoveries (alternate signs starting 2009-2010). Moreover, the persistent negative innovations in 2003-2005 for Romania are associated with the large inflation rates during that period. Monetary policy shocks are positive before 2004, because initially the interest rates were high (approximately 6% in Poland and 15% Romania), but drop thereafter, in tandem with the monetary policy easing during the boom period. Romanian monetary policy shock hikes in 2008Q4 are simultaneous with the onset of the late 2000s crisis. Starting in 2009, both interest

rates register successive cuts to combat the crunch, resulting in primarily negative policy shocks.

Regarding the sources of business cycle fluctuations, the long-run variance decompositions of observed variables at posterior mean estimates (refer to Table 4) reveal both GDP growth rates were mainly affected by the consumer preferences shocks, suggesting demand-driven output developments. This result is similar to Caraiani (2008). Monetary policy and demand shocks explain approximately 10% and 15% of the inflation variability each in Poland, whereas the remainder is determined by supply technology shocks. The Romanian inflation rate is nearly entirely (approximately 90%) governed by supply shocks (given the structure of consumer basket, administered prices hikes and weather conditions affected heavily prices' dynamic). The nominal interest rate in Poland was more symmetrically influenced by consumer preferences and technology shocks, whereas the latter was dominant for Romania.

We also perform a historical decomposition procedure in which we decompose (demeaned) observed GDP growth and inflation rates into the contributions of the three shocks and initial conditions (because of the stochastic initialization of the Kalman filter), refer to Figure 7. Both Polish and Romanian outputs are primarily demand-driven: consumption preference shock ε_a explains much of the GDP dynamics during the pre-crisis excessive economic growth but also explains the sharp drops of 2008-2009 (and for the second phase of the crisis in Poland during 2012-2013). Note that the demand shock modelled here as consumption preference innovations (because of the simple model structure) has a much broader interpretation and includes government expenditure and external shocks, which had positive contributions to GDP growth prior to 2008 and primarily negative during and after the crisis. Technology shock ε_z also had pro-cyclical effects in the two economies, but of a smaller magnitude. It is also worth noting the DSGE model's ability to capture the central banks' attempts to mitigate the negative effects of the crisis, as observed from positive effects of the monetary policy shock ε_i in 2009 in Poland and 2009-2010 in Romania.

The breakdown of the shocks for inflation rates historical decomposition is more balanced. Demand shocks were primarily inflationary before 2008-2009 and disinflationary (or even deflationary) once the crisis

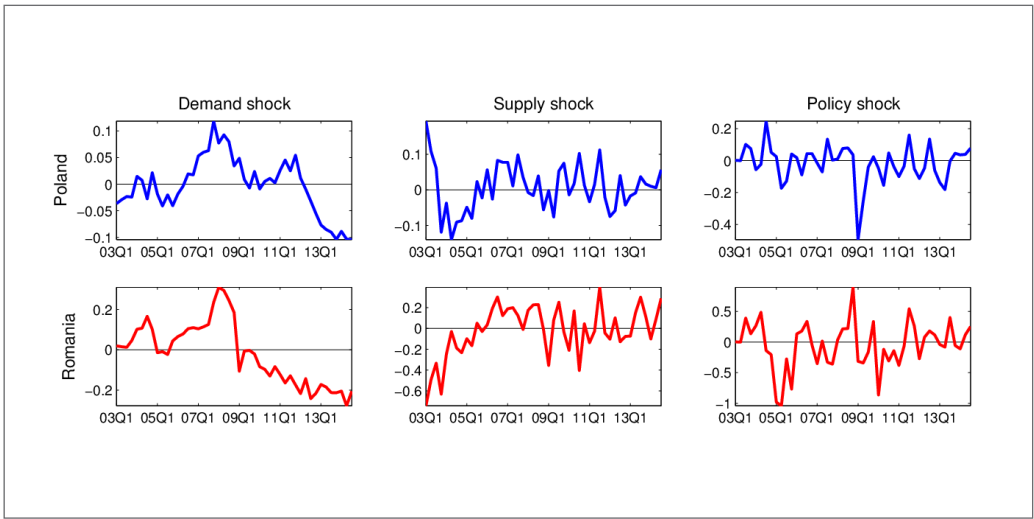


Figure 6. Smoothed shocks

Table 4. Asymptotic variance decomposition, %

	Poland			Romania		
	Policy shock	Demand shock	Supply shock	Policy shock	Demand shock	Supply shock
GDP growth	7	72	21	3	76	21
Inflation	11	16	73	2	7	91
Interest rate	15	35	50	13	20	68

struck. Most erratic short-lived inflation spikes are explained by the technology shocks. These shocks can also be attributed to mark-up, weather conditions or administered prices changes. For example, the large inflation hikes in 2011Q1 in Poland and in 2010Q3 in Romania were caused by the VAT rates increases (interpreted as mark-up shocks); the reduction of Romanian inflation in 2011Q3 was caused by an unexpected large supply of agri-food goods. Additionally, the monetary policy shocks appear to have important contributions for Poland, pushing inflation rates up since 2009 to avoid disinflationary/deflationary pressures. For Romania, the monetary policy shocks are primarily irrelevant. We

interpret this evidence in terms of central banks' degrees of credibility. Because Poland adopted inflation targeting in 1998, the central bank is assessed to achieve a certain level of credibility and was more efficient at shaping the inflation rate, in contrast to the central bank of Romania, which introduced inflation targeting in 2005.

Although we are aware that the results obtained are highly model-specific and that allowing for a richer structure may substantially alter the conclusions achieved in this analysis, the ability of the small DSGE models to generate data-consistent variables and to explain the business cycle fluctuations using the broad definitions of the three shocks are noteworthy.

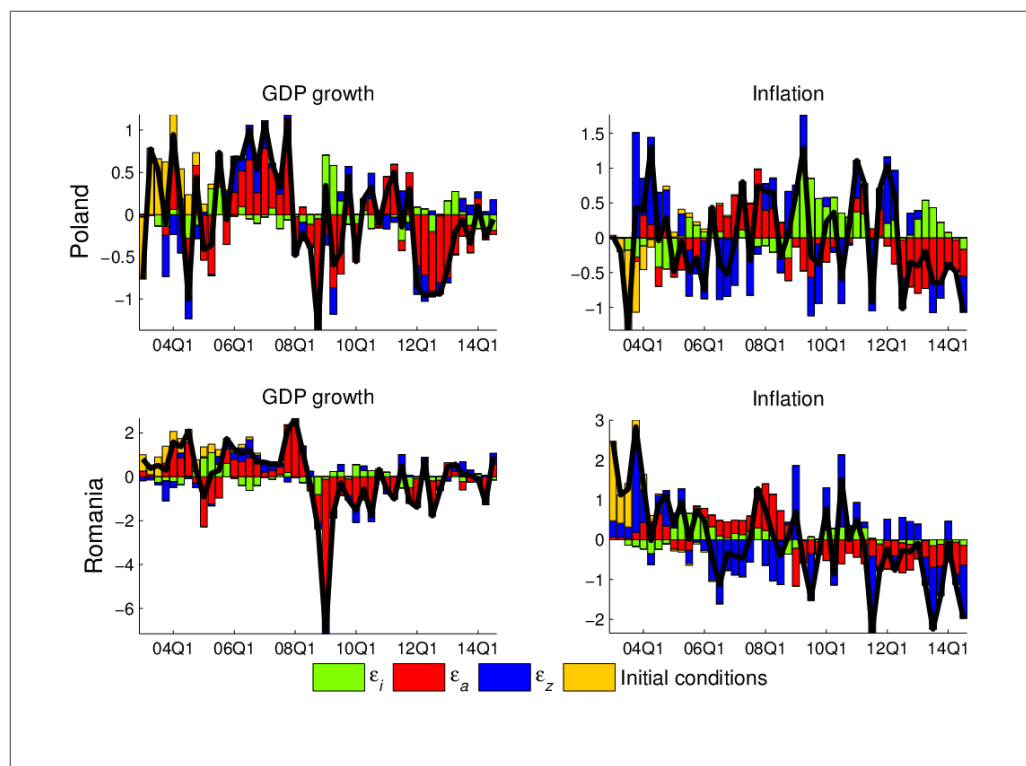


Figure 7. Historical decomposition (demeaned data)

Conclusions

In this paper, we employ and estimate small-scale DSGE models for Poland and Romania. These models include certain standard New Keynesian ingredients, such as price stickiness, price indexation, interest rate smoothing, and consumption habit formation. The system is disturbed by three structural shocks: consumption preference (demand), technology (supply) and monetary policy innovations. In addition to increased tractability, a simple model structure is more appropriate to estimate when samples are short. The calibration scheme and marginal prior distributions are identical, whereas the observed database covers output growth, the consumer prices inflation rate and the 3-month nominal interest rate for 2003Q1 – 2014Q3 period.

The estimated structural coefficients are very similar across the two economies; however, the shocks' standard deviations are larger for Romania, implying impulse response functions with similar shapes

but with slightly different magnitudes. Moreover, the DSGE model-implied deviations of output from the steady state display comparable shapes and developments, suggesting similar business cycle dynamics.

The models were proven to fit the data very well despite their rudimentary structure. Thus, the data moments are matched properly for Poland. Although similar performance was not achieved for Romania, the observed standard deviations of the three series are well inside or close to the model-associated confidence regions. The sample auto- and cross-correlations are situated, with minor exceptions, within the model confidence bands of the simulated corresponding statistics.

The shock processes emphasize the similarities of the two economies and capture certain actual events that occurred during the analyzed period (such as the late 2000s economic crisis) very well. Variance and historical decompositions indicate the output

growth was largely driven by consumption preference shocks, with systematic positive contributions before 2008-2009 and negative thereafter. Additionally, the models are able to replicate the central banks' responses following the crisis strike. Monetary policy easing is interpreted as positive effects of the interest rate rule shocks to the output dynamics and upward inflationary pressures in both economies, but with more pronounced effects in Poland. The inflation rates were driven by a more balanced combination of the three shocks in Poland, and primarily by supply shocks in Romania. Thus, the same model environment was proven to be useful in displaying both common and country-specific features.

For future research, we plan to extend the model with additional channels (such as capital accumulation or to add the external dimension, given that both Romania and Poland are small open economies) to broaden the interpretation of business cycles' driving forces, and to supplement the evaluation toolkit with a forecasting performance procedure.

References

- Adolfson, M., Laséen, S., Christiano, L., Trabandt, M., & Walentin, K. (2013). Ramses II – Model Description (Occasional Paper Series No. 12). Sveriges Riksbank.
- Adolfson, M., Laséen, S., Lindé, J., & Villani, M. (2007). Bayesian estimation of an open economy DSGE model with incomplete pass-through. *Journal of International Economics*, 72 (2), 481 – 511.
- An, S., & Schorfheide, F. (2007). Bayesian Analysis of DSGE Models. *Econometric Reviews*, 26 (2-4), 113-172.
- Benchimol, J. (2014). Risk aversion in the Eurozone. *Research in Economics*, 68 (1), 39-56.
- Bernanke, B., Gertler, M., & Gilchrist, S. (1999). The financial accelerator in a quantitative business cycle framework. In J. B. Taylor & M. Woodford (Eds.), *Handbook of Macroeconomics* (Vol. 1) (pp. 1341-1393). Amsterdam: North-Holland.
- Calvo, G. (1983). Staggered prices in a utility-maximizing framework. *Journal of Monetary Economics*, 12 (3), 383-398.
- Caraiani, P. (2007). An Estimated New Keynesian Model for Romania. *Romanian Journal of Economic Forecasting*, 4 (4), 114-123.
- Caraiani, P. (2008). An Analysis Of Domestic And External Shocks On Romanian Economy Using A DSGE Model. *Romanian Journal of Economic Forecasting*, 5 (3), 100-114.
- Caraiani, P. (2013). Comparing monetary policy rules in CEE economies: A Bayesian approach. *Economic Modelling*, 32, 233-246.
- Castillo, P., Montoro, C. & Tuesta, V. (2006). An Estimated Stochastic General Equilibrium Model with Partial Dollarization: a Bayesian Approach (Working Paper No. 381). Central Bank of Chile.
- Christiano, L., Eichenbaum, M., & Evans, C. (2005). Nominal Rigidities and the Dynamic Effects of a Shock to Monetary Policy. *Journal of Political Economy*, 113 (1), 1-45.
- Christiano, L., Trabandt, M., & Walentin, K. (2011). Introducing financial frictions and unemployment into a small open economy model. *Journal of Economic Dynamics & Control*, 35 (12), 1999-2041.
- Clarida, R., Gali J., & Gertler M. (1999). The Science of Monetary Policy: A New Keynesian Perspective. *Journal of Economic Literature*, 37 (4), 1661-1707.
- Copaciu, M., Neagu, F., & Braun-Erdei, H. (2010). Survey evidence on price-setting patterns of Romanian firms. *Managerial and Decision Economics*, 31 (2-3), 235-247.
- DeJong, D., & Dave, C. (2007). *Structural Macroeconomics*. Princeton, NJ: Princeton University Press.
- Erceg, C., Henderson, D., Levin, A. (2000). Optimal monetary policy with staggered wage and price contracts. *Journal of Monetary Economics*, 46 (2), 281-313.
- Gomes, S., Jacquinot, P., & Pisani, M. (2010). The EAGLE. A model for policy analysis of macroeconomic interdependence in the euro area. European Central Bank Working Paper Series, 1195.
- Grabek, G., Kłos, B., & Koloch, G. (2011). SOE PL 2009 – An Estimated Dynamic Stochastic General Equilibrium Model for Policy Analysis And Forecasting (Working Paper No. 83). National Bank of Poland.
- Grigoraș, V. (2010). Estimation of an open economy DSGE model for Romania. Do nominal and real frictions matter? Advances in Economic and Financial Research DOFIN (Working Paper No. 47). Doctoral School of Finance and Banking. The Bucharest Academy of Economic Studies.

- Ireland, P. (2004). Technology Shocks in the New Keynesian Model. *The Review of Economics and Statistics*, 86 (4), 923-936.
- Kolasa, M. (2013). Business cycles in EU new member states: How and why are they different? (Working Paper No. 156). National Bank of Poland.
- Kydland, F. & Prescott, E. (1982). Time to Build and Aggregate Fluctuations. *Econometrica*, 50 (6), 1345-1370.
- Lucas, R. Jr. (1976). Econometric Policy Evaluation: A Critique. *The Phillips Curve and Labor Markets*, Carnegie-Rochester Conference Series on Public Policy, 1, 19-46.
- Mankiw, G. & Romer, D. (1991). *New Keynesian Economics (Vol. 1), Imperfect competition and sticky prices*. Cambridge, MA: MIT Press.
- Mortensen, D., Pissarides, C. (1994). Job creation and job destruction in the theory of unemployment. *Review of Economic Studies*, 61 (3), 397-415.
- Pagan, A. (2003). Report on modelling and forecasting at the Bank of England. *Bank of England Quarterly Bulletin*, 43 (1), 60-88.
- Rotemberg, J. (1992). Sticky Prices in the United States. *Journal of Political Economy*, 90, 1187-1211.
- Smets, F., & Wouters, R. (2003). An estimated dynamic stochastic general equilibrium model of the euro area. *Journal of the European Economic Association*, 1 (5), 1123-1175.
- Smets, F., & Wouters, R. (2005). Comparing shocks and frictions in US and euro area business cycle fluctuations: a Bayesian DSGE approach. *Journal of Applied Econometrics*, 20 (2), 161-183.
- Taylor, J. B. (1993). Discretion versus Policy Rules in Practice. *Carnegie-Rochester Conference Series on Public Policy*, 39 (1), 195-214.

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Online appendix

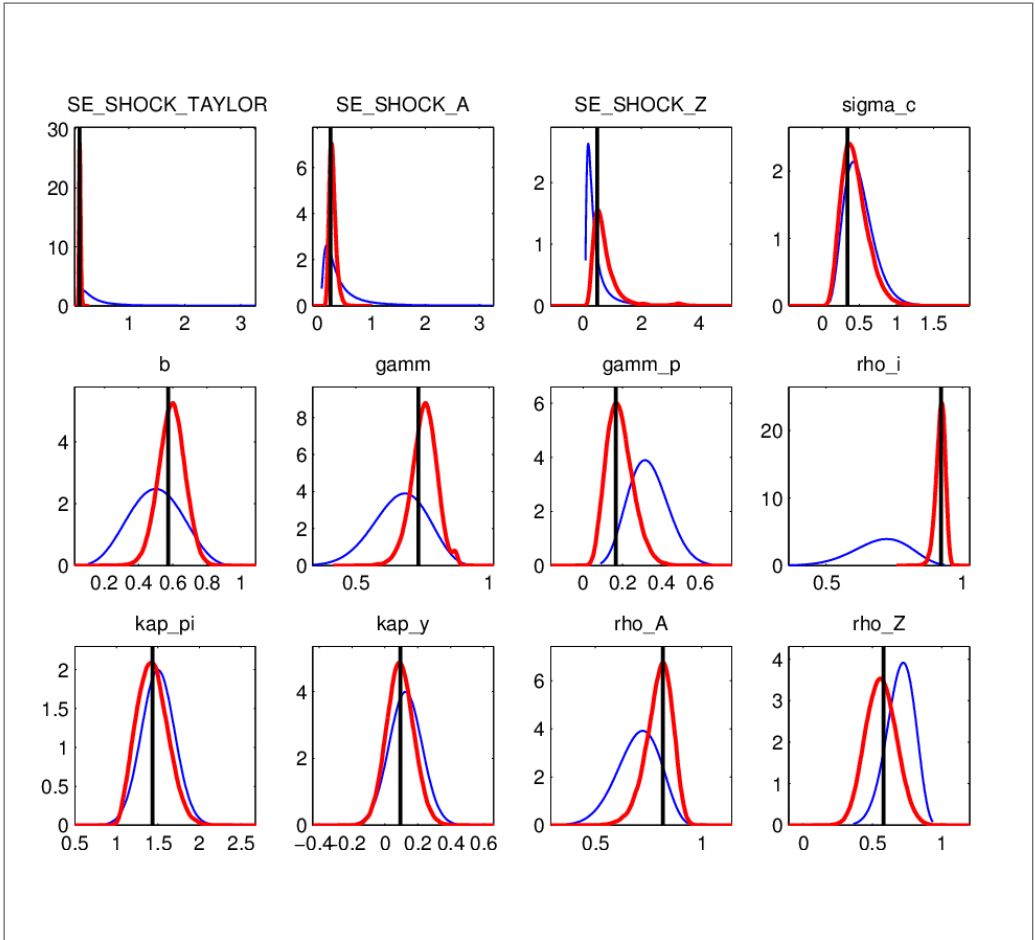


Figure A1. Poland: prior (blue) and posterior (red) distributions, modes (black)

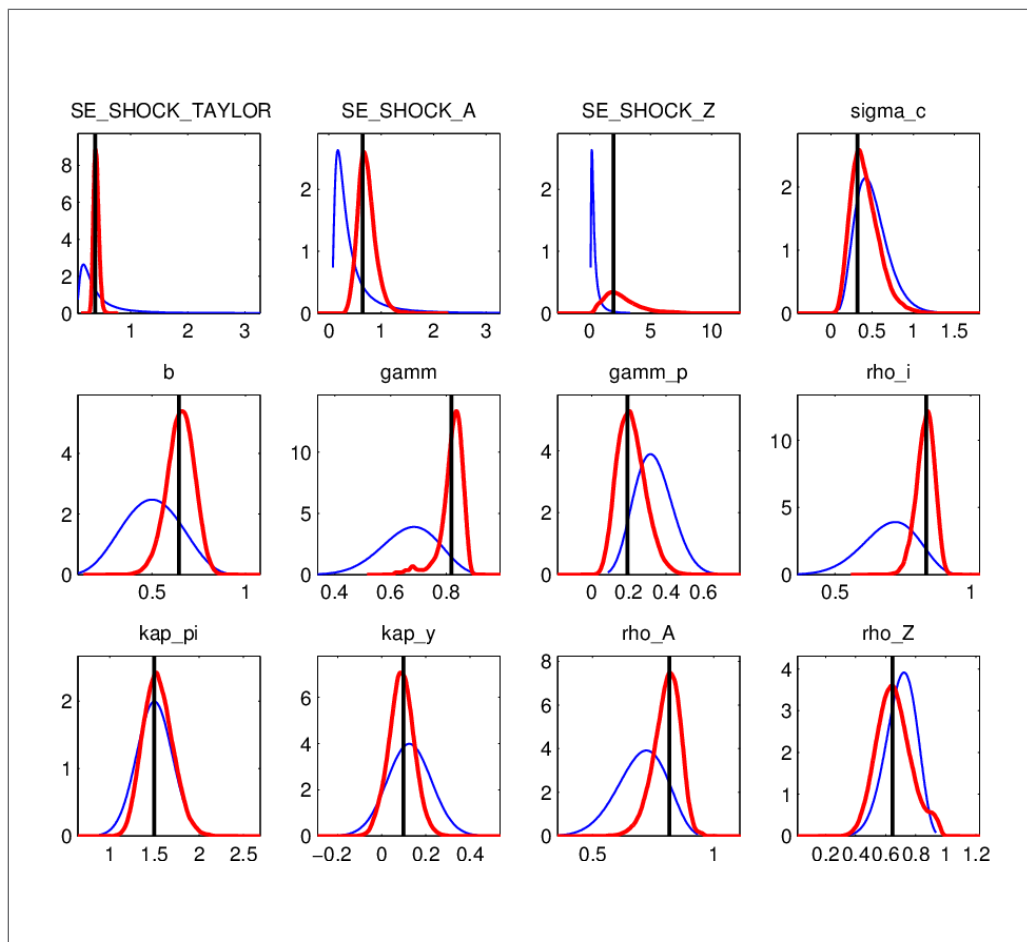


Figure A2. Romania: prior (blue) and posterior (red) distributions, modes (black)

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