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ABSTRACT

We analyse the short-term dynamics of Polish economy with a prominent state-dependent pricing mechanism of Dotsey, King and Wolman (1999). We compare macroeconomic evidence of price rigidity in a small-scale DSGE model with a state-dependent Phillips curve (SDPC) derived by Bakhshi, Khan and Rudolf (2007) to a benchmark model including hybrid New-Keynesian Phillips Curve (NHPC) of Gali and Gertler (1999). To analyse monetary policy transmission mechanism we estimate both models with Bayesian techniques and focus on the comparison of distribution of price vintages, a degree of price stickiness, values of parameters in Phillips curve equations, and impulse responses to macroeconomic shocks. The estimated state-dependent pricing model generates a median duration of prices about 4 quarters compared to 8 quarters in a time-dependent model. In the state-dependent pricing model it takes more time to dampen inflation dynamics after a monetary policy relative to a time-dependent counterpart. The menu cost model is also able to identify higher variance of technology shocks, and higher persistence in preference shocks, while the dynamics of the impulse responses in time- and state-dependent pricing models are hard to distinguish.

JEL Classification: E31

Keywords: state-dependent pricing, inflation, menu costs, monetary policy, Polish economy

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Inflation in Poland under state-dependent pricing

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Abstract

Menu costs offer an appealing explanation of price rigidity in a New-Keynesian DSGE framework, which is provided at the cost of model tractability. In the paper we attempt to analyse the short-term dynamics of Polish economy with a prominent state-dependent pricing mechanism of Dotsey, King and Wolman (1999). We compare macroeconomic evidence of price rigidity in a small-scale DSGE model with a state-dependent Phillips curve (SDPC) derived by Bakhshi, Khan and Rudolf (2007) to a benchmark model including hybrid New-Keynesian Phillips Curve (NHPC) of Gali and Gertler (1999). To replicate a short-term persistence in inflation and output both models include other sources of economic inertia (i.e. habit persistence in consumption, interest rate smoothing in a Taylor-type rule). To analyse monetary policy transmission mechanism we estimate both models with Bayesian techniques and focus on the comparison of distribution of price vintages, a degree of price stickiness, values of parameters in Phillips curve equations, and impulse responses to macroeconomic shocks.

The estimated state-dependent pricing model generates a median duration of prices about 4 quarters compared to 8 quarters in a time-dependent model. Quite surprisingly, in the state-dependent pricing model it takes more time to dampen inflation dynamics after a monetary policy relative to a time-dependent counterpart. The menu cost model is also able to identify higher variance of technology shocks, and higher persistence in preference shocks. Despite those significant differences the dynamics of the estimated impulse responses in time- and state-dependent pricing models are hard to distinguish.

Introduction

Dynamic stochastic general equilibrium (DSGE) models are prominent tools for analyzing short-term deviations of the economy from its steady state in a New-Keynesian style analysis (see Woodford 2003, Gali 2008). Contrary to standard old-fashioned structural multi-equation models, DSGE framework is argued to be resistant to Lucas critique because it relies heavily on rational behaviours of forward-looking microeconomic agents. It is also a dominating framework of monetary policy analysis with structural equations between macroeconomic variables being derived directly from the rules of entrepreneurs maximizing profits, households maximizing utility function, and central bankers following the monetary policy objectives.

One of the milestones in a New-Keynesian modelling is the Calvo (1983) time-dependent stochastic price stickiness. In this particular pricing mechanism firms in every period face a constant probability of resetting a price. This assumption is very helpful in deriving a short-term relation between inflation and output gap known as New-Keynesian Phillips Curve (NKPC). This short-term inflation-output trade-off induces significant real effects of monetary policy shocks in DSGE model.

Although firms in the Calvo price setting take into consideration the probability of price stickiness, the timing of these decisions is random and exogenous. The Calvo pricing mechanism results in a constant average frequency of price adjustment across firms and time. On a microeconomic level it does not depend on a passage of time from the last price adjustment. On a macro scale it is independent of any variable describing current state of the economy (*e.g.* a long-run inflation). It is argued that the Calvo pricing alone is not able to create enough inflation and output persistence widely observed in the data. Thus in empirical research it is enhanced by 'rule-of-thumb' consumers (e.g. Gali, Gertler, 1999) or indexation (Christiano, Eichenbaum, and Evans 2005).

State-dependent pricing models (reviewed empirically in Klenow, Kryvtsov, 2008) – main rivals of time-dependent approach – introduce menu costs to price-setting decisions of representative firm. In opposition to Calvo lottery, price adjustment costs introduce a rationale for the frequency of price changes in a style which is called a 'selection effect'. According to this concept the firms, which are resetting their prices in current period, are those with prices farthest away from their optimal level. Thus in a state-dependent pricing approach the frequency of price changes becomes endogenous, dependent on the shocks and current prices of all representative firms. It considerably complicates the derivation of the Phillips curve. In some of the state-dependent models selection effect may also have some consequences to the degree of money non-neutrality (e.g. Golosov, Lucas, 2007).

Only few papers present fully fledged DSGE models with state-dependent pricing mechanisms which are applied to macroeconomic data. Authors mostly offer only numerical solution or they usually calibrate the parameters to meet the needs of a microeconomic evidence. As the evidence from micro-level dataset is scarce in Poland we focus on the papers that derive close form solution for aggregate inflation equation. In this respect Gertler and Leahy (2008) locally approximate a state-dependent Phillips curve (SDPC) around a zero inflation steady state. Its structure resembles a traditional forward-looking Phillips with the exception that its coefficients are dependent on other 'deep' parameters of the economy. SDPC of Gertler, Leahy (2008) in a DSGE framework calibrated to fit the Klenow-Kryvtsov (2008) micro-price dataset set exhibits the real effects of the monetary policy which are similar to the one obtained in conventional time-dependent pricing models. Bakhshi, Khan and Rudolf (2007) derive state-dependent Phillips curve of Dotsey, King and Wolman (1999, henceforth: DKW), around a positive steady-state inflation in which inflation depends on current and future marginal costs, and both expected and lagged inflation. It also includes state-dependent terms connected to future, and past distribution of firms vintages, defined as a group of firms with the same price. The authors after a series of exercises on a simulated data claim that SDPC offers empirical explanation of intrinsic persistence. This paper of Bakhsi *et al.* (2007), although it has not been challenged with any empirical data, is closely related to our paper.

The model of state dependent pricing

We build on the DKW model, where infinitely many firms indexed by $i \in [0; 1]$ sell differentiated final goods at prices set optimally in the monopolistically competitive market. The only objective of a representative firm is to maximize the sum of discounted expected profits under rational expectations. While resetting price each firm faces different stochastic menu cost which discourage her from changing the price in every period. Menu costs are calculated in the units of labour costs i.e. they are interpreted in terms of the amount of labour necessary to accomplish all activities connected to price changes. We treat those costs as if they were independent (across time and firms) realizations of a continuous random variable, $\xi_{i,t}$, distributed on $[0, B]$ interval. Economically,

B, is an upper bound on menu costs representing potential (opportunity) costs of price adjustment. In each period decisions on resetting the price are conditional on potential real profits from changing the price and random menu costs. In period t only a fraction of firms, drawing menu costs below a given threshold, sets a new price. If a representative firm decides to change the price in period t, it sets new optimal price, P_t^*, that maximize its profits. The firms with relatively high menu costs, at the end of period t, leave their prices unaltered. In a consequence of price-setting decision, otherwise homogenous firms are assigned to different groups (so called 'price vintages') of firms that last changed their price j ($j = 0,1, \dots J$) periods ago, $P_t(i(j)) = P_{t-j}^*$. In period t firms from vintage j solve the similar dynamic optimization problem, and change their price whenever:

$$v_{0,t} - v_{j,t} > \xi_{i,t} W_t, \tag{1}$$

where $v_{0,t}$ and $v_{j,t}$ are sums of discounted expected current and future profits conditional on events of 'setting new price' ($P_t(i) = P_t^*$) and 'no price change' ($P_t(i) = P_{t-j}^*$), respectively, and W_t denotes economy-wide real wage rate in period t(cf. Appendix C DKW – firm pricing decisions).

Let, at the beginning of period t, $\omega_{j-1,t-1}$, $j = 1,2,..J$ denote a fraction of all firms belonging to a price vintage j. In a vintage j in period t a portion of firms, $\alpha_{j,t}$, with relatively low stochastic menu costs, decides to reset its price to P_t^*. Next period they move to the first price vintage ($j = 1$). The rest of the firms from price vintages j, which did not change the price, migrate to a price vintage $j + 1$. In the last vintage J the benefits for all firms from resetting the price is bigger than the upper bound of menu cost B, so all of them reset the price and migrate to the first vintage.[1] The dynamic relations between $\omega_{j,t}$, and $\alpha_{j,t}$ are described by identities (2) and (3):

$$\omega_{j,t} = (1 - \alpha_{j,t})\omega_{j-1,t-1}, \quad j = 1,2 \dots, J - 1 \tag{2}$$

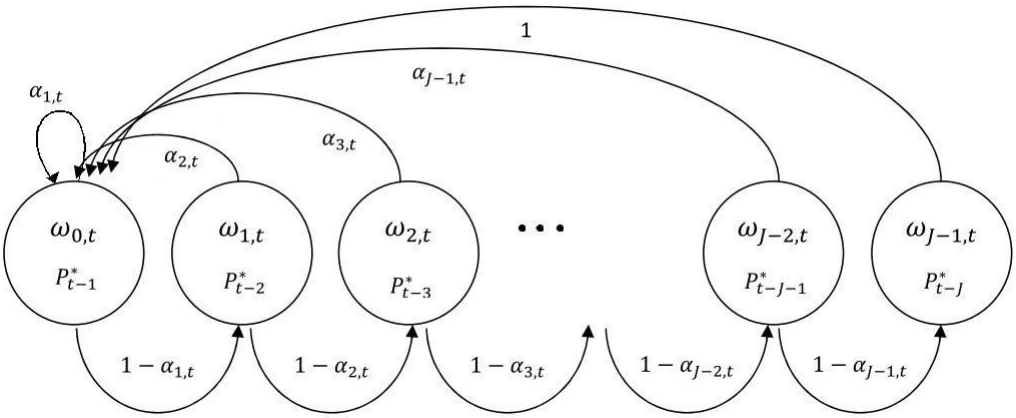
$$\omega_{0,t} = \sum_{j=1}^{J} \alpha_{j,t}\omega_{j-1,t-1}. \tag{3}$$

A law of motion between price vintages given by equations (2) and (3) determines changes in a distribution of firms between price vintages (see Figure 1). DKW pricing mechanism may be described by a time non-homogenous discrete Markov chain with states $1, \dots, J$ and with transition matrices, M_t:

$$M_t = \begin{bmatrix} \alpha_{1,t} & 1-\alpha_{1,t} & 0 & \dots & 0 \\ \alpha_{2,t} & 0 & 1-\alpha_{2,t} & \dots & 0 \\ \dots & \dots & \dots & \dots & \dots \\ \alpha_{J-1,t} & 0 & 0 & \dots & 1-\alpha_{J-1,t} \\ 1 & 0 & 0 & \dots & 0 \end{bmatrix}, \mathrm{t} = 1,2, \dots \tag{4}$$

[1] Due to strictly positive steady-state inflation $\pi^{ss} > 0$ and bounded support of menu cost distribution, there exists a finite number of price vintages, J. The number of non-empty vintages J is a result of firms optimization decisions. It depends on the current shocks and such model parameters: steady state inflation Π, price elasticity of demand and distribution of menu cost G.

Fig. 1. DKW mechanism of migration between price vintages.



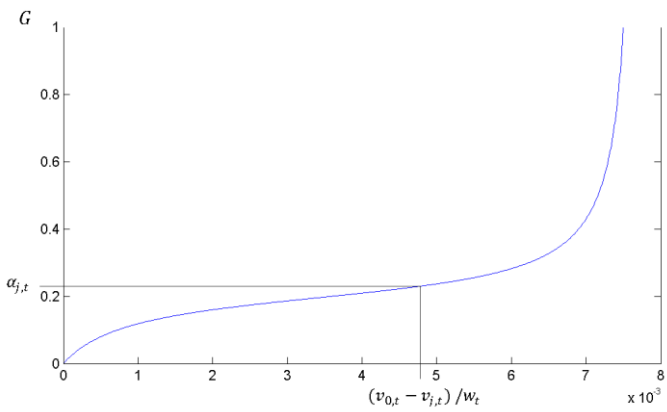
As there are infinitely many firms and menu cost is drawn from a continuous distribution, in every price vintage j there exists a `marginal' firm, $i^*(j)$, with profits of changing the price equal to menu cost:

$$v_{0,t} - v_{j,t} = \xi_{i^*(j),t} W_t \tag{5}$$

Identity (5) with a distribution function G (see Figure 2) defines a fraction of firms that in period t set common optimal price, P^*_t:

$$\alpha_{j,t} = G((v_{0,t} - v_{j,t})/W_t) = G(\xi_{i^*(j),t}) \tag{6}$$

Fig. 2. Menu cost distribution function, and the fraction of firms resetting the price in period t.



Source: Distribution of menu cost according to Dotsey, King, and Wolman (1999) which is a family of tangent-like functions: $G(x) = c_1 + c_2 \cdot tan(c_3 \cdot x + c_4)$, where $c_1 = 0{,}1964, c_2 = 0{,}0625$, and $c_3 = (\arctan((1-c_1)/c_2) - \arctan(-c_1/c_2))/B$, $c_4 = \arctan(-c_1/c_2)$

From the equations (2), (3) and (5) the values of probabilities of resetting the price $\alpha_{1,t}, \alpha_{2,t}, \dots, \alpha_{J,t}$ are not decreasing with j, and the distribution of firms across price vintages $\omega_{0,t}, \omega_{1,t}, \dots, \omega_{J-1,t}$ is non-increasing. Due to a selection effect the later the firm resets its price the bigger the price change is. This phenomenon is not present in exogenous random pricing mechanism of Calvo (1983),

in which the number of firms in consecutive $j = 1,2, \dots$ fractions goes down at the geometric rate and the adjustment probabilities are constant.

Estimation of State Dependent Phillips Curve

Here, we focus on the estimation of Phillips curve derived from DKW pricing mechanism around non-zero steady-state inflation by Bakhshi, Khan, and Rudolf (2007), so called State-Dependent Phillips Curve (SDPC). We introduce one modification to original SDPC and incorporate external habit persistence in consumption (see Abel 1990), which results in a lagged term of output gap:[2]

$$\begin{aligned}\tilde{\pi}_t = \; & E_t \sum_{j=1}^{J-1} \delta_j' \tilde{\pi}_{t+j} + \sum_{j=1}^{\infty} \mu'_j \tilde{\pi}_{t-j} + E_t \sum_{j=0}^{J-1} \hat{\psi}_j' x_{t+j} + \vartheta x_{t-1} \\ & + \sum_{j=0}^{\infty} \eta_j \hat{\Omega}_{t-j} + E_t \sum_{j=0}^{J-1} \gamma_j (\tilde{\omega}_{j,t+j} - \tilde{\omega}_{0,t}) + v_t^{\pi},\end{aligned} \tag{7}$$

where $\tilde{\pi}_t$ is a deviation of inflation from its non-zero steady-state level, x_t is an ouput gap, $\tilde{\omega}_{j,t} = \ln \omega_{jt} - \ln \omega_j$, and $\hat{\Omega}_t = \sum_{j=0}^{J-1} \frac{1}{\omega_0} \Pi^{j(\varepsilon-1)} \omega_j \tilde{\omega}_{j,t}$ are, respectively, expected and past state-dependent unobserved terms.

$$\delta_j' = \frac{1}{\mu_0} \sum_{i=j}^{J-1} (\varepsilon \rho_i - (1-\varepsilon)\delta_i), \text{ for } j = 1, \dots, J-1, \rho_j = \frac{\beta^j \omega_j \Pi^{j\varepsilon}}{\sum_{i=0}^{J-1} \beta^i \omega_i \Pi^{i\varepsilon}}, \delta_j = \frac{\beta^j \omega_j \Pi^{j(\varepsilon-1)}}{\sum_{i=0}^{J-1} \beta^i \omega_i \Pi^{i(\varepsilon-1)}},$$

$$\mu_j = \frac{1}{\omega_0} \sum_{i=j+1}^{J-1} \omega_i \Pi^{i(\varepsilon-1)}, \hat{\psi}'_j = \frac{\hat{\psi}_j}{\mu_0}, \vartheta = \frac{\kappa_2 \rho_0}{\mu_0}, \text{ for } j = 0,1, \dots, J-1$$

$$\hat{\psi}_j = \begin{cases} \kappa_1 \rho_j + \rho_j - \delta_j + \kappa_2 \rho_{j+1} & j = 0,1, \dots, J-2 \\ \kappa_1 \rho_j + \rho_j - \delta_j & j = J-1 \end{cases}, \kappa_1 = \frac{\phi + \alpha + (1-\alpha)\sigma}{1-\alpha}, \kappa_2 = h(\sigma - 1),$$

and matrix formulas on μ'_j, η_j for $j = 1,2, \dots$ are given in Appendix B to Bakhshi, *et al.* (2006).

To compare DSGE models with time- and state-dependent pricing mechanisms we perform a Bayesian estimation of the main parameters. We also compare mean values of parameters of SDPC and NHPC, and the distribution of price vintages ω_j in a steady-state. Then from the posterior distribution of parameters we generate 'empirically-based' impulse response function. They rely on a disturbance of the economy by unanticipated impulse shock estimated and identified on the same dataset within each of the models. Despite the complications of DKW model we perform the estimation in Dynare package (see Adjemain *et. al*. 2011).

We estimate DSGE models for Polish economy on quarterly data from the period 1997-2010. Inflation is measured with CPI (q-to-q), interest rates is short-term interest rate. Because a disinflation process is dominating long-term component in the first 5 years of the sample we perform the estimation on HP-demeaned inflation and interest rates.[3] In the same fashion the output gap was calculated as percentage deviations form HP-trend which is a standard approach in

[2] For the general overview of the SDPC derivation and the structure of DSGE economy we refer the interested reader to the Appendix C.

[3] We have also performed the estimation on non-demeaned inflation and interest rate (only successful for NHPC) with quantitatively different results (incredibly persistent technology shocks).

determining steady-state level of production in the most of DSGE studies for Poland. The figures of the original data and their transformations are depicted in Appendix A.

We start from necessary modifications that facilitate the Bayesian estimation of SDPC derived in (16). Unlike the NHPC, it includes additional $(J-1)$ leads of inflation and output gap, and infinite number of lagged inflation terms. Moreover, there are unobserved characteristics of distributions of firms across price vintages that depend on the realisations of time-heterogeneous Markov process. Firstly, we examine the robustness of SDPC to changes in: distributions of menu costs, steady-state markup and inflation. The exercise shows that a number of price vintages (J) and the fractions of firms in consecutive price vintages are mostly sensitive to changes in markup and steady-state inflation. Secondly, fraction of future price vintages $\omega_{j,t+k}$, price vintages in steady state ω_j, and their weighted absolute deviations $\widehat{\Omega}_{t-k}$ are all unobserved. These components are solutions to dynamic optimization problem (10). From a technical and practical perspective we omit the impact of expected $\omega_{0,t}, \omega_{1,t}, \dots, \omega_{J-1,t}$. Their numerical values are defined by a time-non-homogeneous Markov chain with transition matrices given by (9) and depend on expected opportunity profits from resetting the price, $v_{0,t} - v_{j,t}$. To calculate these quantities for each period in Bayesian estimation with Markov Chain Monte Carlo (MCMC) would be a matter of further numerical complication. Because $\omega_{0,t}, \omega_{1,t}, \dots, \omega_{J-1,t}$ are not observable, their joined estimation with posterior simulation would also introduce additional identification problems.

The other unobservables – steady-state fractions of firms in J price vintages $\omega_0, \omega_1, \dots, \omega_{J-1}$ are solutions to a time-homogeneous Markov chain. Moreover, they are directly related to structural parameters $(\delta_j', \widehat{\psi}'_j, \mu_j')$. To employ these values in posterior MCMC estimation we construct an exponential multinomial with interaction terms that interpolates ω_j reasonably well. The grid is built on a joined domain of a markup $m \equiv \frac{\varepsilon}{\varepsilon-1} \in [1.1; 1.35]$ and an upper bound of menu cost distribution function $B \in [0.0075; 0.05]$ with a steady-state inflation $\pi^{ss} = 4{,}3\%$ p.a. and maximal number of price vintages $J = 10$.

In effect we have estimated a standard three-equation DSGE model with SDPC:

$$\tilde{\pi}_t = E_t \sum_{j=1}^{9} \delta_j' \tilde{\pi}_{t+j} + \vartheta x_{t-1} + E_t \sum_{j=0}^{9} \widehat{\psi}'_j x_{t+j} + \sum_{j=1}^{4} \mu'_j \tilde{\pi}_{t-j} - v_t^{\pi}, \qquad (8)$$

where error term, v_t^{π}, is an autoregressive process of order one for a negative technological shock, $\epsilon_t^{\pi} \sim NID(0, \Sigma^{\pi})$: $v_t^{\pi} = \rho_x v_{t-1}^{\pi} + \epsilon_t^{\pi}$.

The other structural equations are:

- dynamic IS curve with habit persistence:

$$x_t = \gamma \mathrm{E}_t(x_{t+1}) + (1-\gamma) x_{t-1} - \sigma_1(\tilde{\imath}_t - \mathrm{E}_t \tilde{\pi}_{t+1}) + v_t^{x}, \qquad (9)$$

where the parameters depend on σ, and h ($\gamma = \frac{\sigma}{\sigma + h(\sigma-1)}, \sigma_1 = \frac{1}{\sigma + h(\sigma-1)}$), and the error term, v_t^x, is an autoregressive process of order one for a preference shock, $\epsilon_t^x \sim NID(0, \Sigma^x)$: $v_t^x = \rho_x v_{t-1}^x + \epsilon_t^x$.

- Taylor rule with interest rate smoothing:

$$\tilde{\imath}_t = \lambda \tilde{\imath}_{t-1} + (1-\lambda)(\phi_\pi \tilde{\pi}_t + \phi_x x_t) + \epsilon_t^i, \qquad (10)$$

where, ϵ_t^i, is a white noise monetary policy shock: $\epsilon_t^i \sim NID(0, \sigma^i)$.

In next section we compare the dynamics of the system to a benchmark time-dependent pricing framework with the hybrid Phillips curve (NHPC) of Gali and Gertler (1999):

$$\tilde{\pi}_t = \beta_f \mathrm{E}_t(\pi_{t+1}) + \beta_b \pi_{t-1} + \chi_0 x_t + \chi_1 x_{t-1} + \nu_t^{\pi}, \tag{11}$$

where $\beta_f = \frac{\beta\theta}{\theta+\omega(1-\theta(1-\beta))}, \beta_b = \frac{\omega}{\theta+\omega(1-\theta(1-\beta))}$ and $\chi_0 = \frac{\varphi+\alpha+\sigma(1-\alpha)}{1-\alpha}\frac{(1-\omega)(1-\theta)(1-\beta\theta)}{\theta+\omega(1-\theta(1-\beta))}, \chi_1 = \frac{h(1-\sigma)(1-\omega)(1-\theta)(1-\beta\theta)}{\theta+\omega(1-\theta(1-\beta))}$.

Results

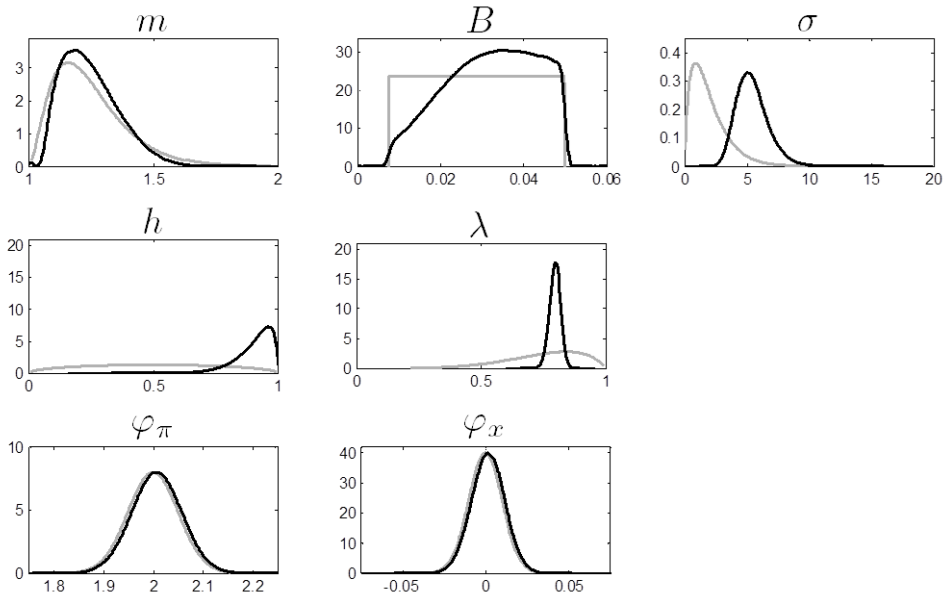
The results included in Table 1, and depicted in Fig. 4 and Fig. 5 come from a simulation of two Markov chains (1 million each with 25% burn-in initial cycles) following Metropolis random-walk algorithm implemented in Dynare. We have started the simulations with possibly loose priors. The parameter, B, describing an upper bound on a menu cost has been initiated with a flat prior on an interval close to a grid domain. With an adjusted gamma distribution we have restricted the prior distribution of gross markup above one. The mean of its prior distribution (1.25) gives an average elasticity of substitution between products equal to 5. The mean of a parameter σ about 2 is a safe assumption in a DSGE modelling strategy. The prior distribution of h, ρ_π, ρ_x was bounded by beta distribution in $[0; 1]$ interval, and standard deviations of shocks come from inverse gamma distribution, which are standard assumptions in many DSGE studies. From the parameters of a Taylor rule it was only λ which has been given considerably loose prior distribution.

Table 1. Characteristics of prior and posterior distribution in the state-dependent pricing model.

Parameter Names	prior distribution			posterior distribution	
	Type of distribution	Mean	Standard deviation	Mean	HPD interval 90%
m	Gamma $(1.\infty)$	1.25	0.15	1.2469	[1.0732; 1.4215]
B	Uniform [0.0075; 0.05]	0.029	-	0.0323	[0.0170; 0.0500]
σ	Gamma $(0.\infty)$	2.0	1.50	5.4286	[3.3479; 7.4149]
h	Beta	0.5	0.25	0.9050	[0.8069; 0.9993]
ϕ_π	Normal	2	0.05	2.0066	[1.9249; 2.0887]
ϕ_x	Normal	0	0.01	0.0014	[-0.0149; 0.0178]
λ	Beta	0.75	0.15	0.7970	[0.7609; 0.8348]
ρ_π	Beta	0.5	0.15	0.3371	[0.1765; 0.4966]
ρ_x	Beta	0.5	0.15	0.4082	[0.2629; 0.5574]
σ^π	Gamma^{-1}	1.0	∞	0.5260	[0.4350; 0.6141]
σ^x	Gamma^{-1}	0.4	∞	0.2460	[0.1886; 0.3007]
σ^i	Gamma^{-1}	1.0	∞	0.2299	[0.1934; 0.2653]

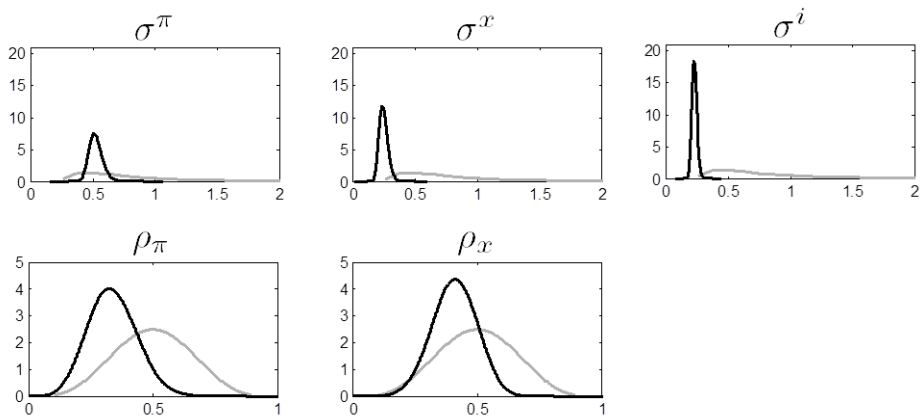
Source: Own calculations with Dynare 4.3 (Adjemian et. al. 2011).

Fig. 4. Posterior distributions of parameters of DSGE with DKW mechanism



Source: own calculations with Dynare 4.3 (Adjemian et. al. 2011)

Fig. 5. Posterior distributions of parameters of DSGE with DKW mechanism



Source: own calculations with Dynare 4.3 (Adjemian et. al. 2011).

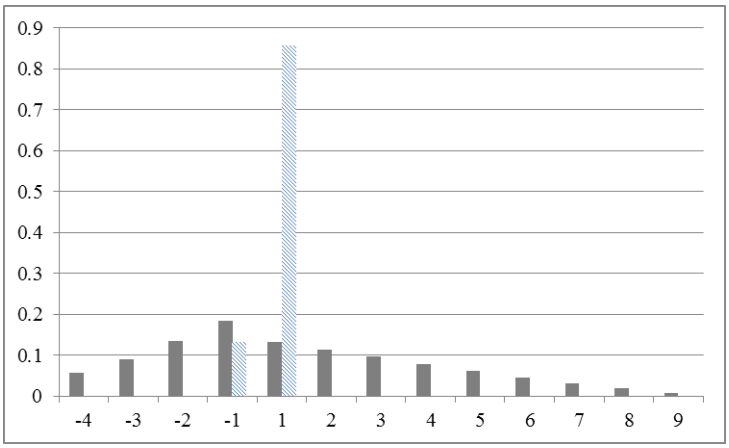
The Bayesian estimation of all of the parameters, except for a markup (m), significantly updates *a priori* knowledge of their distribution (see Figure 4). From a negative asymmetry of B posterior density it seems, that to allow for a degree of price stickiness observed in the data, one needs maximal menu costs to be on average above mean value of its prior density. A mean value of an upper bound of menu costs is 3.23% and in terms of real wages it reads in costs of circa 2.6% of real output. Although it is only an extreme and potential value of menu cost distribution it may make a considerable barrier for price adjustment.

Also a mean of the posterior distribution of σ which explains consumption smoothing preferences takes relatively big value. There is also a strong evidence of the inertial behaviour in a monetary policy (with λ about 0.8) and habit formation (with h about 0.9). All of the three parameters (σ, h, λ) distributions are very close to the time-dependent pricing case (see results in Appendix B). The only

significant difference lies in shock characteristics. In relation to the time-dependent pricing model the variance of technological and preference shocks are increased which is compensated by a lower persistence of a preference shock in the state-dependent pricing model.

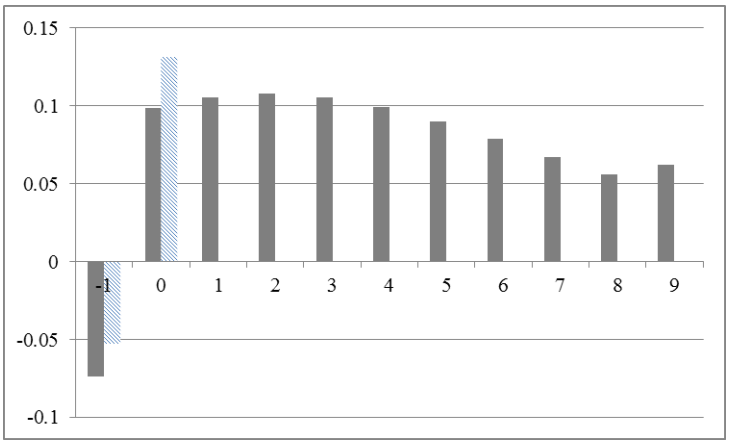
In the next step we make comparisons of the Phillips curve equations in both models because it is potentially an explanation of the differences between the two specifications. As the posterior distributions of the structural parameters (being the functions of posterior distributions of 'deep' DSGE parameters) are of a complex shape we analyse them in terms of medians only (see Fig. 6 and Fig. 7). Compared to NHPC (which is mostly forward looking) inflation expectations in SDPC are of much lower magnitude and they are less and less important with time horizon. SDPC gives an appealing economic (menu costs) explanation of intrinsic inflation persistence. In the perspective of SDPC it is a prevailing force of inflation determination in Polish economy. On the other side according to SDPC estimates there is a role for medium-term output gap expectations in determining inflation. At short-term expectations of the distribution of parameters at the forward-looking terms of SDPC increases with time horizon. A maximum is located at the terms of two quarters ahead, and then those median parameters decay very slowly. The dependence of inflation on past output gap (due to habit persistence) is also little stronger in SDPC compared to NHPC estimates.

Fig. 6. Median of posterior distributions of the parameters at lagged (-) and future (+) inflation in SDPC (in black) and NHPC (shaded)



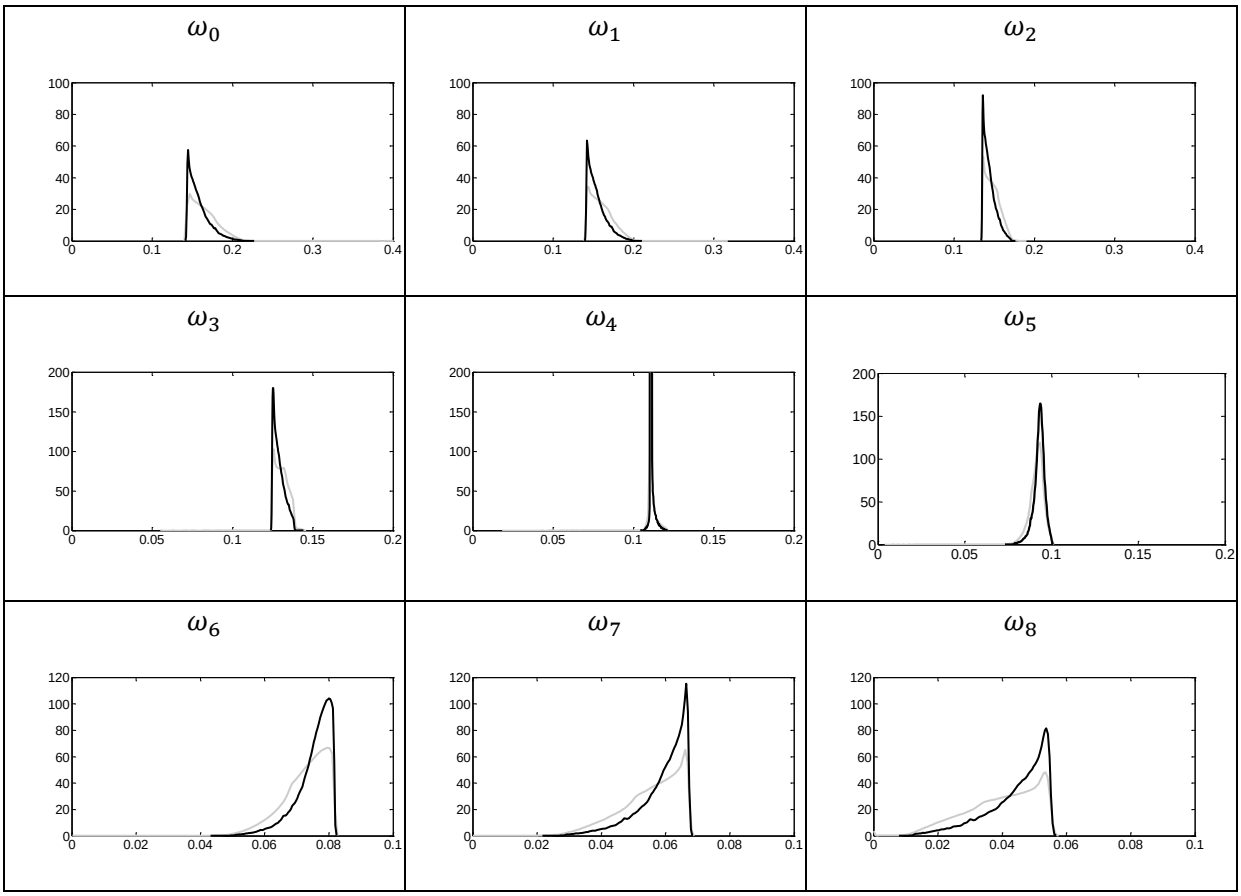
Source: Own calculations. Remark: in SDPC these are μ'_j $(j = -1, \dots, -4)$, and δ'_j $(j = 1 \dots 9)$, and in NHPC: β_b $(j = -1)$, and β_f $(j = 1)$.

Fig. 7. Median of posterior distribution of parameters at lagged (-), current ($j = 0$) and future (+) output gap in SDPC (black) and NHPC (shaded).



Source: Own calculations. In SDPC: ϑ for $j = -1$, $\hat{\psi}_j{}' j = 0, \ldots, 9$, and in NHPC: χ_1 i χ_0,

Fig. 8. Prior and posterior distribution of fraction of firms in price vintages ω_j at steady state in the state-dependent pricing model

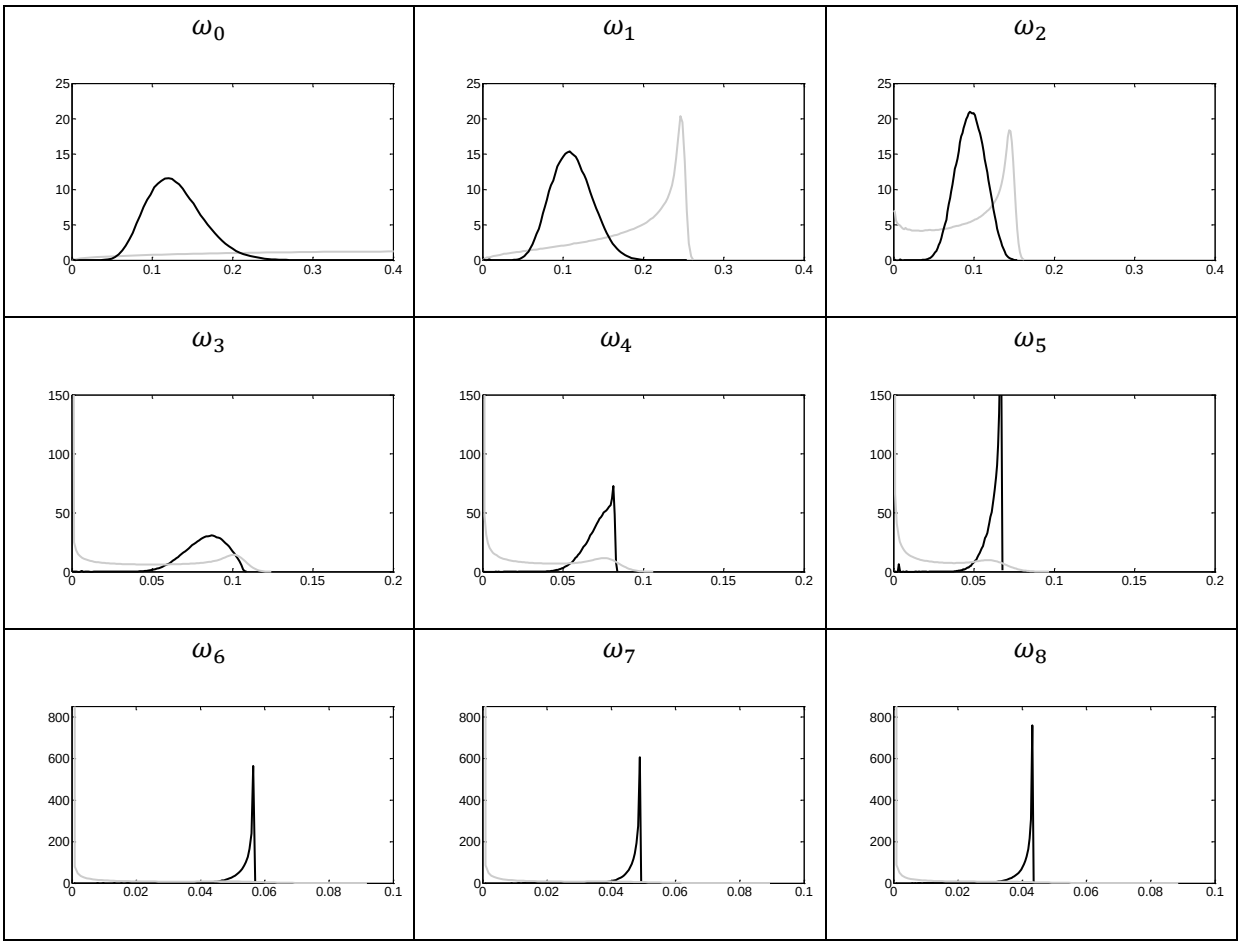


Source: Own calculations.

The next point in a Bayesian analysis of both pricing mechanisms is a comparison of steady-state distribution of their respective price vintages, which are equivalent to ω_j in SDPC (see Fig. 8 for state-dependent and Fig. 9 for time-dependent case). The posterior histograms of firms distributions

across price vintages in a steady state (black line) are depicted in the same picture with *a priori* beliefs of the parameters (grey line) to realize the extent of information update after looking at the data, and the limitations of assumptions on prior distribution of deep parameters. In NHPC the number of price vintages is infinite, and in SDPC estimation we have cut the impact at 10. To make a comparison possible we limit the presentation to the distributions of finite number of price vintages (9 for a convenience).

Fig. 9. Prior and posterior distribution of fraction of firms in price vintages $\omega_j = (1-\theta)\theta^j$ at steady state in the time-dependent pricing model



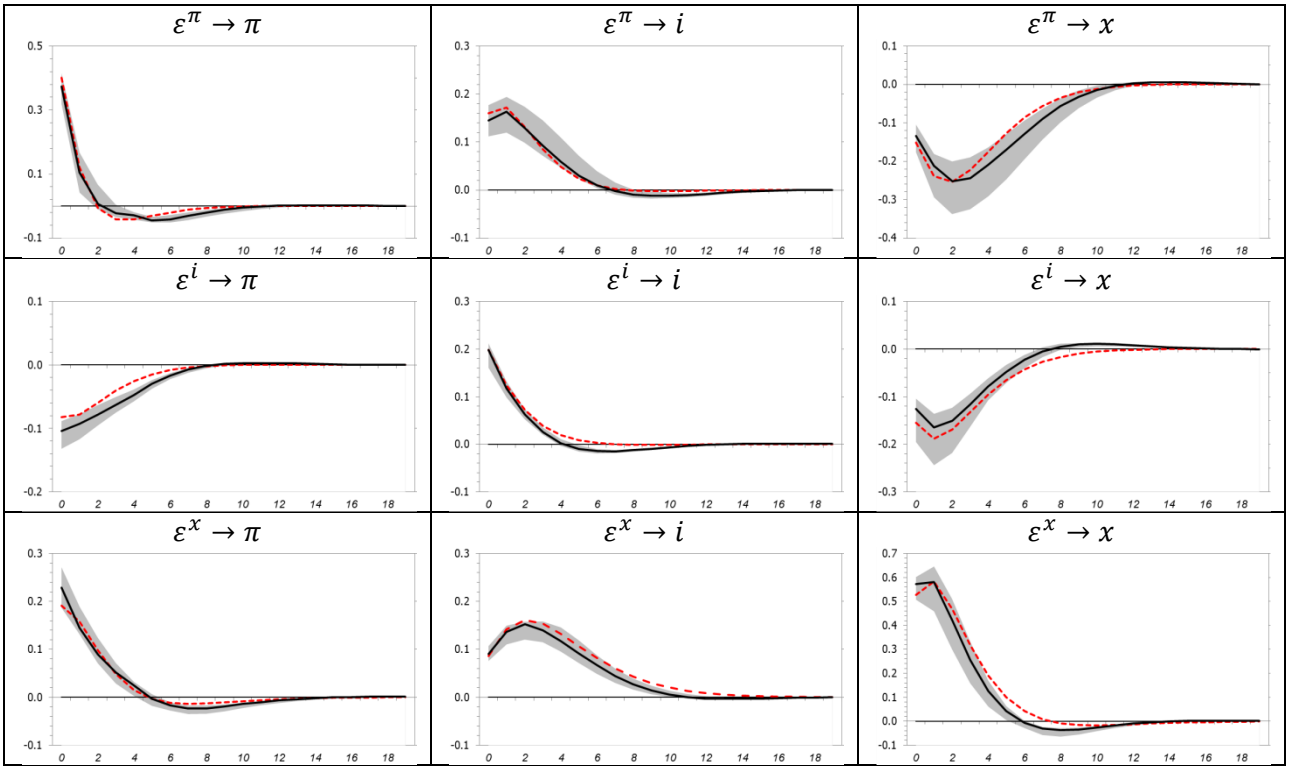
Source: Own calculations.

From Fig. 8 and 9 it is important to learn that the mode of the distribution is decaying faster in SDPC than in NHPC. In effect it gives a median duration of price at about 4 quarters in SDPC, and about 8 quarters in NHPC, the difference that is surprisingly and which is also very important from micro-price studies perspective. These results for state-dependent economy should be treated with higher caution, as the posterior distributions of ω_j in SDPC are to some extent (at least in shape) determined by their prior distribution derived from prior distribution of m and B. Contrary in NHPC the extent of information update clearly increases with the number of a price vintage.

The last part of the Bayesian analysis is about comparing impulse response functions (IRF) in both DSGE models using their posterior distributions. In the Fig. 10 we put a red line for an average IRF

path of responses in the time-dependent pricing model, and a shaded area and a black line are, respectively, 90% HPD interval and median of state-dependent pricing model.

Fig. 10. Impulse response functions of inflation, interest rates, and output gap to standard deviation of empirical distribution of shocks to technology ε^{π}, monetary policy ε^{i} and preferences ε^{x}.



Source: Own calculations with Dynare 4.3.

The results of impulse response functions in both time- and state-dependent model are economically plausible. They also exhibit similar hump-shaped pattern of reaction widely observed in the data. In a first row there are the effects of markup shocks on three observed variables. The unanticipated negative shock to technology makes reoptimising firms to set higher prices, both in DKW and time-dependent pricing model. In response to a higher inflation central bank raises interest rate, what generates negative output gap. In a second row there are the effects of monetary policy shock. An 'extra' raise of central bank interest rate (above the interest rate consistent with the Taylor rule) tend households to delay a part of its current consumption. This generates negative output gap, which results in lower inflation. Third row describes the effects of a preference shock. The shock raises the weight of current utility in the lifetime utility path. This makes current consumption even more valuable, which results in a negative output gap and consequently higher inflation. At last central bank raises interest rate in response to higher economic activity. It is important to remember that the time-dependent and state-dependent IRFs estimated for the Polish economy are generally hard to distinguish. The differences being neglible in economic terms, are statistically significant only for reaction of after monetary policy shock (slower responses of inflation and interest rates, and quicker adjustments in output gap in the model with SDPC), and the responses to technological shocks (faster adjustment despite the higher persistence of these estimated shocks in the state-dependent pricing model).

Though the estimated DSGE model with state-dependent pricing identifies less persistent preference shocks than a comparable time-dependent model. Also the effects of markup shocks for inflation, output and interest rates are less persistent and weaker in state-dependent model. It may be attributed to a 'selection effect' (i.e. firms that change price after the shock are mostly the ones with price far from optimal level).

Conclusions

The estimated state-dependent pricing model generates a median duration of prices above 4 quarters compared to 8 quarters in a time-dependent model. On the other hand in a state-dependent pricing model inflation after a monetary policy shock approaches a steady state path in a similar (or even in a more prolonged) pace relative to a time-dependent pricing model. DKW model is also able to identify higher variance of technology shocks, and higher persistence in preference shocks. Despite those significant differences the dynamics of impulse responses in time- and state-dependent pricing models are hard to distinguish. It may have important consequences in interpreting the results of microeconomic surveys of price stickiness.

The results obtained in the state dependent pricing model should be treated with caution. We have analysed only small-scale models without many important nominal and real frictions. Secondly, the estimation sample contains a 5-year period of considerable disinflation process. We have used HP-filtered series to cope with the issue in a proper manner. The other limit of our study comes from the state-dependent terms which have been omitted to facilitate the Bayesian estimation. The conclusions from IRF analysis should be correct for the Polish economy if small shocks are considered and inflation is close to a 4% inflation steady state. Especially, the lagged response of inflation to the monetary policy shock in state-dependent pricing model may be the result of neglecting short-term changes to the fractions of firms in price vintages due to a 'selection effect'.

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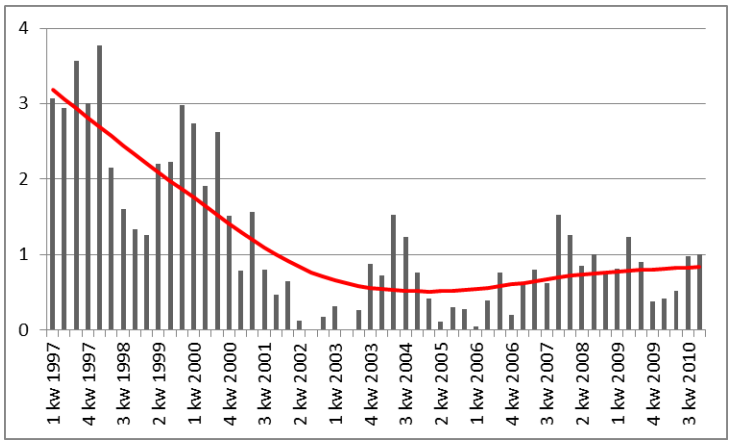
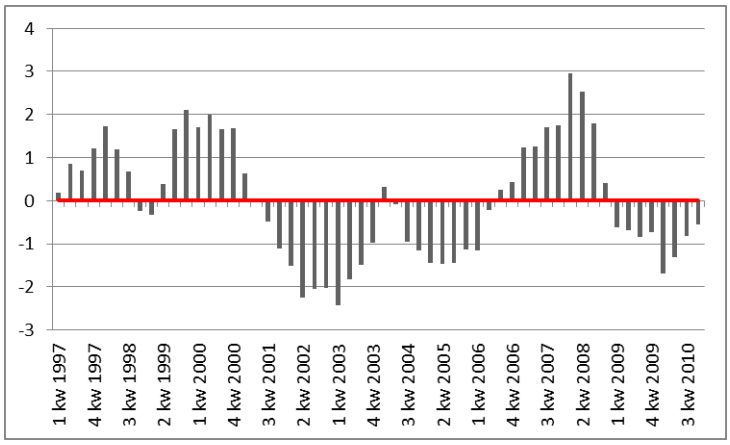
Appendix A. Data for Polish economy (1997-2010)

Fig. 1A. Inflation in Poland (in percentage points per quarter) and HP-filter trend.



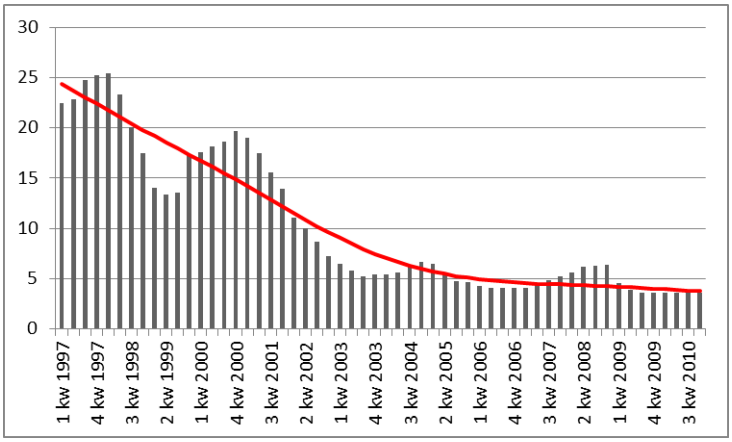
Source: Central Statistical Office (GUS). Own calculations.

Fig. 2A: Output gap in Poland (% deviations from HP-filter trend).



Source: Central Statistical Office (GUS). Own calculations.

Fig. 3A: Short-term interest rate (WIBOR) in Poland and its HP-filter trend.



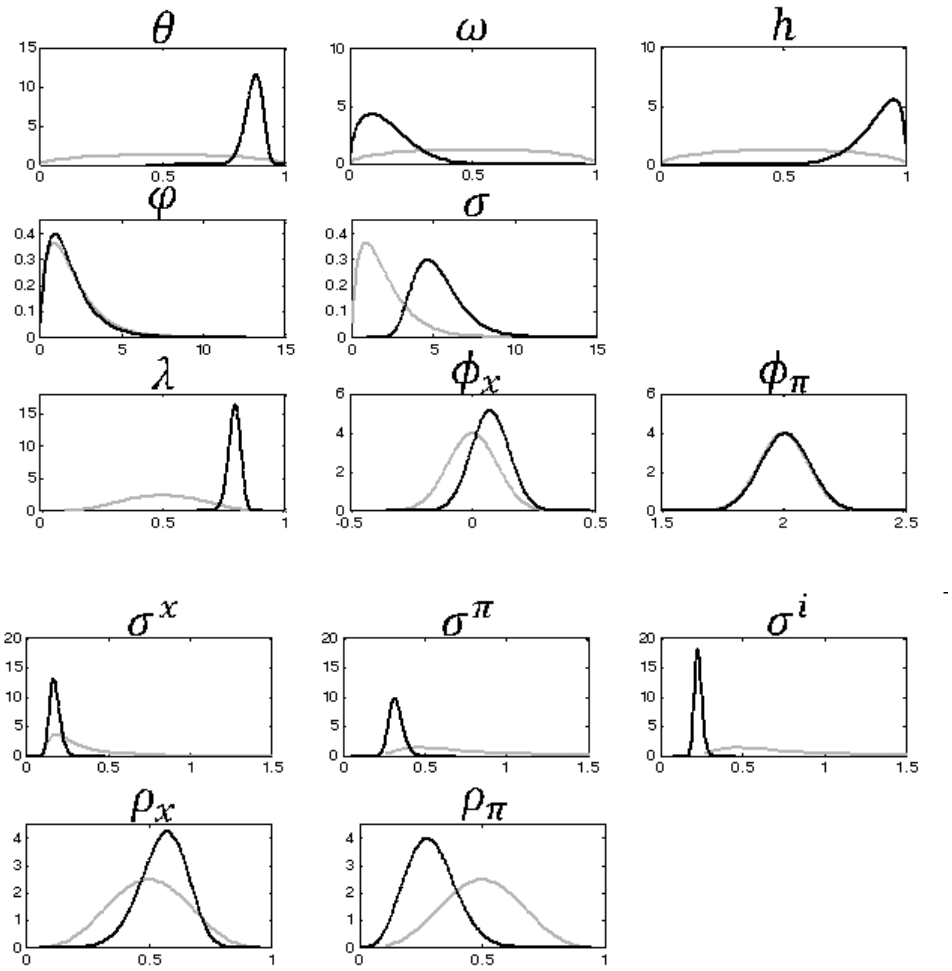
Source: Central Statistical Office (GUS). Own calculations.

Appendix B. The estimation results of a small-scale model with the Gali, Gertler (1999) pricing mechanism

Table A1. Characteristics of prior and posterior distribution in the time dependent pricing model.

Parameter names	Prior distribution			Posteriori distribution	
	Type of distribution	Mean	Standard deviation	Mean	HPD interval (90%)
θ	Beta	0.5	0.25	0.8715	[0.8158; 0.9283]
ω	Beta	0.5	0.25	0.1495	[0.0038; 0.2851]
h	Beta	0.5	0.25	0.8786	[0.7555; 0.9988]
φ	Gamma	2.0	1.50	1.8283	[0.0514; 3.6360]
σ	Gamma	2.0	1.50	5.2919	[3.0074; 7.5548]
ϕ_π	Normalny	2.0	0.10	2.0060	[1.8458; 2.1717]
ϕ_x	Normalny	0.0	0.10	0.0723	[-0.0556; 0.1982]
λ	Beta	0.5	0.15	0.7941	[0.7549; 0.8344]
ρ_π	Beta	0.5	0.15	0.2908	[0.1279; 0.4470]
ρ_x	Beta	0.5	0.15	0.5569	[0.4036; 0.7097]
σ^π	Gamma^{-1}	1.0	∞	0.3234	[0.2548; 0.3899]
σ^x	Gamma^{-1}	0.4	∞	0.1760	[0.1251; 0.2270]
σ^i	Gamma^{-1}	1.0	∞	0.2301	[0.1925; 0.2661]

Fig. 1A Prior and posterior distribution of parameters in the time dependent pricing model.



Appendix C. DSGE model structure

Households

We consider representative households maximizing intertemporal utility from their consumption ($C_t(i)$). and disutility of labor (N_t)[4]:

$$U(\,C_t(i).N_t) = e^{v_t^p}\left(\frac{[C_t(i)/(\bar{C}_{t-1})^h]^{1-\sigma}}{1-\sigma} - \frac{{N_t}^{1+\varphi}}{1+\varphi}\right). \tag{12}$$

where $\sigma > 0$ is a constant relative risk aversion. $\varphi > 0$ is the inverse of Frisch elasticity of labour. and $0 < h < 1$ is a measure of external habit persistence (Abel 1990). which is measured in relation to the average consumption (across all firms) from the previous period $(\bar{C}_{t-1})^h$. and v_t^p is an AR(1) process. which we interpret as a preference shock in period t (i.e. a shock that shifts consumer tastes).

Technology and aggregation

Firms indexed with $j \in (0.1)$ transform labor to products given initial technology level A_0. and aggregate technology shocks v_t^m:[5]

$$Y_t(j) = A_0 e^{v_t^m} N_t^{1-\alpha}. \tag{13}$$

where v_t^m is a stationary AR(1) stochastic process. and $0 < (1-\alpha) < 1$ is a labor share.

With Dixit-Stiglitz (1977) monopolistic conditions we define aggregate consumption and price as:

$$C_t = \left(\int_0^1 C_t(j)^{(\varepsilon-1)/\varepsilon} dj\right)^{\varepsilon/(\varepsilon-1)}. \tag{14}$$

$$P_t = \left(\int_0^1 P_t(i)^{1-\varepsilon} di\right)^{1/(1-\varepsilon)}. \tag{15}$$

where $\varepsilon > 1$ is constant elasticity of substitution between goods or price elasticity of consumption $C_t(j) = \left(\frac{P_t(j)}{P_t}\right)^{-\varepsilon} C_t$.

DKW – firm pricing decisions

In period t the firm decision requires calculation the firm's real values:

$$\begin{aligned} v_{0.t} &= \max_P\{z_{0.t}(P) + E_t Q_{t.t+1} \cdot (1-\alpha_{1.t+1}) \cdot v_{1.t+1} + E_t Q_{t.t+1}\alpha_{1.t+1} \\ &\quad \cdot (v_{0.t+1} - \mathrm{W}_{t+1} \cdot \mathrm{K}_{1.t+1})\} \\ v_{j.t} &= z_{j.t}(P_{t-j}^*) + E_t Q_{t.t+1} \cdot \left(1-\alpha_{j.t+1}\right) \cdot v_{j+1.t+1} + E_t Q_{t.t+1}\alpha_{j.t+1} \\ &\quad \cdot \left(v_{0.t+1} - \mathrm{W}_{t+1} \cdot \mathrm{K}_{j.t+1}\right). \end{aligned} \tag{16}$$

[4] We assume competitive labor market with firms renting labour time of a representative household N_t at an economy wide real wage rate W_t. The household decisions are also subject to standard budget constraint.
[5] Alternatively, one can assume shocks to the monopolistic markup, which is hard to distuinguish from negative technology shocks in the reduced Phillips curve relation (see Smets, Wouters 2003).

$$j = 1.2. \ldots . J - 1.$$

where $z_{j.t}(P^*_{t-j}) = \left(\frac{P^*_{t-j}}{P_t}\right)^{-\varepsilon} \cdot Y_t \cdot \frac{P^*_{t-j}}{P_t} - \Psi_{t.j}$ is the firm's current period real profit if it nominal price is P^*_{t-j}. $\mathrm{K}_{j.t+1}$ is the average menu cost in vintage j and the term $Q_{t.t+k} = \beta^k U'(C_{t+k})/U'(C_t)$ represents stochastic discount factor for the future real profits (see Campbell. 1999).

The solution to the problem of maximization the firm's real value $v_{0.t}$ is given by (cf. Dotsey. King. Wolman [1999]. p. 665):

$$P^*_t = \frac{\varepsilon}{\varepsilon - 1} \frac{\sum_{j=0}^{J-1} \beta^j \cdot E_t Q_{t.t+j} \cdot \frac{\omega_{j.t+j}}{\omega_{0.t}} \cdot MC_{t+j} \cdot P^{\varepsilon}_{t+j} \cdot Y_{t+j}}{\sum_{j=0}^{J-1} \beta^j \cdot E_t Q_{t.t+j} \cdot \frac{\omega_{j.t+j}}{\omega_{0.t}} \cdot P^{\varepsilon-1}_{t+j} \cdot Y_{t+j}}. \tag{17}$$

where MC_t is the real marginal cost and $\frac{\omega_{j.t+j}}{\omega_{0.t}}$ is the probability of nonadjustment the price from t to period $t + j$. In the case of flexible prices ($B = 0$) the formula (17) can be rewritten as $\frac{P^*_t}{P_t} = \frac{\varepsilon}{\varepsilon-1} MC$. Hence the term $\frac{\varepsilon}{\varepsilon-1}$ can be interpret as a monopolistic markup.

Central bank monetary rule

The last type of shocks $\left(\varepsilon^i_t\right)_{t=1.2\ldots}$ is a monetary policy shock. It is interpreted in terms of deviations of a central bank from a current Taylor rule with smoothing.

Non-zero steady state and derivation of SDPC

The steady state of the economy is defined as the constant level of inflation Π^{SS}. total production Y and stationary distribution of prices $\omega_0. \omega_1. \ldots . \omega_{J-1}$. Moreover. denote by $RP^*_t = \frac{P^*_t}{P_t}$ relative optimal price in period t and notice that the steady state value of RP^*_t is constant in time and given by $RP^* = \frac{\varepsilon}{\varepsilon-1} \frac{\sum_{j=0}^{J-1} \beta^j \cdot \omega_j \Pi^{j\varepsilon}}{\sum_{j=0}^{J-1} \beta^j \cdot \Pi^{j(\varepsilon-1)}} MC$.

Hence. DKW pricing mechanism in the steady state is described by time homogenous stationary Markov chain with states $1. \ldots . J$ denoting the price vintages and with transition matrices. M:

$$\mathrm{M} = \begin{bmatrix} \alpha_1 & 1-\alpha_1 & 0 & \ldots & 0 \\ \alpha_2 & 0 & 1-\alpha_2 & \ldots & 0 \\ \ldots & \ldots & \ldots & \ldots & \ldots \\ \alpha_{J-1} & 0 & 0 & \ldots & 1-\alpha_{J-1} \\ 1 & 0 & 0 & \ldots & 0 \end{bmatrix} \tag{18}$$

The probabilities $\alpha_1. \alpha_2. \ldots . \alpha_J$ with $\alpha_J = 1$ are the solution to the constrain optimization problem:

$$v_0 = \max_{RP}\{z_0(RP) + (1-\alpha_1) \cdot v_1 + \alpha_1 \cdot (v_0 - \mathrm{W} \cdot \mathrm{K}_1)\}$$

$$v_j = z_j(RP^*) + \left(1 - \alpha_{j.t+1}\right) \cdot v_{j+1} + \alpha_j \cdot \left(v_0 - \mathrm{W} \cdot \mathrm{K}_j\right). \tag{19}$$

$$\alpha_j = G((v_0 - v_j)/\mathrm{W})$$

$$j = 1.2. \ldots . J - 1.$$

where $z_j(RP^*) = (RP^*)^{1-\varepsilon} \cdot Y \cdot \Pi^{j(\varepsilon-1)} - MC \cdot Y$. In consequence. the sums of conditional discounted current and future profits $v_0. v_{j.}$are constant in time.

To derive the Phillips Curve one has to solve the optimal problem for firms and aggregate the price distribution. The Dixit-Stiglitz (1977) price aggregation entails: $P_t^{1-\varepsilon} = \sum_{j=0}^{J-1} \omega_{j.t} \left(P_{t-j}^*\right)^{1-\varepsilon}$. Then.on substituting the formula for relative optimal prices RP_t into the above equation one obtains:

$$1 = \sum_{j=0}^{J-1} \omega_{j.t} \left(RP_{t-j}^* \frac{P_{t-j}}{P_t}\right)^{1-\varepsilon} \tag{20}$$

Log-linearization around steady-state leads to:

$$\widetilde{rp}_t^* = \frac{1}{\omega_0}\left[\sum_{j=0}^{J-2} \tilde{\pi}_{t-j} \sum_{i=j+1}^{J-1} \omega_i \Pi^{i(\varepsilon-1)} - \sum_{j=1}^{J-1} \omega_j \Pi^{j(\varepsilon-1)} \widetilde{rp}_{t-j}^* + \frac{1}{\varepsilon-1} \sum_{j=0}^{J-1} \omega_j \widetilde{\omega}_{j.t} \Pi^{j(\varepsilon-1)}\right]. \tag{21}$$

where variables with a tilde are deviations from steady-state: $\widetilde{\omega}_{j.t} = \ln \omega_{j.t} - \ln \omega_j$. $\tilde{\pi}_{t-j} = \ln \Pi_{t-j} - \ln \Pi$. $\widetilde{rp}_{t-j}^* = \ln(RP_t^*) - \ln(RP^*)$. $j = 0.1. \ldots J-1$.

After log-linearization of formula for relative optimal price (see Appendix A to Bakhsi. Khan i Rudolf. 2006) and assuming $\tilde{q}_{t.t+j} = 0$. we obtain:

$$\widetilde{rp}_t^* = E_t \sum_{j=0}^{J-1} \left[\widetilde{\omega}_{j.t+j} - \widetilde{\omega}_{0.t} + \widetilde{mc}_{t+j} + \varepsilon \sum_{i=1}^{j} \tilde{\pi}_{t+i} + x_{t+j}\right] \rho_j - E_t \sum_{j=0}^{J-1} \left[\widetilde{\omega}_{j.t+j} - \widetilde{\omega}_{0.t} + (\varepsilon-1) \sum_{i=1}^{j} \tilde{\pi}_{t+i} + x_{t+j}\right] \delta_j. \tag{22}$$

where x_t is an output gap. $\delta_j = \frac{\beta^j \omega_j \Pi^{j(\varepsilon-1)}}{\sum_{i=0}^{J-1} \beta^i \omega_i \Pi^{i(\varepsilon-1)}}$ and $\rho_j = \frac{\beta^j \omega_j \Pi^{j\varepsilon}}{\sum_{i=0}^{J-1} \beta^i \omega_i \Pi^{i\varepsilon}}$.

Consumer's habit persistence leads to the following relationship between percentage deviation of real marginal cost from its steady-state value and output gap:

$$\widetilde{mc}_t = \kappa_1 x_t + \kappa_2 x_{t-1} + \frac{\phi+1}{1-\alpha} d_t + \frac{\phi+1}{\alpha-1} v_t^m.$$

where $\kappa_1 = \frac{\phi+\alpha+(1-\alpha)\sigma}{1-\alpha}$. $\kappa_2 = h(\sigma-1)$.

Substituting this equation together with $E_t v_{t+j}^m = 0$ for $j = 1.2. \ldots J-1$ into (13) gives:

$$\widetilde{rp}_t^* = E_t \sum_{j=1}^{J-1} \tilde{\pi}_{t+j} \sum_{i=j}^{J-1} (\varepsilon\rho_i - (1-\varepsilon)\delta_i) + E_t \sum_{j=0}^{J-1} \left(\rho_j - \delta_j\right)(\widetilde{\omega}_{j.t+j} - \widetilde{\omega}_{0.t}) + \kappa_2 \rho_0 x_{t-1} + E_t \sum_{j=0}^{J-1} \hat{\psi}_j x_{t+j} + \frac{\phi+1}{1-\alpha} d_t + \frac{\phi+1}{\alpha-1} v_t^m . \tag{23}$$

where:

$$\hat{\psi}_j = \begin{cases} \kappa_1\rho_j + \rho_j - \delta_j + \kappa_2\rho_{j+1} & j = 0.1.\ldots.J-2 \\ \kappa_1\rho_j + \rho_j - \delta_j & j = J-1 \end{cases}$$

Comparing (21) and (14) we obtain:

$$\tilde{\pi}_t = E_t \sum_{j=1}^{J-1} \tilde{\pi}_{t+j} \delta'_j + E_t \sum_{j=0}^{J-1} \gamma_j (\widetilde{\omega}_{j.t+j} - \widetilde{\omega}_{0.t}) + \frac{1}{\mu_0} \kappa_2 \rho_0 x_{t-1} + E_t \sum_{j=0}^{J-1} \frac{1}{\mu_0} \hat{\psi}_j x_{t+j} - \sum_{j=1}^{J-2} \frac{\mu_j}{\mu_0} \tilde{\pi}_{t-j} + \sum_{j=1}^{J-1} \frac{\omega_j}{\mu_0 \omega_0} \Pi^{j(\varepsilon-1)} \widetilde{rp}_{t-j}^* + \frac{1}{1-\varepsilon} \widehat{\Omega}_{\mathrm{t}} + \frac{1}{\mu_0} \frac{\phi+1}{1-\alpha} d_t + \frac{1}{\mu_0} \frac{\phi+1}{\alpha-1} v_t^m \tag{24}$$

$\delta'_j = \frac{1}{\mu_0} \sum_{i=j}^{J-1} (\varepsilon\rho_i - (1-\varepsilon)\delta_i)$ for $j = 1.2.\ldots.J-1$. $\gamma_j = \frac{1}{\mu_0}\left(\rho_j - \delta_j\right)$ for $j = 0.2.\ldots.J-1$.
$\mu_j = \frac{1}{\omega_0} \sum_{i=j+1}^{J-1} \omega_i \Pi^{i(\varepsilon-1)}$ for $j = 0.1.2.\ldots.J-2$. $\widehat{\Omega}_{\mathrm{t}} = \sum_{j=0}^{J-1} \frac{1}{\omega_0} \Pi^{j(\varepsilon-1)} \omega_{\mathrm{j}} \widetilde{\omega}_{\mathrm{j.t}}$.

To derive State Dependent Phillips Curve from equation (24) one needs to recurrently substitute $\widetilde{rp}_{t-1}^*. \widetilde{rp}_{t-2}^*. \ldots.$ formulas from (22) .(see Appendix B in Bakhshi. Khan. Rudolf. 2006).

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