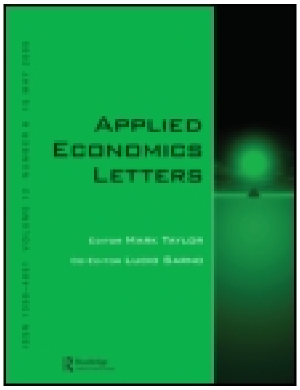

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How frequently do consumer prices change in transition countries?

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How frequently do consumer prices change in transition countries?

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A common feature of recent literature regarding inflation dynamics is the Calvo pricing mechanism. Using this model and aggregate series, I estimate the price change probability and the mean time between price changes in 13 transition countries. The average price change probability is much larger than suggested by the New Keynesian Phillips Curve (NKPC) literature. The corresponding mean time between price changes is slightly longer than 6 months. Moreover, a forward-looking pricing has been found only for four countries.

Keywords: price change probability; transition; forward-looking pricing; CEE countries

JEL Classification: E31; P22

I. Introduction

New Keynesian Phillips Curve (NKPC) literature (e.g. Gali and Gertler, 1999; Sbordone, 2002, 2005) suggests that prices in the United States are ‘frozen’ for approximately 18 months. In Western Europe, firms change their prices once a year (Álvarez *et al*., 2005). Some recent microeconomic studies (Bils and Klenow, 2004; Baumgartner *et al*., 2005; Lünemann and Mathä, 2005; Coenen *et al*., 2007), however, provides some evidence on relatively high price change probability.

NKPC literature and microeconomic studies utilize different sets of data and thus the results are not directly comparable. Aggregation may influence the obtained outcomes to a great extent. To make the results comparable, a microeconomic background should meet the aggregate data analysis. A simple way to estimate the probability of price changes using the aggregate price series offers the Calvo model.

The historical and institutional background clearly suggests that the prices in transition countries should be more rigid than in economies with long-standing market traditions. In this article, I estimate the price change probability and the mean time between price changes in Central and Eastern European (CEE) countries. I also trace out whether prices are set in accordance with the forward-looking mechanism by regressing the quotation series against the actual lead values of the output gap and then by conducting the impulse response analysis.

II. The Model

In a discrete time version of the Calvo (1983) pricing mechanism, current price is a function of discounted sum of the lagged price quotations ν:

$$p_t = \delta \sum_{j=0}^{\infty} (1-\delta)^j \nu_{t-j} \tag{1}$$

The time-independent probability of price change in period t is equal to δ. A representative firm receives a price-change signal regardless of the previous signal receptions. Thus, I assume that the price quotation is a white noise process satisfying the standard Ordinary Least Square (OLS)-residual properties. Rearranging Equation 1 gives

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$$p_t = \delta\nu_t + (1-\delta)p_{t-1} \quad (2)$$

Using the previous assumption that ν is a white noise process, I rewrite Equation 2 as an first-order Autoregressive (AR(1)) model:

$$p_t = \theta p_{t-1} + \varepsilon_t \quad (3)$$

where $\theta \equiv 1-\delta$ and $\varepsilon_t \equiv \delta\nu_t$. The residuals ε_t capture information on price-change signal, so the lack of information on ν is neutralized. Hence, the probability of a price change is

$$\delta = 1-\theta \quad (4)$$

and the mean time between price changes is

$$T = \frac{1}{\delta} \quad (5)$$

III. Estimating the Price Change Probability

I estimate the price change for a group of 13 CEE countries: Belarus, Bulgaria, Croatia, the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Romania, Russia, Slovenia and the Slovak Republic. For the sake of comparison, I use the total Consumer Price Indices (CPIs), which are available for the entire group. Since the price series are not stationary, I employ the Hodrick–Prescott (HP) filter with the standard smoothing parameter. I present the estimation results for both seasonally unadjusted and adjusted series. The employed samples are monthly and start, in most cases, in January 1996 (see Table 1 for description). Since the sudden speed-ups in inflation have produced some structural breaks, I resample the series for Belarus, Bulgaria, Romania and Russia.

Estimation results for the seasonally unadjusted series

For the group of 13 countries, the mean price change probability is 14.8% (see the upper panel of Table 2). The corresponding mean time between price changes is 6.7 months. It is reasonable to presume that the large impact of the public sector combined with the centralistic tradition should have produced a smaller price change probability. In fact, I found the smallest value of δ for Poland, a country which is sometimes acknowledged to be the leader of the market transition in Central Europe. In a sharp contrast to the Polish case, the largest price change probability is estimated for Russia, where the price duration slightly exceeds one quarter. This frequency is similar to the average speed of transmission from changes in the growth of broad money to inflation (Nikolić, 2000). The results are also in tact with the findings concerning the influence of monetary policy variables on prices in Commonwealth of Independent States (CIS) presented by Starr (2005). Equally frequent price changes are found in Croatia and Slovenia. Among the investigated countries, price changes also occur twice a year in Belarus. The other economies form a group with price change probabilities between 11% and 16%. The 12% probability for Romania is smaller than the survey-based evidence presented by Copaciu *et al.* (2010). The estimated time between price changes is 3 months longer than the results obtained using micro data. The discrepancy might be caused by the sample size.

None of the countries, however, is characterized by a nominal price stickiness similar to the one assumed by the NKPC literature. The 8.8-month T for Poland is twice smaller than the 18 months reported by NKPC literature. Moreover, the Kashyap (1995) hypothesis, which stated that prices changed more often during the periods of high inflation, is not confirmed. Since the transition countries often suffer from relatively high inflation, connecting the large values of δ to inflation seems to be reasonable. The CEE data, however, provide little evidence on price change probability–inflation relationship. Figure 1(a) plots the estimated probabilities (in percentage) against the variance of the monthly inflation. The trend line tends slightly upwards, but the fit is very poor. An equally poor fit is obtained after plotting the probabilities against the average monthly inflation, as Fig. 1(b) depicts.

Table 1. Sample description

Country	Sample	Country	Sample
Belarus (BEL)	January 2000 to February 2008	Lithuania (LTH)	January 1996 to February 2008
Bulgaria (BUL)	March 1997 to February 2008	Poland (POL)	January 1996 to February 2008
Croatia (CRO)	January 1996 to February 2008	Romania (ROM)	April 1997 to February 2008
Czech Republic (CZ)	January 1996 to February 2008	Russia (RUS)	January 1999 to February 2008
Estonia (EST)	January 1996 to February 2008	Slovenia (SLO)	January 1996 to February 2008
Hungary (HUN)	January 1996 to February 2008	Slovak Republic (SLK)	January 1996 to February 2008
Latvia (LTV)	January 1996 to February 2008		

Table 2. Estimates of price change probabilities

Country	Probability	Time	Country	Probability	Time
Probability: Equation 4					
Belarus	17.1	5.9	Bulgaria	12.9	7.7
Croatia	20.2	5.0	Czech Republic	12.6	7.9
Estonia	11.9	8.4	Hungary	13.7	7.3
Latvia	11.7	8.5	Lithuania	14.4	7
Poland	11.4	8.8	Romania	12	8.3
Russia	20.3	4.9	Slovak Republic	16.1	6.2
Slovenia	18.6	5.4	Average probability = 14.8% Average time = 6.7		
Probability: Equation 6					
Belarus	18.7	5.3	Bulgaria	13.8	7.2
Croatia	22.5	4.4	Czech Republic	13.5	7.4
Estonia	12.7	7.9	Hungary	14.7	6.8
Latvia	12.5	8	Lithuania	15.5	6.4
Poland	12.1	8.3	Romania	12.8	7.8
Russia	22.7	4.4	Slovak Republic	17.5	5.7
Slovenia	20.6	4.9	Average probability = 16.1% Average time = 6.2		

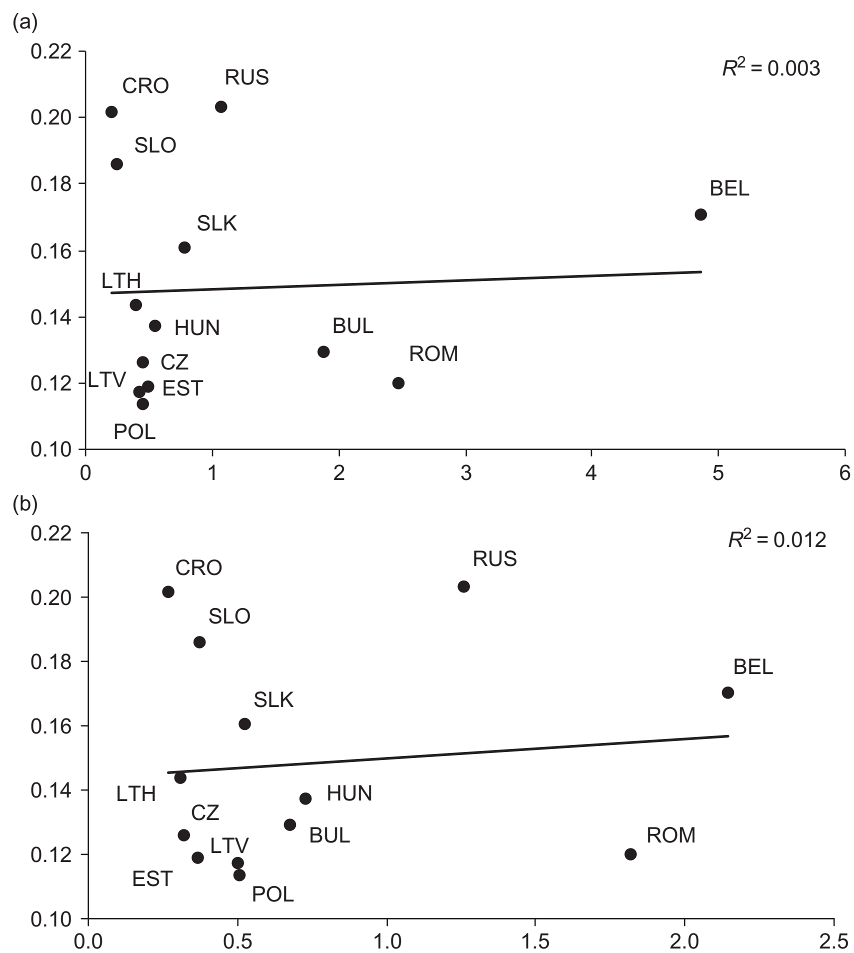

Fig. 1. (a) Variance and (b) average monthly inflation versus price change probability
Note: BEL, Belarus; BUL, Bulgaria; CRO, Croatia; CZ, Czech Republic; EST, Estonia; HUN, Hungary; LTV, Latvia; LTH, Lithuania; POL, Poland; ROM, Romania; RUS, Russia; SLO, Slovenia; SLK, Slovak Republic.

Equation 1 limits the minimum price change frequency to monthly interval. The assumed frequency could be inadequate for the periods of high inflation, so I estimate the instantaneous probability of price change (Bils and Klenow, 2004):

$$P = -\ln(1-\delta) \tag{6}$$

The corresponding mean time between price changes is calculated as $T_P = 1/P$. The estimates are summarized in the lower panel of Table 2. No additional insight is provided by allowing more frequent changes. The estimated average probability for the whole group is 1.3 percentage points greater and the mean time is approximately 2 weeks shorter.

Seasonality in price changes

The results reported in section 'Estimation results for the seasonally unadjusted series' were obtained by running the AR(1) model for the seasonally unadjusted price series. To capture the influence of seasonality, I ran the AR(1) model with seasonal dummies. Since most of the coefficients were insignificant, I re-estimated the model only for significant monthly dummies. Table 3 reports the coefficients, probabilities (in percentage) and the mean time.

After including the seasonal dummies, the mean probability of a price change for the group of 13 countries is 2.5 percentage points smaller than the probability reported in section 'Estimation results for the seasonally unadjusted series'. The corresponding mean time between price changes is 8.9 months. For Belarus, Bulgaria and Romania, however, the estimated values of δ are slightly larger. The largest impacts of seasonality on price change are found in the Czech Republic, Latvia and Poland.

The last row of Table 3 lists the mean value of the month coefficients for countries in which the coefficient was found to be significant. Not surprisingly, January is the month that influences the price changes most. The coefficient is significant and positive in 12 out of 13 countries. Centrally regulated prices along with the 'menus' supplied by the foreign companies are the most probable factors that affected the price changes in January. The values of coefficients obtained for June–August are negative, which might be connected to the decreasing food prices.

IV. Price Quotation

Once the residuals from the AR(1) model are obtained, it is possible to take a closer look at the price quotation series. The quotations are calculated in accordance to Equation 3. Then, using the finite sum $\sum_{j=0}^{k}$ with an arbitrary chosen lag-length k, these series are employed to check whether Equation 1 fits the actual data. Figure 2 depicts the results for Croatia and Poland.[1] The simulated price series (solid lines) fit the actual data reasonably well. They mimic not only the dynamics, but also the magnitude of the trend-adjusted series. Finally, the artificial price series obtained by the discounted sum of ν_t are much closer to the actual series than the simulations based on ε_t. These experiments, however, provide little evidence for a forward-looking pricing mechanism.

Pricing in the discrete time version of the Calvo model consists of two equations. Besides Equation 1, the mechanism is described by the price quotation:

$$\nu_t = \delta \sum_{j=0}^{\infty} (1-\delta)^j E_t\left(p_{t+j} + \lambda \hat{y}_{t+j}\right) \tag{7}$$

where $\hat{y}$ stands for the output gap. Solving the system of Equations 1 and 7 for p_t, one obtains a solution

$$p_t = p_{t-1} + \frac{\delta^2 \lambda}{1-\delta} E_t\left(\sum_{t=0}^{\infty} \hat{y}_{t+i}\right) \tag{8}$$

which is asymptotically unstable. (This justifies the use of HP filter.) Now it is important to note that Equation 3 might be interpreted as the Wold decomposition of the price series. A univariate stationary time-series process can be decomposed into a deterministic and independent Gaussian process (Bierens, 1994). In this case, the independent process is ν_t. Estimating Equation 3 is similar to the first step of the Granger-causality test procedure presented by Sargent (1976). Checking whether the discounted sum of the (future) output gap significantly influences the price quotations, as Equation 8 suggests, is therefore equivalent to testing for Granger causality from the output gap to prices.[2]

I start with regressing the ν_t series against the actual lead values of output gap

[1] I concentrate on the countries with short and long mean time T. Other figures and series are available upon request. The value of k was set to 5.

[2] Rudd and Whelan (2005) presented similar conclusion for inflation dynamics.

Table 3. Model with seasonal dummies

Country	Months												δ	T
	Jan	Feb	Mar	Apr	May	Jun	Jul	Aug	Sep	Oct	Nov	Dec		
Belarus	0.014	0.007						−0.012			0.007	0.011	17.5	5.7
Bulgaria	0.011				−0.008	−0.017			0.009	0.003		0.004	13.7	7.3
Croatia	0.004				0.004	−0.003	−0.005						14.2	7.0
Czech Republic	0.013						0.006						8.3	12.1
Estonia	0.005			0.003				−0.004					9.8	10.2
Hungary	0.009	0.005	0.003		0.003		−0.004	−0.007			−0.003	−0.004	9.5	10.5
Latvia	0.008						−0.005	−0.007		0.003			7.0	14.4
Lithuania	0.005							−0.004					14.3	7.0
Poland	0.006			0.003			−0.006	−0.006	0.003				8.2	12.2
Romania							−0.005	−0.007					12.3	8.1
Russia	0.012	0.005						−0.008	−0.005				15.9	6.3
Slovak Republic	0.016	0.004											10.9	9.2
Slovenia			0.004		0.003			−0.004					18.4	5.4
Mean value	0.009	0.005	0.003	0.003	0.000	−0.010	−0.003	−0.007	0.002	0.003	0.002	0.004	12.3	8.9

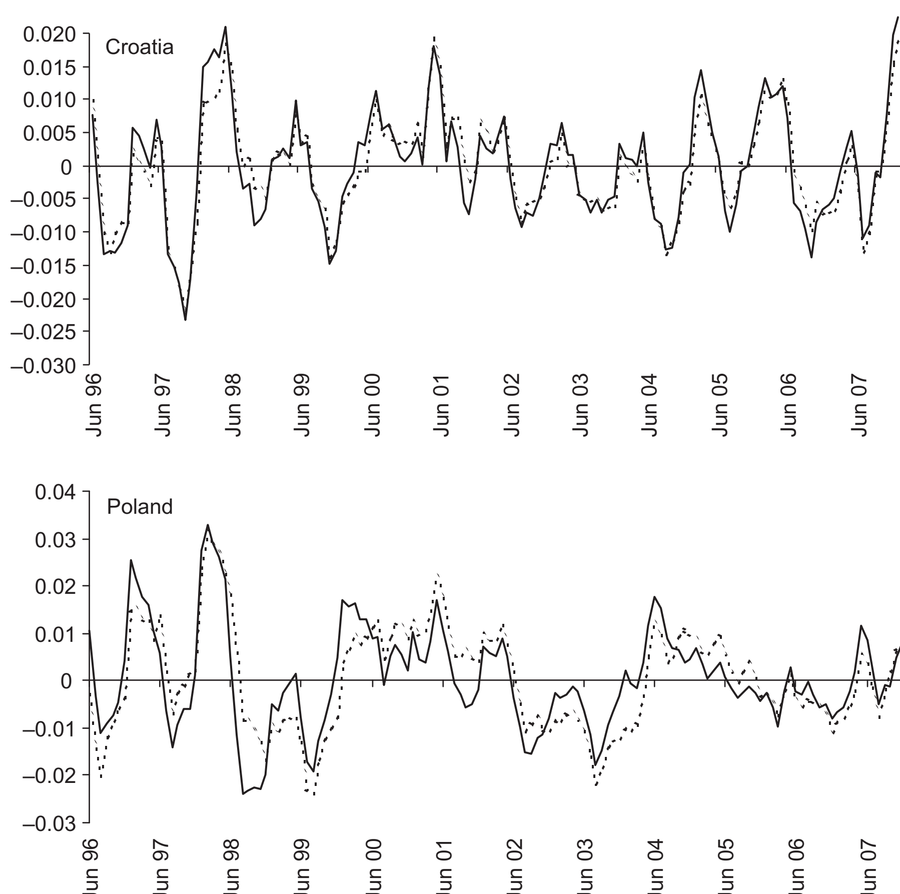

Fig. 2. Actual and simulated price series for Croatia and Poland

$$\nu_t = \sum_{i=0}^{l} \theta_i \hat{y}_{t+i} \tag{9}$$

and test the null hypothesis H_0: $\theta_1 = \theta_2 = ... = \theta_l = 0$, where l is chosen subject to the minimization of the Schwarz Criterion. Then I carry out an impulse response analysis. I estimate the SEs for the response of price quotations Φ_j to a one-unit shock in the lead output gap as the square root of the first diagonal element of the product:

$$G_j \sum_{\hat{\theta}}^{NW} G_j' \tag{10}$$

where[3]

$$G_j = JG_{j-1}\Theta' + J \otimes \Phi_{j-1}, \quad J = I_{(l)},$$
$$\Theta = \begin{bmatrix} \theta_1 & \theta_1 & ... & \theta_{l-1} & \theta_l \\ 1 & 0 & ... & 0 & 0 \\ 0 & 1 & ... & 0 & 0 \\ \vdots & & \ddots & & \vdots \\ 0 & 0 & ... & 1 & 0 \end{bmatrix}$$

Using the estimated responses and the corresponding SEs, I check the overall significance of price quotation response. The probabilities for the first $(K-1)$ individual responses[4] are reported in the first columns of Table 4. The last two columns of Table 4 show the sum of the coefficients and the p-values of the Wald test that the sum is equal to 0. Figure 3 depicts the plotted values of responses along with the asymptotic two-SE bounds (dotted lines). To answer the question whether the Calvo-like pricing behaviour describes the situation in Central Europe properly, I use the results of both tests. The answer would be positive if the null of no response of price quotation to output gap is rejected and the sum of θ-coefficients is positive and significant.

Regarding the forward-looking pricing behaviour, the CEE countries form three groups. The results obtained for the first cluster (Belarus, Bulgaria, Estonia, Russia, the Slovak Republic and Slovenia) suggest that the forward-looking Calvo-like pricing mechanism does not explain the price dynamics. The agents in Croatia, the Czech Republic, Hungary and Lithuania set the prices in accordance to the forward-looking mechanism. For Croatia and the Czech Republic, both hypotheses are decisively rejected. For a slightly less restrictive level of significance, the null can also be rejected for Lithuania and Hungary. The third group consists of countries, for which the responses were found to be significant, but the sums of coefficients were not. This is the case of Latvia, Poland and Romania.

V. Summary

Recent empirical works provide some evidence regarding price change frequencies that is not

Table 4. Impulse response analysis

	Probabilities for first l responses										
Country	1	2	3	4	5	6	7	8	9	Sum of coefficients	Probability
Belarus	0.725									0.013	
Bulgaria	0.766									0.006	
Croatia	0.05	0.032	0.171	0.02	0.308	0.000				0.053	0.045
Czech Republic	0.592	0.005	0.928	0.001	0.001	0.006	0.098	0.018		0.059	0.013
Estonia	0.596	0.601								–0.001	0.883
Hungary	0.747	0.708	0.000	0.205	0.226	0.446	0.168	0.000	0.000	0.038	0.097
Latvia	0.047	0.286	0.000							–0.004	0.647
Lithuania	0.215	0.159	0.384	0.651	0.044	0.011				0.004	0.016
Poland	0.495	0.371	0.031	0.049	0.205	0.000				0.023	0.243
Romania	0.264	0.003								–0.022	0.319
Russia	0.000	0.836	0.000							–0.020	0.000
Slovak Republic	0.523									–0.008	
Slovenia	0.922									–0.002	

[3] For G_1, see Lütkepohl (2005). Note, however, that Equation 10 differs slightly from the formula presented by Lütkepohl.
[4] Lütkepohl (2005) showed that in a VAR(1) model for K variables the number of individual responses to be checked for significance is equal to $l(K-1)$. Since in this case $K-1 = 1$, the number of lags l is equal to the number of individual responses.

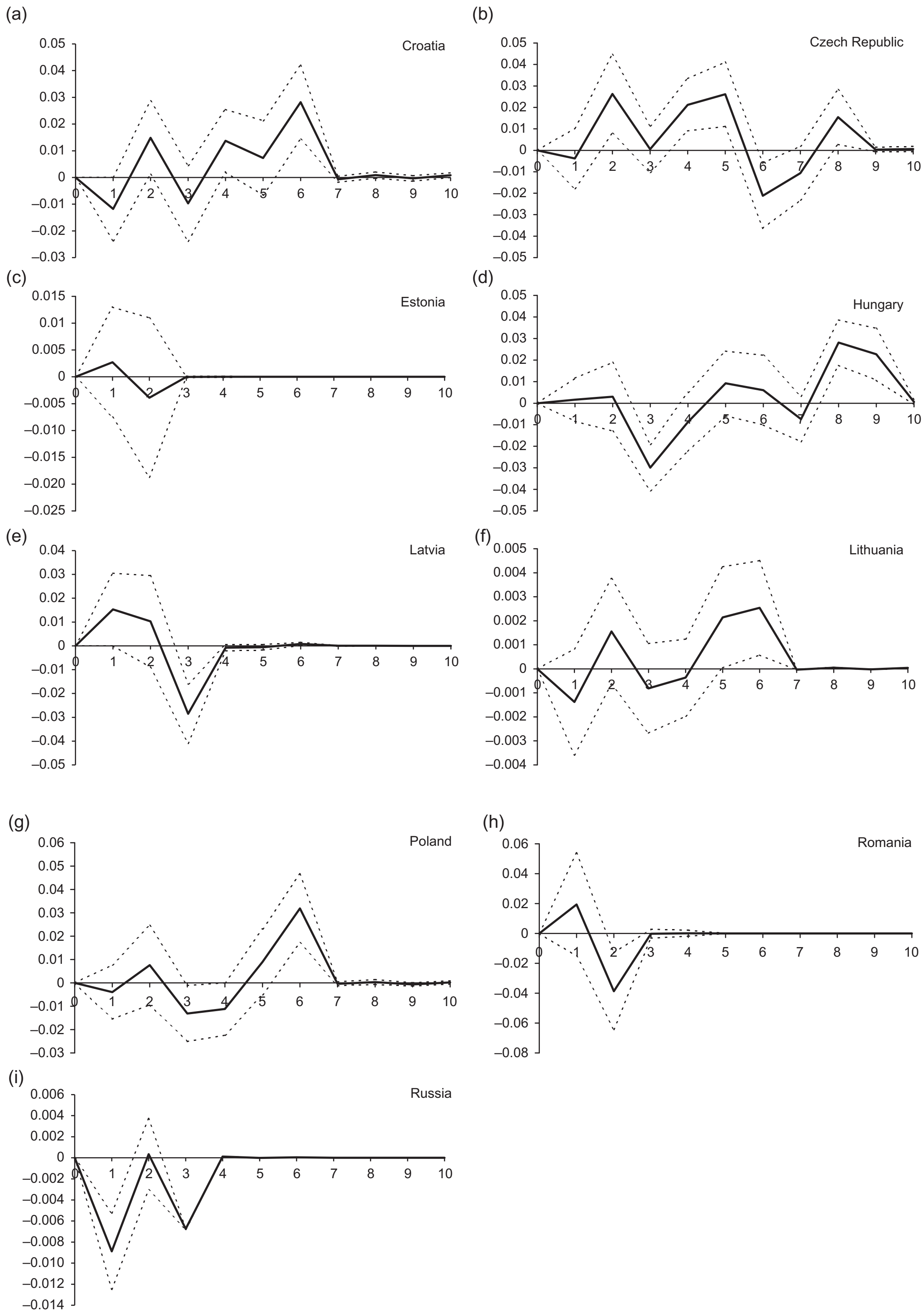

Fig. 3. Impulse response: (a) Croatia, (b) Czech Republic, (c) Estonia, (d) Hungary, (e) Latvia, (f) Lithuania, (g) Poland, (h) Romania and (i) Russia

consistent with the NKPC literature. Unlike the latter, however, they concentrate on micro data. Thus, the discrepancies between the microeconometric results and the NKPC might be explained by the (dis)aggregation level. In this article, I employed the aggregate CPI series, but the results were much closer to the ones obtained by the micro studies.

For the group of 13 European transition countries, the mean probability was 14.8%, while the mean time between the price changes was slightly longer than 6 months. The results were clearly not in tact with the NKPC literature, which usually considers prices to be rigid for more than 1 year. The impact of seasonality was rather small. The estimated average probability of a price change after the inclusion of seasonal dummies decreased by approximately 2.5 percentage points, and the corresponding time between price changes was 2.5 months longer. The sign and significance of the distributed lead coefficients along with the impulse response analysis indicated that the Calvo-like pricing mechanism is suitable to describe agents' behaviour only in Croatia, the Czech Republic, Hungary and Lithuania.

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