Changes in Inflation Dynamics under Inflation Targeting? Evidence from Central European Countries

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1 Introduction

Understanding the nature of short-term inflation dynamics poses a major challenge for monetary policy. Sound knowledge of inflation properties is especially pressing for countries whose economies underwent dramatic structural changes and where institutional setting of monetary policy considerably changed in order to engineer sharp disinflation process. While taming inflation was traditionally considered costly in terms of lost output, a better notion of the role of expectations equipped policy makers with a hope that credible monetary policy can achieve disinflation without causing detrimental effect on real economic activity. This concept has become a hallmark of the New Keynesian Phillips curve (NKPC).

The NKPC was designed as a structural model which is based on optimization process at the micro level and thus should be invariant to policy changes. However, this claim is not fully supported by recent experience. There are numerous examples why parameters of the NKPC model can be susceptible to variation. There has been vast evidence about changes in monetary policy setting and its impact on inflation for developed countries (Baxa et al., 2010; Boivin, 2006; Kim and Nelson, 2006; Koop et al., 2009; Sims and Zha, 2006; Trecroci and Vassalli, 2010). A more aggressive monetary policy stance (Davig and Doh, 2008) or implementation of credible monetary policy regimes such as inflation targeting (Benati, 2008) have been considered key drivers in decrease of inflation persistence. On the contrary, the cross-country panel studies such as Ball and Sheridan (2004), Mishkin and Schmidt-Hebbel (2007) or Brito and Bystedt (2010) find rather mixed evidence on the relative performance of inflation targeters vs. non-targeters both in developed and emerging countries.

The macroeconomic developments are inseparable from changes on microeconomic level. When a change in macroeconomic setting occurs, agents' behavior might gradually adapt to new conditions. In particular, recent evidence suggests that firms' decisions on the frequency of price adjustment are prone to be state-dependent. As corroborated by most of microeconomic studies on price-setting, the price adjustment is mainly influenced by the level and variability of inflation (see Klenow and Malin, 2010, for a survey). Similarly, Fernandez-Villaverde and Rubio-Ramirez (2008) show within the DSGE framework that movements of pricing parameters are indeed correlated with inflation. Therefore, there is no a priori reason why either reduced-form or 'structural' parameters of the NKPC shall be time invariant.

This also seems to hold from purely modeling perspective. Estrella and Fuhrer (2003) and others offer a suite of methodological explanations why structural models derived from agents' optimizing behavior and based on the assumption of rational expectations do not guarantee immunity to the Lucas critique. Therefore, the stability of structural model and its ability to withstand the Lucas critique should not be an *a priori* assumption but rather a hypothesis subject to empirical testing. Arguably, the need for its verification is stronger in transition countries than anywhere else.

Countries in Central Europe (CE) represent a unique sample for analysis of inflation dynamics within a context of changes in economic system and monetary policy conduct. They jointly underwent economic transition to market economy, which could have induced similar changes both in price setting and expectation formation behavior. The three CE countries, the Czech Republic, Hungary, and Poland seem to be particularly suitable for cross-country analysis. They are relatively similar small open economies with strong regional and historical affinity and all introduced inflation targeting as disinflation strategy. On the other hand, their actual monetary policy conduct shows notable differences, e.g. in the role given to exchange rate. Therefore, each national central bank faced distinct challenges when building the credibility of the regime. Under these conditions, one can run a natural experiment to assess the effectiveness of different monetary policy settings vis-à-vis changes in inflation process. The lesson learned from these countries is of general interest given that inflation targeting is still preferred monetary framework being adopted by emerging countries around the globe. It should be also stressed that unlike many other emerging and transition economies, the OECD and EU membership of CE countries makes the data reliable and internationally comparable.

Alongside the results of the cross-country analysis which brings some new insights and offers a lesson for other emerging countries, we make two other contributions to the existing literature. First, we extend the open-economy version of the NKPC proposed in Galí and Monacelli (2005) into hybrid form and time-varying context and show what one can learn about possible changes in inflation dynamics under inflation targeting regime. We track evolution of both the reduced form and the 'structural' parameters. Second, we introduce some methodological improvements in estimation of the NKPC. In our two-step procedure closely related to Kim (2006) we estimate a time-varying regression model with stochastic volatility using Bayesian techniques. As documented by several authors (see e.g. Koop and Korobilis, 2009), changes in inflation volatility commonly occur and might induce spurious variantion in estimated coefficients when not treated appropriately. In addition, we propose to use Bayesian model averaging method to tackle the thorny issue of instrument selection. Sensitivity of results to the choice of the conditioning instrument set was shown to be very relevant in forward-looking models, with the NKPC being a prominent example in this context (see e.g. Mavroeidis, 2005).

Results can be summarized as follows. First and foremost, structural changes in economy in tandem with shifts in monetary regimes indeed produced some changes in inflation process, although the reduced-form parameters of the NKPC are generally quite stable. In any case, one can draw these conclusions only when structural instability, including volatility of inflation shocks, is ex-ante accounted for. Second, we find that the overall nature of the inflation process alike its time evolution differs across the selected CE countries. Intrinsic inflation persistence has dropped substantially only in the Czech Republic to currently insignificant levels. The volatility of inflation shocks decreased quickly a few years after the adoption of inflation targeting in both the Czech Republic and Poland. By contrast, the nature of the inflation process in Hungary does not seem to have changed much over the last 15 years in spite of the fact that country adopted inflation targeting a decade ago. Third, the inflation-output trade-off seems to be blurred by potentially important supply shocks during the transition, which cannot be fully captured by the original NKPC model. However, results tend to reveal non-linear relationship between domestic economic activity and inflation. Indeed, the coefficient of the output gap consistently increases only in periods when the output is significantly away from its potential. Similar conclusions can be drawn for foreign inflation factors, tracked by the terms of trade. They turn significant in specific periods such as major exchange rate devaluations. Yet, there is also some indication suggesting that the foreign factors might already be well reflected in inflation expectations themselves. Finally, the overall changes in inflation process reflect the changes in price-setting behavior of firms that we detect both in cross-country and temporal dimension.

Our empirical findings have some noteworthy policy implications for (emerging) countries that adopt inflation targeting. Although all three CE countries officially adopted inflation targeting more than a decade ago, intrinsic inflation persistence has not decreased considerably in two of them and remains at higher levels than in developed countries. Therefore, it seems that adoption of inflation targeting itself does not automatically produce any changes in inflation process and the specific way the framework is implemented might matter. Indeed, if inflation targeting is not sufficiently credible, the economic agents might chiefly take into account observed inflation levels rather than the inflation target. This can be arguably related to the role of the exchange rate in the monetary policy. Although inflation targeting shall aim at domestic price level, the Hungarian monetary policy has paid special attention to exchange rate and the exchange rate channel was considered the most efficient channel of monetary policy transmission (Vonnak, 2008).

The paper is organized as follows. In section 2, we review relevant literature. Section 3 presents our empirical framework and data. All the results and their interpretation appear in section 4. The final section concludes and suggests some avenues for future research.

2 Related Literature

From the empirical perspective, the NKPC owes its growing popularity to the seminal papers of Galí and Gertler (1999) (GG hereafter) and Galí, Gertler, and López-Salido (2001, GGL). Despite the theoretical appeal of the (hybrid) NKPC, consecutive studies have produced rather conflicting empirical evidence with results varying across economies, data sets, and – most notably – estimation methods (e.g. Rudd and Whelan, 2005; Mavroeidis, 2005; Lindé, 2005). To account for characteristics of small open economies Galí and Monacelli (2005) derive a small open economy version of the NKPC for CPI inflation, which in addition to marginal cost includes also the terms of trade . Mihailov et al. (2011a) provide some favorable empirical evidence on this model based on data of selected OECD countries. With respect to theoretical underpinnings, it is probably the closest empirical study to our own. Alternative approaches include Batini et al. (2005), who propose an open-economy NKPC where the marginal cost is affected by import prices and external competition, confirming that this model fits the UK data well. Rumler (2007) extends the marginal cost measure to include the cost of intermediate inputs (both domestic and imported) and finds some plausible evidence for the euro area countries.

A few recent studies consider the effects of structural changes in the economic system and monetary policy regime and explore how these are propagated into the parameters of the NKPC. Most of the evidence is available for the US. The intrinsic inflation persistence was found to be an empirical artefact driven by specification bias inherent to fixed-coefficient models (Hall et al., 2009) or to variation in the long-run inflation trend (Cogley and Sbordone, 2008). Several authors also document that the nature of inflation process changes with macroeconomic environment (Zhang and Kim, 2008) and monetary policy regimes (Cogley et al., 2010; Kang et al., 2009). However, there are also studies (Stock and Watson, 2007) claiming that the US inflation persistence has not changed for decades.

The evidence on changes in inflation dynamics in other economies is less abundant.[1] Benati (2008) examines data for several developed inflation-targeting countries (Canada, New Zealand, Sweden, Switzerland, and the UK) and the euro area concluding that inflation persistence decreased almost to zero once credible monetary regimes had been implemented and, therefore, that inflation persistence is not structural. Hondroyiannis et al. (2009) apply a specific time-varying framework to data for France, Germany, Italy, and the UK, concluding consistently with previous evidence for the US (Hall et al., 2009) that the backward-looking parameter of the time-varying NKPC is almost negligible. Tillmann (2009) explores how the explanatory power of the forward-looking NKPC in the euro area evolves across time finding that the explanatory power of the model varies substantially across the underlying monetary regimes and is influenced by events such as the ERM crisis, the Maastricht treaty, and the launch of EMU. Koop and Onorante (2011) use dynamic model averaging (Raftery et al., 2010) to study the relationship between inflation and inflation expectations in the euro area. They find strong support for forward-looking behavior, interestingly mainly since the start of the recent financial crisis.

The changes in inflation dynamics in transition countries have not been paid much attention to so far. The existing time-invariant estimates of the NKPC for countries in question mostly imply that inflation is relatively more persistent than in advanced economies and that external factors seem to matter (Franta et al., 2007; Vašíček, 2011; Mihailov et al., 2011b). Based on this evidence, however, it is very difficult to draw any conclusion related to the monetary policy effects on the temporal and cross-country variation in inflation process. Only Hondroyiannis et al. (2008) provide some evidence based on a time-varying model for panel of seven new EU member states. The panel estimation for (in our view) a highly heterogeneous group of seven new members does not seem to be appropriate given that the economic structures and monetary policy frameworks of these countries are very different. They find that the inflation persistence in these countries is practically nonexistent, which contradicts practically all the country-specific time-invariant evidence.

Our study is conceptually related to papers tracking structural changes in inflation process by means of time-varying estimation of the NKPC such as Cogley et al. (2010); Hall et al. (2009); Hondroyiannis et al. (2008, 2009). Nevertheless, those works aim almost exclusively at major developed countries. We share the focus on small open economies with Mihailov et al. (2011a,b), but they do not in turn consider possibility of structural changes, which might be relevant for emerging countries in general, and transition countries in particular. Our study also bears some resemblance to studies using alternative empirical frameworks, which aim specifically on the changes in inflation persistence along changes in monetary policy regimes (Benati, 2008; Kang et al., 2009). Finally, our aim to shed some light on the viability of inflation targeting regime brings us spiritually close to cross-country (panel) studies (e.g. Mishkin and Schmidt-Hebbel, 2007; Brito and Bystedt, 2010). Yet, the approach of these studies is of aggregate and unstructural nature. On the contrary, our approach allows revealing that effects of inflation targeting regime can be different even in relatively similar countries and therefore structural country-specific analysis might be more appropriate than aggregate estimation with heterogenous country panels.

[1]There are numerous studies initiated by the ESCB Inflation Persistence Network but they are mainly based on micro data.

3 Model and Estimation Strategy

3.1 Open-economy Hybrid NKPC

We start our exposition with the seminal hybrid NKPC model laid out in GG:

$$\pi_t = \gamma_f E_t \pi_{t+1} + \gamma_b \pi_{t-1} + \lambda s_t + \varepsilon_t, \tag{1}$$

where π_t denotes inflation, $E_t \pi_{t+1}$ represents inflation expectations conditional on the information up to time t, s_t is a proxy for the marginal cost (as a deviation from the steady-state), and ε_t is an exogenous inflation shock, such that $E_{t-1}\varepsilon_t = 0$. Unlike GG, we assume that parameters γ_f, γ_b, and λ are potentially time-varying, i.e., they may evolve over time because of the dynamic economic conditions in the converging economies under study. We provide more detailed motivation and justification for time-varying model in the next subsection.

Inflation persistence enters equation (1) not only through backward-looking term γ_b but potentially also through parameter λ as long as markets that determine evolution of the forcing variable (output gap, in our case) are rigid. Moreover, inflation process can exhibit rather persistent properties (in terms of autocorrelation) even under quite stable economic conditions where lagged inflation provides a good guess about the future inflation path.

From monetary policy perspective, we are however chiefly interested in the intrinsic (or structural) price rigidity tracked by the backward-looking term, insomuch as it captures persistence inherent to the inflation process itself. It contributes to lower ability of monetary authorities to make disinflation policy costless and, to a certain extent, implies their limited credibility. High instrinsic persistence can also signal poorly anchored inflation expectations.

The reduced-form parameters are non-linear functions of three structural parameters: a subjective discount factor, β, the probability that prices remain fixed, θ, and a fraction of backward-looking price setters, ω.

$$\begin{aligned}
\lambda &\equiv (1-\omega)(1-\theta)(1-\beta\theta)\phi^{-1} \\
\gamma_f &\equiv \beta\theta\phi^{-1} \\
\gamma_b &\equiv \omega\phi^{-1} \\
\phi &\equiv \theta + \omega(1-\theta(1-\beta))
\end{aligned}$$

The structural parameters may provide a closer view of the nature of the structural changes that have been affecting the economies in question. Specifically, one might be interested in finding out whether the fraction of backward-looking setters has decreased, for example, as a result of the inflation-targeting regime, or how the average time for which prices remain fixed $(1/(1-\theta))$ drifts over time.

Given that the CE countries under study are all small open economies, we derive a new version of hybrid NKPC model in the spirit of Galí and Monacelli (2005), accounting for the potential impact of external factors on inflation. Recently, Mihailov et al. (2011b) used the pure small-economy NKPC model of Galí and Monacelli (2005) and evaluated the relative importance of domestic and external drivers in the new member states. Our version of the model can be viewed as an extension of their approach to the hybrid form and time-varying framework.

In line with Galí and Monacelli (2005), we now assume that CPI inflation can be expressed as:

$$\pi_t = \pi_{H,t} + \alpha \Delta TT_t, \tag{2}$$

where $\pi_{H,t}$ is *domestic inflation*, ΔTT_t denotes the current-to-past period change in the terms of trade[2] and parameter α measures openness of the economy. Analogously to (1), the dynamics of domestic inflation are given by:[3]

$$\pi_{H,t} = \gamma_f E_t \pi_{H,t+1} + \gamma_b \pi_{H,t-1} + \lambda s_t. \tag{3}$$

Plugging (3) into (2) and making use of the fact that $\pi_{H,t} = \pi_t - \alpha \Delta TT_t$, we get:

$$\pi_t = \gamma_f E_t(\pi_{t+1} - \alpha \Delta TT_{t+1}) + \gamma_b(\pi_{t-1} - \alpha \Delta TT_{t-1}) + \lambda s_t + \alpha \Delta TT_t,$$

and after some rearrangement we obtain a hybrid open-economy NKPC model of the form:

$$\pi_t = \gamma_f E_t \pi_{t+1} + \gamma_b \pi_{t-1} + \lambda s_t + \alpha\{\Delta TT_t - \gamma_f E_t \Delta TT_{t+1} - \gamma_b \Delta TT_{t-1}\}. \tag{4}$$

To motivate economic interpretation of the last term[4] in (4), it is useful first to consider the two extreme cases when either γ_f or γ_b is equal to one.[5] If $\gamma_f = 1$, then the term becomes $(\Delta TT_t - E_t \Delta TT_{t+1})$ and model (4) collapses into the pure open-economy model introduced by Mihailov et al. (2011b). Intuitively, as pointed out by Mihailov et al. (2011b), current demand for domestic goods in the pure NKPC would increase when $\Delta TT_t > E_t \Delta TT_{t+1}$ because the relative price of domestic goods is lower than that anticipated in the future, and this increased demand causes upward pressure on current inflation. Conversely, when $\Delta TT_t < E_t \Delta TT_{t+1}$, current-period demand for domestic goods would fall, as agents expect their relative price to decline in the future, and this exerts downward pressure on current inflation.

In a fully backward-looking setting, implied by $\gamma_b = 1$, the term shrinks to $(\Delta TT_t - \Delta TT_{t-1})$. Again, the effect on inflation can be inferred by comparing the two terms in brackets, i.e., by investigating whether $\Delta TT_t > \Delta TT_{t-1}$ or $\Delta TT_t < \Delta TT_{t-1}$ holds true. The crucial difference, however, is that backward-looking agents now anticipate the future path of the terms of trade with respect to their past value, since the lagged value is used as a simple way to make a forecast. Note that this implies, other things being equal, higher inflation inertia than in the closed-economy model, because the terms of trade now serve as another channel contributing to persistence.

When the universe is formed by both forward and backward-looking agents, one simply compares ΔTT_t to the linear combination of $E_t \Delta TT_{t+1}$ and ΔTT_{t-1}, where coefficients γ_f and γ_b serve as multiplicative constants or weights. Hence, the linear combination[6] can be viewed as a weighted average of the next-to-current difference in the terms of trade anticipated by forward-looking and backward-looking agents. Since a difference in the terms of trade is nothing else but a change in the relative prices of imports (in terms of exports), it can, in a certain respect, be interpreted as a measure of import *inflation.* Thus, the hybrid open-economy NKPC consistently uses the same hybrid formation for inflation expectations no matter whether they are defined as a rise in the general price level of goods and services or as the relative price of imports in terms of exports.

[2] Galí and Monacelli (2005) use an inverse definition of the terms of trade, i.e., they define it as the import price index over the export price index.

[3] We leave out the error term for expositional ease.

[4] To spare space, we will simply refer to this term as TT_t in the subsequent sections.

[5] Although we do not impose the restriction $\gamma_f + \gamma_b = 1$ a priori, the results usually show close-to-convexity properties.

[6] If we restrict the coefficients to sum to 1, we obtain a convex combination with the straight interpretation of a weighted average.

3.2 Econometric Framework

Model (4) cannot be estimated directly due to fact that $E_t\pi_{t+1}$ is, in essence, a latent quantity which must be proxied by some observable variable. Since inflation expectations taken from surveys cover only a very short time span, we proceed by making a common assumption that economic agents form their expectations rationally and replace $E_t\pi_{t+1}$ by π_{t+1}. Note, however, that this leads to endogeneity bias, as future inflation is by construction correlated with the error term. To see this, let $\vartheta_{t+1} \equiv \pi_{t+1} - E_t\pi_{t+1}$ be the unpredictable forecast error and rewrite (4) into the following form:

$$\pi_t = \gamma_f \pi_{t+1} + \gamma_b \pi_{t-1} + \lambda s_t + \alpha TT_t + e_t, \qquad (e_t \equiv \varepsilon_t - \gamma_f \vartheta_{t+1}). \tag{5}$$

To obtain time-invariant parameter estimates in model (5) one routinely resorts to GMM techniques and, not surprisingly, this has also been common practice in previous research focusing on CE countries (Franta et al., 2007; Vašíček, 2011; Mihailov et al., 2011b). Still, within the context of economic changes presented above and with regard to the goal of our analysis, a pursuit of this strategy might not be the best option available. Even though a model with fixed parameters can be quite good approximation of the true state on average, it does not allow for closer inspection of changes, which are potentially related to the shifts in monetary regimes and/or to adaptation mechanisms initiated by transition. As such, a use of time-varying coefficient model is clearly preferable in our setting. On the other hand, economic intuition is probably not enough to recklessly adopt time-varying framework and some statistical motivation is also necessary.

Looking at the GMM results for different time-spans presented in Table B.1 (see Appendix), one can generally observe a clear tendency to instability in coefficients. This is also confirmed by formal GMM breakpoint tests in Table B.2 (Andrews-Fair Wald and LR-type tests), which allow not only for changes in the coefficients but also in the variance covariance matrix of the error terms.[7] Yet, in light of specific economic reasons discussed above, we tend to believe that gradual smooth changes are a priori more plausible framework than single (or multiple) structural breaks.[8]

For this reason, we also checked parameter instability via Flexible least squares (FLS, Kalaba and Tesfatsion, 1989).[9] Although FLS are not a formal testing procedure; they provide handy descriptive tools for analyzing variation in parameters. Similarly to provided structural stability tests, FLS results (see Figure B.1 in Appendix) suggest that there exists systematic parameter instability in all three countries, which is the most pronounced in the Czech Republic and quite less so in Poland and Hungary. At the same time, coefficients' path for alternative values of the smoothing parameter showed that gradual changes are more plausible framework than a presence of single break.

[7] Given the sample size, it seems reasonable to test breakpoint around middle of the sample so that coefficients can be reliably estimated in each subsample. Consequently, we assumed breakpoint in 1Q 2003, albeit alternative breaks in earlier or later dates provide very similar results. This break point is also reasonable given the evidence that inflation targeting might affect agents' behavior around two or three years after the implementation.

[8] Recently, Chen and Hong (2012) proposed two tests with power against one-off as well as gradual changes. They are based on the mutual comparison of fixed and time-varying coefficient model with respect to squared residuals and model's fit. Even though the tests are not directly useful to our needs as they still cannot effectively determine the nature of time-varyingness, the relation between variability in coefficients and model's capacity to explain the data deserves some deeper exploration.

[9] The objective function of FLS consists of two subcriteria, namely: goodness-of-fit determined by squared residuals (measurement error, R^2_M) and the smoothness of coefficients given by the sum of their squared first differences (dynamic error, R^2_D). Relative weight assigned to each criterion is regulated through smoothing constant μ. One of the useful outputs of FLS is the residual efficiency frontier, which consist of all pairs of (weighted) measurement and dynamic error that are compatible with the minimum value of the objective function. It allows to check whether small changes in variability of coefficients lead to a substantial improvement in measurement error.

Against this backdrop, we broadly stick in our core empirical framework to the strategy proposed by Kim (2006), who tackles the issue of endogeneity in linear models with dynamic coefficients following a random walk. In principle, Kim (2006) shows that it is possible to get consistent estimates of time-varying coefficients by employing a two-step procedure. In the first step, one runs the OLS regression[10] of endogenous variables on a set of instruments that are uncorrelated with the error term in (5) and stores the standardized residuals. In the second step, the standardized residuals are added as additional regressors into (5) and the whole system with time-varying coefficients can be cast into the state-space form and estimated with a few modifications using the Kalman Filter in a quite traditional fashion. Details on modifications in Kalman filter formulas are given in Kim (2006) and Kim (2008).

Despite the practical appeal of the two-step procedure, there are still some thorny issues to be answered in the NKPC context: *i*) economic theory does not postulate what instruments should be used in the first step, which leads to the common problem of instrument selection, *ii*) the standard estimation of linear state-space models using the Kalman filter assumes that Gaussian shocks to the target variable are constant over time. However, this is unlikely to hold for the inflation process (as argued for example in Koop and Korobilis, 2009). Applying methods that ignore possible variation in the volatility of the error term may lead to serious bias of the estimated time-varying coefficients.[11]

To address both these issues we slightly modify original procedure of Kim (2006). More specifically, we use Bayesian model averaging (BMA) instead of traditional OLS in the first step and estimate a time-varying model with stochastic volatility[12] in the second. Bayesian model averaging (see Hoeting et al., 1999) is a relatively new method that was introduced to a wider audience in the mid-1990s. It provides a coherent framework to account for model uncertainty and instrument sensitivity. Unlike the 'traditional' approach to estimation of the NKPC, where a researcher typically selects instruments (and thus conditions her model) in a quite subjective manner, BMA effectively weights all the possible models based on the posterior model probability. Thus, the aim of model averaging is not to find the best model or to select the best possible set of instruments, but rather to use information from all models and average the outcome with respect to their 'reliability', induced by data and priors. To our knowledge the BMA approach is new in the NKPC literature, although similar ideas have already been tossed around in the context of rational expectations models (see Wright, 2003).

The need to apply a more formal approach to instrument selection is again demonstrated in Table B.1 where we used GMM with some instrument sets typically exploited in the literature. One can clearly observe notable variation in estimated parameters across different instrument sets, which mirrors the considerable uncertainty related to their choice. Under this situation, BMA stands for natural and effective approach capable of solving problems of this kind. As again demonstrated in Table B.1, the selection of instruments by means of BMA leads to different results than other instrument sets, sometimes changing coefficients' statistical significance.[13]

[10] Kim (2008) considers all possible alternatives on the relation between endogenous variable and instruments. In particular, in line with Kim (2006) one option is to assume time-varying relation between endogenous variables and instruments. For reasons that will become clear later we do not adopt this approach here.

[11] Indeed, we find that stochastic volatility does matter and its omission leads to highly erratic and unstable results. Due to their apparent failure, we do not present these results here, but they are available upon request.

[12] Although Kim (2006) procedure in principle allows dealing with heteroskedasticity by means of GARCH, the prevailing view in the literature is that stochastic volatility models are more flexible and usually outperform the ARCH-type models (Kim et al., 1998).

[13] Sensitivity of the parameter estimates to the choice of instruments for the case of the Czech republic is more

Let Z be a $T \times k$ matrix summarizing the information set available to economic agents. Under standard assumptions, the unrestricted model can be represented as:

$$y_t = a + Z_t \delta + \epsilon_t \qquad \epsilon \sim N(0, \sigma^2), \tag{6}$$

where y_t denotes the outcome variable (such as π_{t+1}), a is an intercept, and δ is a vector of parameters. Since economic theory leaves us rather agnostic about the 'true' model, the researcher may have some uncertainty over which instruments to include or exclude. All possible combinations of instruments form the model universe $M = [M_1, M_2, \ldots, M_K]$, where $K = 2^k$. The BMA solution to the problem is to weight the outcomes of all the models by their posterior probability. The fitted value $\hat{y}_t^{BMA}$ can be then expressed as:

$$\hat{y}_t^{BMA} = \sum_{k=1}^{K} \hat{y}_{t,k} p(M_k | y, Z), \tag{7}$$

where $\hat{y}_{t,k}$ denotes a fitted value conditional on the model k, and weights $p(M_k|y, Z)$ are the posterior model probabilities that arise from Bayes' theorem:

$$p(M_k|y, Z) = \frac{p(y|M_k, Z)p(M_k)}{p(y|Z)} = \frac{p(y|M_k, Z)p(M_k)}{\sum_{s=1}^{K} p(y|M_s, Z)p(M_s)}, \tag{8}$$

where $p(y|M_k, Z)$ denotes the marginal likelihood of the model, $p(M_k)$ is the prior probability that M_k is the 'true' model, and the denominator represents integrated likelihood, which is constant over the model universe. The expressions for the marginal likelihood $p(y|M_k, Z)$ depend on the problem at hand and vary across different kinds of models. In a linear regression setting, the marginal likelihood has a closed-form solution or can be obtained by approximation (depending on the nature of the priors on the coefficients).[14] Before running BMA, the researcher needs to specify the model universe (set of instruments), the model priors $P(M_k)$, and the parameter priors $P(\varpi|M_k)$, with $\varpi \equiv (a, \delta', \sigma^2)'$.

In our setting, y_t represents the endogenous variables in (5)[15] and the instrument set includes four lags of inflation, the output gap, the unit labor cost, long-term interest rates, the interest rate spread, unemployment, the nominal effective exchange rate, and the crude oil price. We aimed to include the most comprehensive set of instruments consistently with previous papers, subject to data availability. We use the hyper-g prior on the coefficients proposed by Liang et al. (2008) and run the Bayesian adaptive sampling algorithm (Clyde et al., 2010) to obtain the posterior probabilities over models.

formally treated in Plašil (2010), who shows that there are huge differences between instrument sets in terms of RMSC criterion (Hall et al., 2007). In theory, we could use this approach as well, but enumeration for all possible combinations of instruments would be computationally cumbersome, which calls for the use of methods like BMA, where tools to deal with potentially large instrument sets are readily available.

[14] BMA for linear models has been implemented in several statistical products. Here, we make use of the BAS package (Clyde et al., 2010), which is freely available in R.

[15] As we have shown above, the endogeneity problem enters the model through the replacement of inflation expectations with the observable value of future inflation. The forcing variable (the unit labor cost or the output gap) is usually considered exogenous. However, we believe that endogeneity of the output gap cannot be rejected a priori. For this reason, we formally treat the output gap as endogenous in the first-step regression and test for the presence of endogeneity in the second step by inspecting the statistical significance of the coefficient on the endogeneity correction term. The terms of trade in all specifications are considered exogenous. All other variables are predetermined since they enter the equation with some lag.

Note that BMA assumes a time-invariant relation between the target variable and the set of instruments. In light of our considerations above, it may seem necessary (or reasonable) to account for the time-varying nature of the parameters rather than model uncertainty. Recent evidence, however, suggests that traditional time-varying parameter models perform rather poorly in inflation-forecasting exercises and are outperformed by procedures accounting for model uncertainty (see Koop and Korobilis, 2009).[16]

To finish the first step we get residuals $\hat{v}_t = y_t - \hat{y}_t^{BMA}$, estimate Σ_v by $\hat{\Sigma}_v = \sum_{t=1}^{T} \frac{1}{T} \hat{v}_t \hat{v}_t^{'}$, and obtain the standardized residuals $\hat{v}_t^* = \hat{\Sigma}_v^{-1/2} \hat{v}_t$. These residuals are used as the auxiliary regressors in the second step and may be viewed as the endogeneity correction terms.

The hybrid NKPC (5) with added correction terms, time-varying coefficients, and stochastic volatility can be expressed as follows (see Nakajima, 2011, for general representations of time-varying regression and VAR models with stochastic volatility):

$$\pi_t = c_t' \kappa + x_t' f_t + \psi_t, \qquad \psi_t \sim N(0, \sigma_t^2) \tag{9}$$

$$f_{t+1} = f_t + u_t, \qquad u_t \sim N(0, \Sigma) \tag{10}$$

$$\sigma_t^2 = \gamma \exp(h_t) \tag{11}$$

$$h_{t+1} = \rho h_t + \eta_t, \qquad \eta_t \sim N(0, \sigma_\eta^2), \tag{12}$$

where $c_t \equiv (v_{t,\pi}^*, v_{t,gap}^*)'$ is a vector of the endogeneity correction terms, $x_t \equiv (\pi_{t+1}, \pi_{t-1}, s_t, TT_t)'$ is a vector containing key model covariates, κ is a vector of constant parameters, and $f_t \equiv (\gamma_{f,t}, \gamma_{b,t}, \lambda_t, \alpha_t)'$ represents a vector of time-varying coefficients.

The time-varying coefficients are constrained to follow a random walk, which allows for both permanent and transient shifts. Such a specification is designed to capture gradual smooth changes and/or structural breaks in the coefficients. Disturbances in (9), denoted ψ_t, are normally distributed with the time-varying variance σ_t^2. The log-volatility, $h_t = \log(\sigma_t^2/\gamma)$, is modeled as an AR(1) process.

The system of equations (9)-(12) forms a non-linear state space model with state variables α_t and h_t. The presence of stochastic volatility (the source of non-linearity) makes traditional estimation difficult because the likelihood function is intractable. However, Bayesian inference is still possible and we can estimate the model efficiently using Markov chain Monte Carlo (MCMC) methods.[17] To obtain the results, we drew M = 70,000 samples from the posterior distribution and discarded the first 50,000 samples as a burn-in period. Below we report the results for the default (quite loose) coefficient priors implemented by Nakajima (2011) in his code. As a robustness check we also experimented with other parameter settings in the prior densities, but the results do not seem to be severely affected by the choice of prior. Nevertheless, the mixing properties of the Markov chain improved as the priors got tighter. To check for convergence, we computed inefficiency factors (Geweke, 1992), which measure how well the Markov chain mixes. The inefficiency factors were

[16] Recently, Raftery et al. (2010) proposed a new method called dynamic model averaging (DMA) that accounts for both model uncertainty and parameter evolution. Since forecasting exercises have shown that BMA and DMA perform comparably at short horizons, and given that DMA is still computationally unfeasible for larger instrument sets, we regard BMA as a reasonable option. Note that DMA requires full enumeration of all models, which is memory and time consuming for K greater than, say, 2^{20}.

[17] Nakajima (2011) shows how to sample from the posterior distribution of coefficients using a Gibbs sampler and provides all the necessary computational details. See Nakajima (2011) also for the reference to his Ox and Matlab codes, which were (after some modifications) used for the estimation.

usually quite low (below 50). Occasionally, however, they reached values close to 100 for some coefficients. Nevertheless, this still implies that we get about M/100 = 200 uncorrelated samples, which is considered enough for posterior inference (see Nakajima, 2011). As a robustness check we also obtained posterior distribution of coefficients by sampling only every tenth draw, as this can reduce potential autocorrelation in the chain. Results, however, remained vastly identical.

As indicated above, one might also be interested in the structural parameters of the NKPC model. However, their direct estimation from a highly non-linear state-space model would be extremely difficult in practice. Since under quite mild conditions there exists one-to-one mapping between the reduced-form coefficients and the structural parameters, we avoid direct estimation of the structural parameters and instead use a non-linear solver to obtain their value from the median of the reduced-form coefficients' posterior distribution.[18]

3.3 Data

Our dataset combines time series taken from several data sources (ECB, Eurostat, OECD, IMF and national statistical offices). They were mainly downloaded from the E(S)CB data warehouse, which integrates series collected by the key supranational data providers. We used seasonally-adjusted (SA) data or performed our own adjustment based on X12 ARIMA when SA series were not directly available and statistical tests detected seasonality. Due to the limited data availability induced by the transition from a command to a free-market economy we are forced to use a relatively short time span, running from 1996 Q1 to 2010 Q4. One also has to take into account lower data quality – especially at the beginning of the sample as the statistical services in CE countries still faced some difficulties in meeting newly adopted statistical standards. In this respect, the results should be interpreted with some caution. In line with Galí and Monacelli (2005) the inflation rate is measured as the annualized quarter-on-quarter (log) difference in the harmonized index of consumer prices. To proxy the marginal cost we stick to the output gap taken from the OECD Economic Outlook[19] rather than the commonly used unit labor cost (labor share of income, LIS). The latter measure performed rather poorly in the cross-correlation pre-analysis and in the pre-estimation exercise. The terms of trade series are calculated as the ratio of the import price index to the export price index as taken from the Eurostat database.

In addition to the lags of the variables described above, our initial instrument set includes (four lags of) the unit labor cost, unemployment, the nominal effective exchange rate, the crude oil price, the long-term interest rate, and the interest rate spread. The spread is defined as the difference between 3M and overnight interbank interest rates.[20] As noted above, the number of four lags corresponds to that in most previous studies (see for example Galí et al., 2005). The results of the BMA procedure, which document the relative strengths of the individual instruments, are relegated to the Appendix. The R-square of the models with the highest posterior probability was around 0.8 for all three countries.

It is important to highlight that the inflation rate (especially for Hungary and Poland), along with some other variables, show a clear non-stationary pattern. Since it is not evident whether the non-stationarity is a result of the time-varying environment or is of an intrinsic nature, we rendered inflation stationary by shortening the estimation period to 1999 Q1–2010 Q4 and re-estimated

[18] We fixed the subjective factor β to 0.99.

[19] It seems to correspond by and large to the output gap obtained by the HP filter.

[20] We resort to this rather simplistic definition due to the limited availability of other interest rate data in the given period.

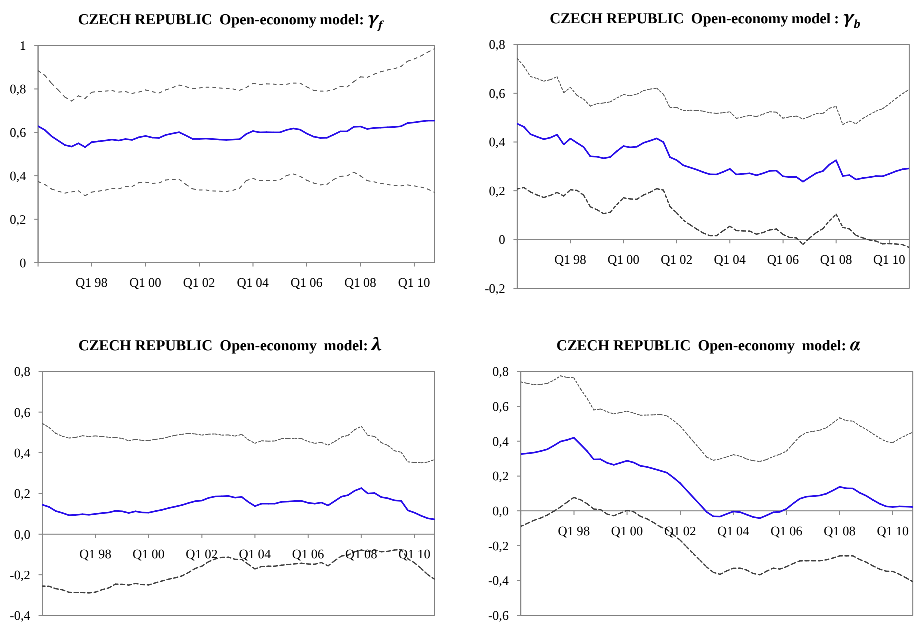
model (4). Given that the overall results remained largely identical we report the outcomes for the longer time span only.

4 Results

4.1 Open-economy NKPC with Time-varying Parameters

Figures (1–3) present the estimated time-varying open economy reduced-form coefficients for the Czech Republic, Hungary and Poland respectively.

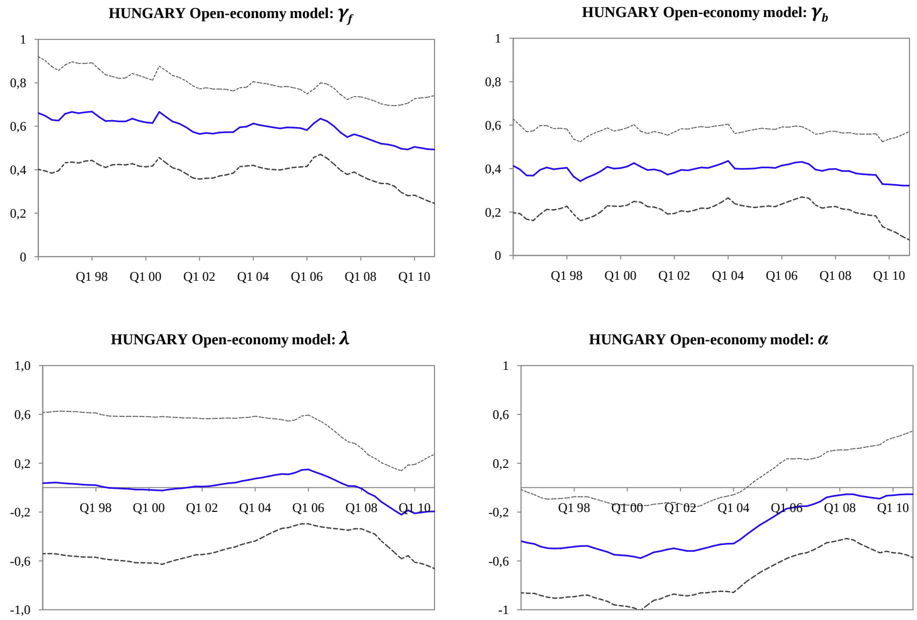
Figure 1: Czech Republic, Open-economy NKPC Coefficients



In general, we find evidence that the forward-looking inflation term is more important than the backward-looking one. This implies that inflation expectations formed in forward-looking fashion play an important role and are (at least partially) anchored in the CE countries. Consequently, monetary policy might be able to affect future inflation by influencing inflation expectations as such, for example by making a credible commitment to future policy actions. In this respect, central banks do not need to rely on interest rate changes only. Yet, the backward-looking term that arguably tracks the intrinsic inflation persistence remains largely significant (with a partial exception for the Czech Republic). This shows somewhat different picture to that found in studies for developed countries that use comparable econometric framework (e.g. Cogley et al., 2010; Hall et al., 2009). These studies usually argue that a proper treatment of potential structural instabilities enables to fully abstract from the existence of intrinsic inflation persistence.

The overall evolution in backward and forward-looking coefficients underlines their relative stability. However as usual, there are some exceptions to the rule which should not go unnoticed.

Figure 2: Hungary, Open-economy NKPC Coefficients



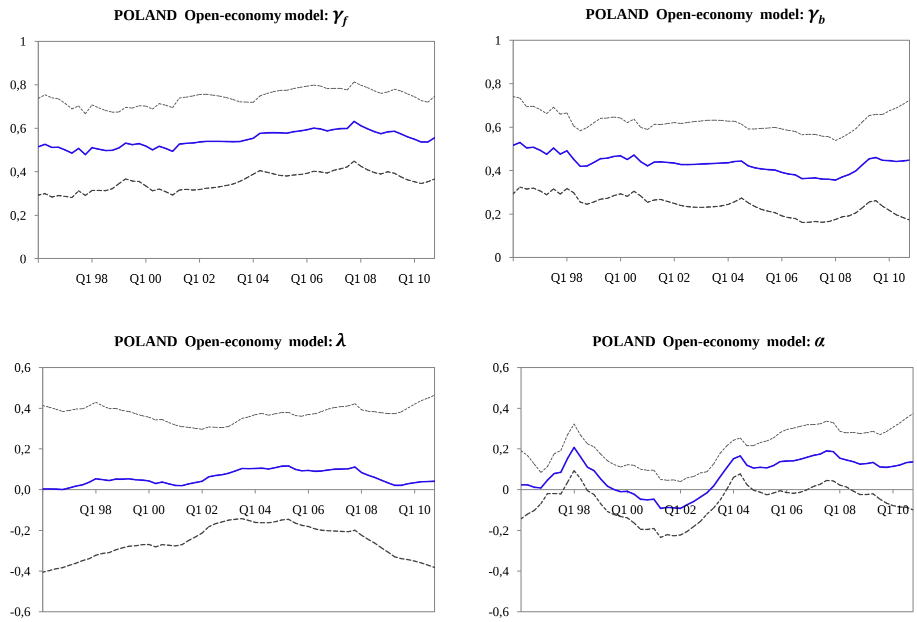
Despite considerable uncertainty as evidenced by the wide credible intervals one can occasionally observe that median as well as the whole posterior distribution tends downwards or upwards. The most conspicuous instance is the decrease of the backward-looking term γ_b to practically insignificant levels in the Czech Republic. Correspondingly, the coefficient of the forward-looking term exhibits a tendency of slight increase. Similar albeit less pronounced pattern can be found for Poland just until the outbreak of financial crisis. In 2008, the observed resemblance seems to come to a halt. On the contrary, in case of Hungary both coefficients are remarkably stable during the whole period under study.

Somewhat interestingly, we do not observe any peak or change in trend around the time when inflation targeting was adopted (i.e. in 1998 in the Czech Republic, in 1999 in Poland and in 2001 in Hungary), thus a shift to the new regime did not bring an immediate effect. For the first two countries the backward and forward-looking coefficients start to deviate around two years after the adoption of inflation targeting.[21] These changes were accompanied by overall slump in inflation rate below the inflation target in both countries. In the Czech Republic, the disinflation appeared shortly after the Czech National Bank decided to move from periodic setting of targets for the end of the year to continuous targeting of headline inflation within a predefined target range.[22] The National Bank of Poland originally set inflation targets in a similar manner as in the Czech Republic, but in 1999 and 2000 the actual inflation ran well above the upper bounds of the announced targets. The

[21]For Poland Lyziak (2003) and Orlowski (2010) document that inflation expectations actually became anchored to the target path about two years after inflation targeting was adopted.

[22]Initially, the target was continuously decreasing, from 3-5% to 2-4% between 2002 and 2005. However, since the inflation rate already often crawled below the inflation target, the effects of the subsequent shift to point targets in 2005 and the change of the targeted inflation rate from 3% to 2% in 2009 were negligible.

Figure 3: Poland, Open-economy NKPC Coefficients



inflation expectations were anchored to inflation targets shortly after the crawling peg was replaced with a pure float of Polish Zloty in April 2000.

On the contrary, the stability of coefficients in Hungary is somewhat surprising given that, at least formally, monetary policy framework has changed significantly over the past years.[23] The main policy difference vis-à-vis the former two countries lies arguably in the role attributed to exchange rate (Vonnak, 2008). Simultaneously with a shift to inflation targeting, the previous exchange rate regime of crawling band was replaced by the 'shadow' ERM II regime of fixed exchange rate with a fluctuation band of +/-15% around the central parity against the euro (the other two countries did not declare any specific exchange rate target and maintained free floating for most of the time). Although this de-facto meant a formal adoption of exchange rate target (alongside the official inflation target), Hungarian national Bank was not fully able to fulfill it in practice and some inflationary depreciation periods followed.[24] This is also supported by the fact that the exchange rate was identified as one of the most important factors of inflation expectations (see the BMA results in Appendix). With coefficients' evolution in mind, this narrative evidence seems to suggest that the continuous targeting is a preferable communication vehicle of monetary policy intentions alike it seems preferable to disregard the explicit exchange rate targets consistently with the Impossible trinity hypothesis (Obstfeld et al., 2005).

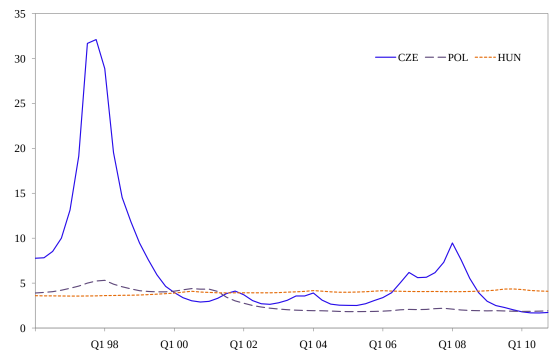
In this regard it also useful to inspect more closely the estimated volatility of inflation shocks, which is depicted in Figure (4). Most notably, the volatility of inflation shocks decreased quickly

[23] Inflation targets were announced at the end of the year for the following one until 2007. A policy based on a predefined medium-term target (set at 3%) was implemented only in 2008.

[24] The most significant one in terms of its effect on inflation occurred in 2004, when inflation increased from 3% to 7%.

a few years after the adoption of inflation targeting in the Czech Republic and Poland (standard deviation commonly around 2.5 in the Czech Republic and 2.0 in Poland) but remained rather stable and on average higher in Hungary (around 4). The most significant inflation shocks can be found in the Czech Republic (mostly induced by administrative measures) but these were short-lived and do not seem to have affected longer-term properties of the inflation dynamics. This arguably demonstrates that one-off shocks do not affect inflation expectations if they are well anchored.

Figure 4: Stochastic Volatility



The identification of inflation forcing variables turns rather challenging. The coefficient λ, measuring the impact of real economic activity (output gap) on inflation remains insignificant over the entire sample in all three countries. The median estimates vary in the common range of 0 - 0.2, but the credible intervals are too wide to draw any particular conclusion.[25] If we limit our attention solely to the estimated median, we can observe its mild increase for all three countries starting in 2001 and its synchronic decline since the onset of global economic crisis in 2008.

As a robustness check, we also explored performance of alternative domestic forcing variables: *(i)* the real unit labor cost, and *(ii)* the unemployment rate, but they both led to even less satisfactory outcomes than the output gap itself. Moreover, to address the uncertainty related to individual measures of real economy output we estimated a factor from all available series potentially tracking domestic inflation pressures (real GDP, domestic demand, industrial production, index of domestic wages and salaries, real unit labor cost, unemployment rate, consumer confidence indicator, current level of capacity utilization). The resulting factor tracked the common variation of all series (besides real unit labor cost, whose development is entirely idiosyncratic) relatively well, and was in fact quite similar to the original output gap. Consequently, the estimated coefficient λ was also almost identical.[26]

These results may have a number of explanations. First, the hypothesis often put forth for developed countries is that Phillips curve has flattened in the last decades. Indeed, if central banks are successful in keeping the inflation rates close to the target and its volatility is limited, it is rather difficult to find a stable relations vis-à-vis output gap, which is substantially more volatile. Second, the trade-off between inflation and economic activity can be non-linear, which means that

[25]Note, that in the conventional close-economy set-up this coefficient is usually higher and significant for some periods. These results are available upon request.

[26]These results are not reported to save a space and are available upon request.

the slope of the Phillips curve might depend on the actual size of the output gap: in normal times without recessions and with only mild output gaps, the relationship implied by the Phillips curve is negligible, but when larger expansions or recessions occur, the curve steepens (Stock and Watson, 2010). This argument can partially apply here, as there is some apparent volatility of the GDP growth rates in all three countries. The most blatant example relates to economic boom in the Czech Republic and Poland in 2005-2007 with growth rates exceeding 6% annually. The output gap was then reaching the highest positive values which was accompanied by the apparent (though insignificant) increase in coefficient λ. Third, and perhaps most importantly within the context of the transition or emerging economies, the low λ may be associated with factors specific to transition such as supply shocks caused by changing production structure of the economy or the fading impact of changes in regulated prices. These factors cause shifts of the Phillips curve rather than movement along it.[27] Fourth and finally, there is some intuition that open trade and capital flows weaken the effect of domestic real activity on inflation (Razin and Yuen, 2002; Razin and Loungani, 2005). This is consistent with alternative hypothesis of Phillips curve flattening, which points to the effects of globalization rather than to monetary policy (Borio and Filardo, 2007). All three countries under study are small open economies highly integrated with international markets, especially the euro area. As a result, a large proportion of domestic production is destined for foreign markets and a significant share of both intermediate and final products is imported. Therefore, domestic consumer inflation shall be rather determined (at least partially) by external factors.

As advocated by Galí and Monacelli (2005) and subsequent authors, the terms of trade, which track relative changes in import and export prices, can thus be considered a second forcing variable for the inflation dynamics. Yet, the corresponding estimate of coefficient α brings a mixed evidence. A certain pattern of trade-off between domestic and foreign inflation factors can be found in the Czech Republic where foreign factors dominate in the first half of the sample and domestic in the latter. For Poland, we find a predominance of foreign factors with the corresponding coefficient α being significant in several periods following large depreciations (1997 Q4–1998 Q2, 2003 Q4–2004 Q3) and again at the onset of the late-2000s recession (2007 Q1–2008 Q2). For Hungary the estimated coefficient α is significantly negative on part of the sample. This goes against to the underlying theory, however coincides with the findings of Mihailov et al. (2011b).

A possible explanation why effects of external factors on inflation are only of temporary (and possibly non-linear) nature might be that these are already reflected in inflation expectations. If domestic firms engage in foreign trade, their inflation expectations are likely to be influenced by the exchange rate. For instance, in Hungary the exchange rate was on depreciating path for the most of the time and thus it arguably made agents take future currency depreciation already into account when forming their inflation expectations. Indeed in Hungary, the NEER turned out to be the most relevant variable in the first-step BMA results, but external factors' effect in (4) proved to be very limited and disputable. On the contrary, Poland experienced a few sudden and generally unexpected depreciation periods (as noted above), which caused a huge temporary blip in terms of trade with significant impact on inflation.

As in case of the domestic forcing variables we checked the robustness of our results using *(i)* a simple first difference of the terms of trade as well as it is deviation from the HP-filtered trend, but the results based on theoretic model are still preferable (as also found by Mihailov et al. (2011a)), *(ii)* both the difference and deviations from the HP-trend of the NEER but these variables were

[27]However, the first-step BMA results show that real factors are often relevant for inflation forecasts (inflation expectations). It may follow that the mutual relationship between inflation and real drivers is actually more complicated than the standard NKPC suggests.

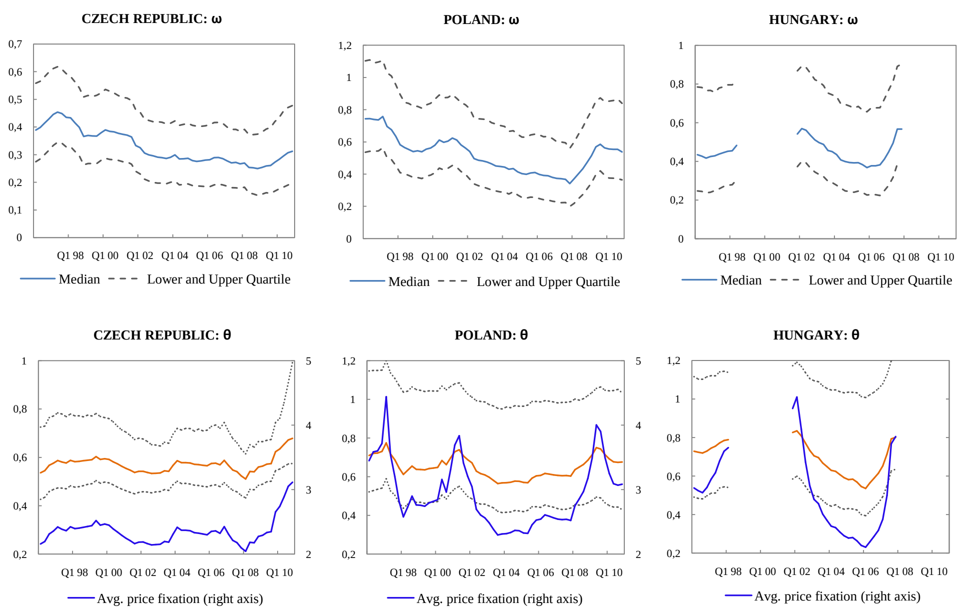
generally insignificant.[28]

4.2 Are the NKPC Structural Coefficients Truly Structural?

While structural coefficients are standardly reported in papers based on the time-invariant framework, time-varying studies usually do not go that far. The idea that deep structural coefficients vary in time is rather controversial. However, as put forth above, there are numerous reasons why this may be the case, especially in transition countries. For example, the structural nature of the Calvo parameter can be hardly justified in a fully microfunded model, because the exogenous probability of the price changes is likely depend on the average size of nominal shocks.

We use the estimates of the reduced-form coefficients to obtain a sequence of structural coefficients, namely, *i)* the share of backward-looking price setters ω and *ii)* the average time for which prices remain fixed as a function of θ. As already mentioned above, the structural coefficients were derived under the assumption that subjective factor, β, is fixed to 0.99. The results are reported in Figure (5). It is also important to note, that since estimates of λ are insignificant, one has to read the results with a great caution. Moreover, for Hungary, the reduced-form coefficient of the output gap, λ, turns negative in some periods, which impedes obtaining structural coefficient θ in reasonable range (these periods are dropped and the corresponding coefficient series is discontinuous albeit still holds some information).

Figure 5: Estimates of Structural Coefficients



On the upper panel we can appreciate some differences in the share of the 'rule of thumb' firms, ω. The most economically consistent picture can be drawn for the Czech Republic. The share of

[28]These results are available upon request.

backward-looking firms that adjust prices simply to the observed inflation in the previous period has been slowly trending downwards. These developments correspond to our expectations. First, during the transition, firms faced continuously increasing competition and needed to change their pricing policy with respect to the market conditions. Second, over time, the share of the administered prices in overall inflation (which are typically set in a backward-looking manner) decreased. Third, forward-looking price setting is arguably subject to learning. Therefore, a decreasing share of backward-looking price setters may signal that agents are becoming more sophisticated and form their expectations in a rational (i.e., forward-looking) rather than adaptive (i.e., backward-looking) fashion. These findings seem to suggest some kind of convergence in price setting to the euroarea where the backward-looking price setting was found negligible (Galí et al., 2001). Similar, although a bit more humpy pattern[29] can be found for Poland where this share decreases from substantially higher levels. The pattern for Hungary where the estimate is subject to the largest uncertainty is again humpy and some decreasing trend is apparent only from 2000. Interestingly, for all three countries we can observe increase in this share as the global crisis unfolds. The cross-country variation in the share of backward-looking price setters can be driven by multiple institutional differences in price setting and expectation formation behaviour (which are unfortunately not well explored in cross-country context, yet).

On the bottom panel we depict the length of price fixation derived from the structural parameter θ, capturing overall price rigidity. Again we find rather stable pattern for the Czech Republic and a humpy pattern for Poland and Hungary. For the former country we find the shortest period of average price fixation that slowly increases from two quarters up to three. This result may be associated with decreased volatility of inflation, which in turn translates into longer periods when prices remain unchanged. For Poland, we can observe clear pattern of trendless oscillation around mean value close to three quarters. However, there are two notable peaks; the first around the late-2000s recession and the second around the global financial crisis (late 2007 till mid-2009). A concurrent upward shift in structural parameter θ and a deep slump in output suggest the existence of downward price rigidities, which mirror the increase of average fixation during these periods. The values of average price fixation for Hungary are subject to very substantial variation and uncertainty, which can be attributed to very low and sometimes even negative estimates of λ (unreasonable values of price fixation are not depicted in the chart). If anything, we again can find some pattern of variation over the time.[30] The cross-country differences can be linked to numerous institutional disparities. One reason noted in Rumler (2007) can be that more open economies can be prone to lower structural price rigidity as firms import from volatile international markets and need to change prices more often. This could explain why the average price fixation is lower in relatively more open Czech economy than Polish one (discarding the unreliable results for Hungary).

5 Conclusions

This paper aims to shed some light on possible changes in inflation dynamics under the presence of structural changes in economy and monetary policy regime. It analyzes the dynamics of inflation through the lens of the New Keynesian Phillips curve nested within a time-varying framework using data of three CE countries. Although originally, the NKPC was proposed as a structural model

[29] This is mainly due to low estimates of λ in some periods and highly-nonlinear mapping between reduced-form and structural coefficients.

[30] These findings can be confronted with stylized facts on pricing behavior available maily for developed countries (Taylor, 1999; Klenow and Malin, 2010).

of inflation dynamics which is invariant to policy changes, it is likely that substantial changes on the macroeconomic level coupled with large-scale restructuring of whole economies resulted also in significant changes at the microeconomic level.

The changes in inflation dynamics are often linked to monetary policy. In particular, the recent decrease in intrinsic inflation persistence was related to a more aggressive stance by central banks but also to adoption of stable monetary policy regimes with well-defined nominal anchor. In the period under study, the countries in Central Europe went through a unique historical episode where monetary and exchange rate regimes changed substantially. All three countries in our sample adopted the inflation-targeting framework, which is generally believed to stabilize inflation and reduce its persistence and variability through anchoring inflation expectations (although the international evidence is not entirely conclusive).

In general, we find that considerable structural changes experienced by economies under study coupled with shifts in monetary regimes therein produced some notable changes in inflation process but their nature is still rather heterogeneous across countries - despite their relative closeness. It seems that adoption of inflation targeting per-se does not guarantee desirable outcomes such as the decrease of intrinsic inflation persistence or inflation volatility. The details of implementation might matter as well, which supports the need of structural country-level analysis. If inflation targeting is not sufficiently credible, economic agents might chiefly take into account observed inflation levels rather than the inflation target. However, the predominance of the forward-looking component in the NKPC implies that forward-looking inflation expectations that are usually guided by declared inflation targets play a substantial role.

Refering to the cross-country differences, we found that intrinsic inflation persistence decreased substantially only in the Czech Republic. A predominantly forward-looking nature of the inflation process is commonly reported in recent studies on the US and large EU countries (Hondroyiannis et al., 2009; Cogley and Sbordone, 2008; Benati, 2008). This implies that lower inflation can be achieved by anchoring inflation expectations and does not need to be accompanied by output or employment losses. These findings are also supported by the estimated volatility of inflation shocks, which decreased quickly a few years after the adoption of inflation targeting in the Czech Republic but also in Poland. Indeed, predominantly forward-looking nature of inflation process and short-lived inflation shocks point to well-anchored inflation expectations. On the contrary, there is a surprisingly high stability of almost all characteristics describing inflation process in Hungary. An intuitive explanation points to the role of exchange rate in Hungarian monetary policy. Indeed, inconsistency between inflation and exchange rate target is well-described as Impossible trinity (Obstfeld et al., 2005).

The overall changes in inflation process can be traced down to changes in microeconomic behavior, in our case principally the price setting behavior of firms. We found some evidence that the 'structural' coefficients describing this behavior might not be entirely stable across time and their variation can be in turn related to macroeconomic environment.

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Appendix

A: BMA results[31]

Figure A1: Czech Republic

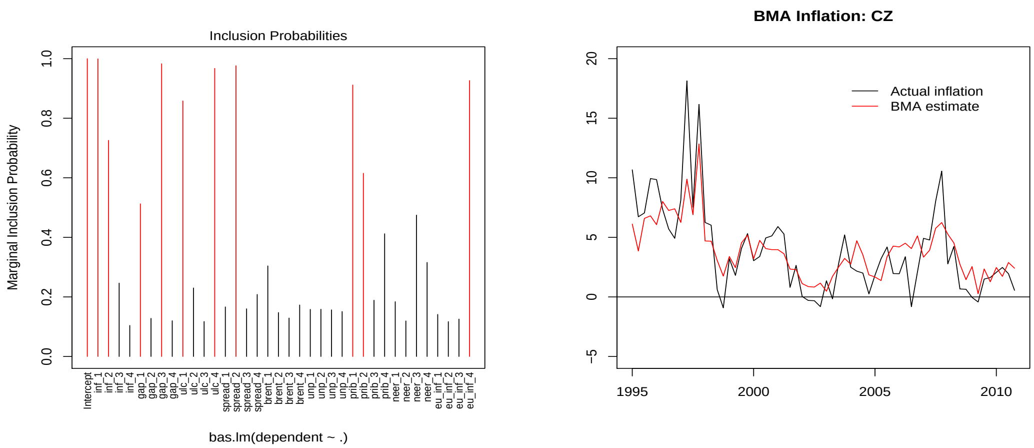
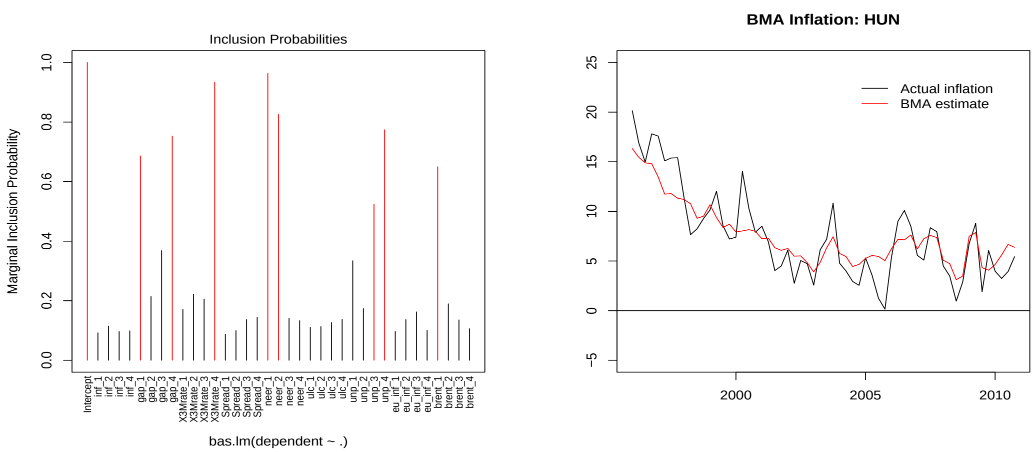
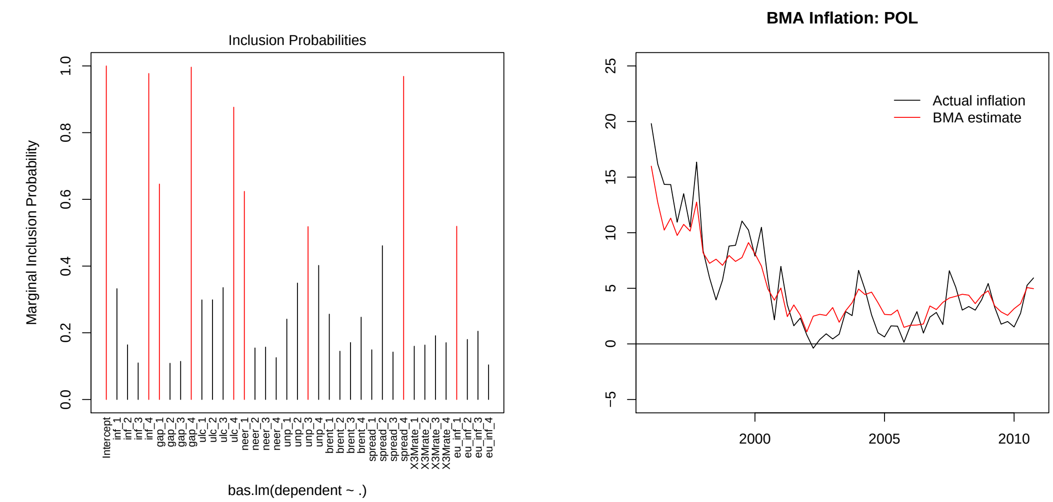


Figure A2: Hungary



[31] Red bars indicate posterior inclusion probability higher than 0.5.

Figure A3: Poland



B: (In)stability of coefficients

Figure B1: Flexible Least Squares: Residual efficiency frontier

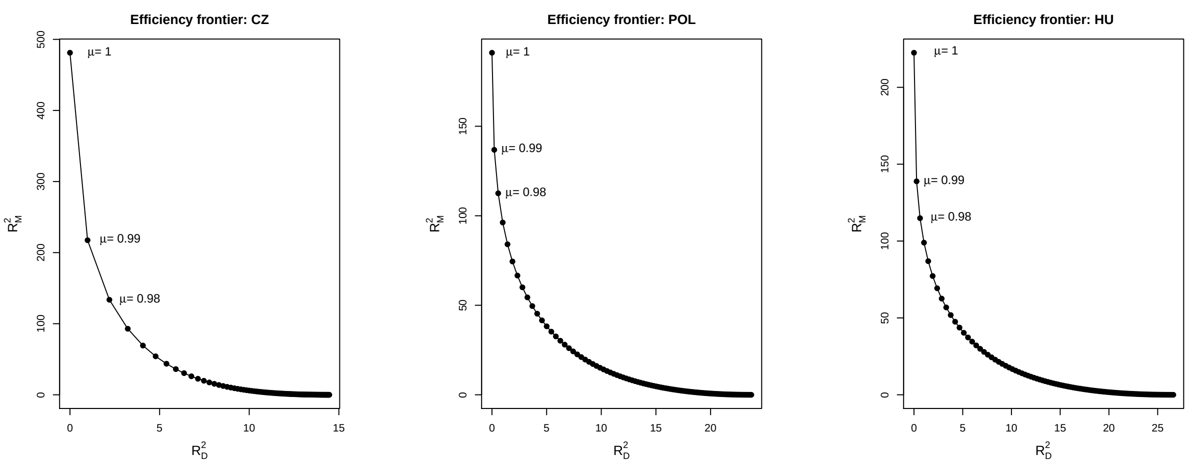


Table B1: Time invariant GMM estimates across different time-spans and instrument sets

	GG99		**GGL05**		**all_lags1_2**		**all_4lags**		**BMA**	
	coef	***pval***	**coef**	***pval***	**coef**	***pval***	**coef**	***pval***	**coef**	***pval***
CZ										
γ_f	0.54	*0.00*	0.43	*0.32*	0.41	*0.22*	0.41	*0.11*	0.59	*0.00*
γ_b	0.45	*0.01*	0.56	*0.13*	0.55	*0.05*	0.61	*0.01*	0.41	*0.00*
λ	0.05	*0.62*	-0.02	*0.81*	0.00	*0.99*	-0.07	*0.61*	-0.01	*0.93*
α	0.13	*0.85*	1.60	*0.00*	2.08	*0.13*	1.45	*0.11*	1.73	*0.00*
Subsample 96-02										
γ_f	0.44	*0.21*	0.58	*0.17*	0.10	*0.90*	0.28	*0.47*	0.51	*0.00*
γ_b	0.56	*0.07*	0.47	*0.20*	0.84	*0.13*	0.77	*0.05*	0.48	*0.00*
λ	0.00	*1.00*	-0.14	*0.46*	0.16	*0.83*	-0.09	*0.71*	-0.04	*0.80*
α	0.95	*0.05*	1.55	*0.00*	2.79	*0.07*	1.50	*0.02*	1.96	*0.00*
Subsample 03-10										
γ_f	0.61	*0.03*	0.71	*0.00*	0.67	*0.03*	0.64	*0.01*	0.68	*0.00*
γ_b	0.41	*0.09*	0.31	*0.10*	0.37	*0.14*	0.39	*0.09*	0.29	*0.01*
λ	0.05	*0.55*	0.04	*0.72*	0.04	*0.69*	0.02	*0.73*	0.05	*0.47*
α	-0.43	*0.65*	1.09	*0.22*	-0.58	*0.57*	0.04	*0.97*	1.35	*0.06*
POL										
γ_f	0.58	*0.00*	0.59	*0.01*	0.81	*0.00*	0.71	*0.00*	0.63	*0.00*
γ_b	0.43	*0.00*	0.42	*0.01*	0.29	*0.09*	0.35	*0.03*	0.40	*0.05*
λ	0.01	*0.96*	-0.03	*0.81*	0.04	*0.77*	-0.02	*0.91*	-0.01	*0.92*
α	0.02	*0.82*	-0.02	*0.79*	0.07	*0.41*	-0.01	*0.88*	-0.04	*0.57*
Subsample 96-02										
γ_f	0.51	*0.03*	0.60	*0.27*	0.73	*0.00*	0.64	*0.00*	0.38	*0.06*
γ_b	0.49	*0.00*	0.43	*0.15*	0.35	*0.03*	0.39	*0.00*	0.32	*0.04*
λ	-0.16	*0.52*	-0.17	*0.50*	-0.03	*0.77*	-0.01	*0.94*	0.40	*0.15*
α	-0.03	*0.77*	-0.04	*0.70*	0.01	*0.95*	0.05	*0.34*	0.09	*0.08*
Subsample 03-10										
γ_f	0.54	*0.01*	0.52	*0.13*	0.65	*0.04*	0.38	*0.27*	0.45	*0.03*
γ_b	0.54	*0.00*	0.49	*0.01*	0.64	*0.07*	0.76	*0.00*	0.67	*0.00*
λ	0.04	*0.82*	0.03	*0.91*	-0.14	*0.69*	-0.09	*0.66*	0.01	*0.95*
α	0.21	*0.00*	0.35	*0.02*	0.60	*0.03*	0.02	*0.94*	-0.21	*0.74*
HUN										
γ_f	0.27	*0.68*	0.88	*0.11*	0.45	*0.12*	0.49	*0.15*	0.64	*0.00*
γ_b	0.56	*0.00*	0.42	*0.00*	0.57	*0.03*	0.58	*0.11*	0.51	*0.00*
λ	-0.03	*0.85*	0.06	*0.72*	-0.02	*0.88*	-0.01	*0.96*	-0.04	*0.75*
α	-0.75	*0.31*	-0.15	*0.81*	-1.18	*0.66*	-1.09	*0.66*	0.23	*0.82*
Subsample 96-02										
γ_f	0.78	*0.55*	0.37	*0.39*	0.95	*0.17*	0.77	*0.18*	0.36	*0.26*
γ_b	0.31	*0.45*	0.30	*0.04*	0.25	*0.78*	0.41	*0.36*	0.40	*0.25*
λ	0.06	*0.84*	0.01	*0.98*	0.11	*0.76*	0.07	*0.79*	-0.51	*0.32*
α	0.95	*0.39*	0.72	*0.06*	1.36	*0.40*	0.86	*0.59*	0.76	*0.42*
Subsample 03-10										
γ_f	0.76	*0.11*	0.55	*0.04*	0.46	*0.09*	0.67	*0.03*	0.41	*0.01*
γ_b	0.42	*0.24*	0.41	*0.00*	0.50	*0.03*	0.40	*0.08*	0.67	*0.00*
λ	0.05	*0.84*	0.10	*0.65*	0.01	*0.98*	0.07	*0.75*	-0.10	*0.58*
α	-1.49	*0.17*	-0.71	*0.37*	-1.02	*0.59*	-1.08	*0.65*	-1.01	*0.23*

Table B2: Break point tests with break point at 2003 Q1

Model	Czech Republic A-F Wald	A-F Lik.Rat.	Poland A-F Wald	A-F Lik.Rat.	Hungary A-F Wald	A-F Lik.Rat.
GG99	57.932	136.735	15.436	45.476	33.054	4148.214
	0.000	**0.000**	**0.004**	**0.000**	**0.000**	**0.000**
GGL05	12.809	73.019	19.798	36.941	7.231	42.208
	0.0123	**0.000**	**0.000**	**0.000**	**0.124**	**0.000**
All lags 1,2	181.38	919.697	44.99	476.562	26.653	88.952
	0.000	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**
All lags 1-4	49.135	244.063	1.574	13.403	8.074	18.010
	0.000	**0.000**	0.813	**0.010**	**0.089**	**0.001**
BMA	3.876	5.774	12.562	24.344	2.656	13.520
	0.423	0.217	**0.014**	0.001	0.617	**0.009**