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

Jaromír Baxa, Miroslav Plašil, Bořek Vašíček *

Abstract

The purpose of this paper is to provide a novel look at the evolution of inflation dynamics in selected Central European (CE) countries. We use the lens of the New Keynesian Phillips Curve (NKPC) nested within a time-varying framework. Exploiting a time-varying regression model with stochastic volatility estimated using Bayesian techniques, we analyze both the closed and open-economy version of the NKPC. The results point to significant differences between the inflation processes in three CE countries. While inflation persistence has almost disappeared in the Czech Republic, it remains rather high in Hungary and Poland. In addition, the volatility of inflation shocks decreased quickly a few years after the adoption of inflation targeting in the Czech Republic and Poland, whereas it remains quite stable in Hungary even after ten years' experience of inflation targeting. Our results thus suggest that the degree of anchoring of inflation expectations varies across CE countries. In addition, we found some evidence that the 'structural' parameters of the NKPC are somewhat related to the macroeconomic environment.

JEL Codes: C11, C22, E31, E52.

Keywords: Bayesian model averaging, Central European countries, inflation dynamics, New Keynesian Phillips curve, time-varying parameter model.

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Nontechnical Summary

Understanding the nature of short-term inflation dynamics poses a major challenge for monetary policy. The traditional Phillips curve postulated that there is a stable trade-off between inflation and economic activity. At the same time, it has almost become common wisdom that inflation is very persistent. Consequently, taming inflation was deemed to be costly in terms of output loss. However, a better understanding of the role of expectations has changed the perceptions of monetary policy conduct. Since inflation is believed to be affected not only by current and past monetary policy, but also by the commitment to future monetary policy actions, a credible monetary policy that anchors inflation expectations can achieve disinflation at no cost in terms of real output. This idea was formalized into the New Keynesian Phillips curve (NKPC), which appeared during the 1990s. Its main ingredient is a forward-looking inflation term tracking the effect of inflation expectations on the current value of inflation.

The NKPC was proposed as a structural model of inflation dynamics, in the sense that it is a result of an optimization process at the micro level and thus is invariant to policy changes. In practice, however, there are numerous reasons why the nature of the inflation process can evolve over time. Importantly, the implementation of a credible monetary policy framework can stabilize inflation and reduce its persistence and variability through anchored inflation expectations. Macroeconomic changes can in turn feed back to the microeconomic environment. The countries in Central Europe went through a unique episode where both macro and microeconomic factors might have played a role in triggering changes in inflation dynamics in the last two decades. Their economies underwent significant structural changes coupled with changes in monetary and exchange rate regimes.

This paper aims to provide evidence on inflation dynamics in some CE countries that have adopted inflation targeting (the Czech Republic, Hungary, and Poland) by means of the NKPC nested within a time-varying framework. To estimate a model with time-varying parameters we resort to Bayesian techniques. We track the overall inflation dynamics along with changes in ‘structural’ parameters such as price stickiness.

We find that the nature of the inflation process differs across the selected CE countries. Although the forward-looking component dominates the inflation dynamics in all three countries, which is a sign of (at least partially) anchored inflation expectations, inflation is considerably less persistent in the Czech Republic than in Hungary and Poland and the persistence has been constantly decreasing. In addition, the volatility of inflation shocks decreased quickly a few years after the adoption of inflation targeting in the Czech Republic and Poland, while it remains rather stable in Hungary even ten years after inflation targeting was adopted. These two results show that inflation expectations are well anchored in the Czech Republic while this is less true for Poland and Hungary.

We have found some evidence that ‘structural’ coefficients are not stable over time as is commonly believed. We show that the average time for which prices remain fixed is negatively correlated with both the level and the volatility of inflation. That is, if inflation is high and volatile, firms tend to change prices more frequently. The share of backward-looking price setters, who simply adjust their prices for observed inflation rather than in a forward-looking fashion, is changing smoothly, with a predominantly downward-sloping trend. It seems that this is being driven by long-term factors such as increasing competition, decreasing administered prices, and the learning capacity of price setters.

1. Introduction

The New Keynesian Phillips curve (NKPC) has become a workhorse macroeconomic model for studying the relation between inflation and real economic activity, notably in the domain of optimal monetary policy conduct. In broader terms, the NKPC is a core element of New Keynesian DSGE models (Smets and Wouters, 2003). The NKPC is built around the concept of staggered price-setting (or wage-setting), motivated either by staggered contracts (Taylor, 1980; Rotemberg, 1982) or by a probabilistic approach (Calvo, 1983). The NKPC was proposed as a structural model of inflation dynamics (Galí and Gertler, 1999; Galí et al., 2001), in the sense that it is a result of an optimization process at the micro level and thus is invariant to policy changes. In practice, however, there are potentially numerous reasons why the nature of the inflation process can evolve across time, and related empirical evidence seems to confirm this claim.

In general terms, the (macro)economic structure is constantly changing, and when we look at past decades these changes have been quite substantial. There is a long history of empirical research on changes in the business cycle and inflation persistence. Kim and Nelson (2006) and McConnell and Perez-Quiros (2000) provided groundbreaking evidence for the US and Stock and Watson (2003, 2005) for other countries, which initiated a debate about the Great Moderation. Corvoisier and Mojon (2005) find that the mean inflation of OECD countries has been subject to two or three structural breaks since the 1960s. The decrease of inflation persistence has been attributed to a more aggressive monetary policy stance in the US (Davig and Doh, 2008) and to the implementation of credible monetary policy regimes such as inflation targeting elsewhere (Benati, 2008). The linkage between changes in inflation dynamics and monetary policy is further corroborated by evidence about structural changes in monetary policy itself (Baxa et al., 2010; Boivin, 2006; Kim and Nelson, 2006; Koop et al., 2009; Sims and Zha, 2006; Trecroci and Vassalli, 2010).

From the microeconomic point of view, there are various reasons why agents' behavior might evolve over time, which in turn induce changes in the key 'structural' parameters of the NKPC. Some of these changes may even be triggered by changes on the macroeconomic level. The most obvious case is firms' decisions on the frequency of price adjustment. Most microeconomic studies on price-setting find the level and variability of inflation to be one of the key determinants of the frequency of price changes (Klenow and Malin, 2010). Fernandez-Villaverde and Rubio-Ramirez (2008) show within the DSGE framework that movements of pricing parameters are indeed correlated with inflation.

The countries in Central Europe went through a unique episode where both macro and microeconomic factors might have played a role in triggering changes in inflation dynamics in the last two decades. Their economies underwent significant structural

changes coupled with changes in monetary and exchange rate regimes. It is likely that these factors also implied significant changes at the microeconomic level. In particular, the level of inflation and monetary policy credibility could have induced changes in the price-setting behavior of individual firms. While there is some micro evidence (Babetskii et al., 2007; Konieczny and Skrzypacz, 2005; Coricelli and Horváth, 2010), changes in the nature of the overall inflation process are practically undocumented.

The purpose of this paper is to fill this gap and provide evidence on inflation dynamics in CE countries (the Czech Republic, Hungary, and Poland) through the lens of the NKPC nested within a time-varying framework. First, we estimate a standard hybrid version of the NKPC (Galí and Gertler, 1999) and track the overall inflation dynamics along with changes in ‘structural’ parameters such as price stickiness. Second, to study the impact of external drivers on inflation we estimate an open-economy version of the NKPC in the spirit of Galí and Monacelli (2005). We slightly depart from their original (purely forward-looking) model and consider a hybrid version of the NKPC, just as in the case of the closed-economy form. In our two-step procedure closely related to Kim (2006) we estimate a time-varying regression model with stochastic volatility using Bayesian techniques. In addition, we use Bayesian model averaging in order to tackle the issue of instrument selection, as it has been shown to be very relevant in forward-looking models in many previous papers.

Our results can be summarized as follows. First, we find that the nature of the inflation process differs across the selected CE countries. Despite the fact that the forward-looking component dominates the inflation dynamics in all three countries, inflation is considerably less persistent in the Czech Republic than in Hungary and Poland. Second, the changes in the inflation process over time are also rather heterogeneous. Inflation persistence as tracked by the coefficient on the backward-looking term has decreased substantially in the Czech Republic. This has been accompanied by a corresponding increase in the forward-looking term. Additionally, the volatility of inflation shocks decreased quickly a few years after the adoption of inflation targeting in both the Czech Republic and Poland, suggesting that inflation targeting and other policy changes influenced the inflation dynamics in these two countries. By contrast, the nature of the inflation process in Hungary does not seem to have changed much over the last 15 years. Third, the estimated coefficients of the domestic driving variable were often statistically insignificant. This feature may be linked to potentially important supply shocks during the transition, which cannot be fully captured by the original NKPC model. The relative importance of foreign inflation factors, tracked by the terms of trade, is relatively negligible as well, suggesting that the foreign factors might already be well reflected in inflation expectations themselves. Fourth, we find some evidence that both the level and the volatility of inflation are negatively correlated with the average time for which prices remain fixed. Therefore, it seems that the price-setting behavior of economic agents is somewhat related to the macroeconomic environment rather than being fully

invariant to it as the benchmark NKPC assumes. Our findings have some noteworthy policy implications. Previous research suggested that the implementation of a credible monetary policy regime contributed to a decrease of inflation persistence in the most developed countries. Although all three CE countries officially adopted inflation targeting a decade ago, inflation persistence has not changed considerably in Poland and Hungary and remains at high levels when compared to the Czech Republic or developed countries. This could be related to the fact that inflation targets in these countries are less credible and economic agents chiefly take into account observed inflation levels rather than the inflation target.

The paper is organized as follows. In section 2, we review relevant literature, focusing particularly on empirical aspects of NKPC estimation. 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). GG introduced estimation by GMM techniques and proposed a ‘hybrid’ modification of the original forward-looking model. This modification encompasses an effort to provide some structural justification for inflation persistence that the ‘pure’ version of the NKPC was unable to capture.

Despite the theoretical appeal of the NKPC, consecutive studies have produced rather conflicting empirical evidence with results varying across economies, data sets, and – most notably – estimation methods. The econometric approach of GG was heavily criticized by a few later authors (e.g. Rudd and Whelan, 2005; Mavroeidis, 2005), mainly on the grounds of the questionable behavior of the GMM estimator in the NKPC context. The common criticisms include sensitivity to the choice of instrument set, weak identification, and small sample bias. To overcome the potential pitfalls related to the GMM estimator, Ireland (2001) and Lindé (2005) advocate a system approach using the full information maximum likelihood method as it provides more efficient parameter estimates than limited information (i.e., ‘single-equation’) methods such as the GMM. In their response to Lindé and other critiques Galí et al. (2005) claim (with reference to Cochrane, 2001) that the issue of which estimation approach is preferable is completely open since there are no theorems or Monte Carlo simulations that suggest that one outperforms the other. In addition, they show that when the NKPC is correctly specified one obtains fairly robust results across estimation methods.

The stock of econometric techniques has progressively expanded. Some authors use Bayesian techniques (e.g. Smets and Wouters, 2003) or the minimum distance approach (Sbordone, 2005; Christiano et al., 2005). Kleibergen and Mavroeidis (2009) proposed robust versions of the GMM estimator. Other papers stick to the VAR framework and assess the validity of the NKPC by testing the set of restrictions (in the spirit of Campbell and Shiller, 1987). Fanelli (2008) analyzes the idea that forward-looking agents calculate their expectations within a VAR-like setting (with inflation and forcing variables), which allows one to deal with the issue of the feedback effect from inflation to the forcing variables. This paper rejects the validity of the NKPC for the euro area. Carriero (2008) obtains similar negative evidence with the US data, suggesting, however, that this result may indicate failure of the rational expectations hypothesis rather than NKPC-consistent forward-looking behavior. The relevance of rational expectations is further tested by Nunes (2010), who estimates the NKPC for the US economy considering firms represented by rational expectations as well as firms represented by survey expectations. He finds that although survey expectations can be a determinant of inflation dynamics, rational expectations seem to be predominant. Dees et al. (2009) use global VAR to solve the weak instrument problem. In particular, they construct valid instruments using weighted averages of the global variables. Harvey (2011) points to the problem that the NKPC cannot appropriately account for nonstationarity, which is usually dealt with in an ad-hoc fashion such as by the application of detrended variables. Instead, he proposes a model where lagged inflation in the NKPC is replaced by an unobserved random walk component. Kontonikas (2010) generalizes the NKPC using the ARDL bound approach (Pesaran et al., 2001), which is suitable for variables with any order of integration. He finds with US data starting in the 1960s that higher marginal costs increase inflation.

Leaving aside the question of estimation, there are two other strands of literature that seek to improve the model's fit. The first strand tries to find a good proxy for the marginal cost or another appropriate inflation-forcing variable (notably for open economies), while the latter studies the effects of changes in the economic system and monetary policy on inflation dynamics.

In the empirical literature, firms' marginal costs are proxied by the labor income share (LIS) – a measure based on Cobb-Douglas production technology. While the measure may be applicable for the US and other major countries, some modifications need to be made for small open economies. In general terms, 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). Galí and Monacelli (2005) derive a small open economy version of the NKPC for CPI inflation, which includes (the difference in) the terms of trade as an additional forcing variable (above the marginal cost). While this model assumes complete exchange-rate pass-through, Monacelli (2005) relaxes this assumption. Mihailov et al. (2011a) provide the first empirical evidence based on this

model. Batini et al. (2005) 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 to include the cost of intermediate inputs (both domestic and imported) and finds some plausible evidence for the euro area countries. Regardless of the proposed corrections of the marginal cost to account for external effects, some authors (e.g. Rudd and Whelan, 2007) cast severe doubts on the appropriateness of the LIS measure itself, claiming that the LIS is in fact intrinsically countercyclical in nature. Consequently, Mazumder (2010) proposed a new measure that corrects the LIS by relaxing overly restrictive assumptions such as free adjustment of labor input at a fixed wage rate. This measure of the marginal cost turns out to be procyclical. Mazumder (2011) claims that the cyclicity of the selected marginal cost proxy is crucial for the sign of the corresponding coefficient in the NKPC. Paradoxically, however, if the marginal cost is procyclical, as is commonly believed, its coefficient in the NKPC has a counter-intuitive negative sign.

A few recent studies more closely related to our research fall into the second strand of literature. They consider the effects of changes in the economic system and monetary policy and explore how these changes are propagated into changes in the parameters of the NKPC. In general terms, these studies allow the nature of inflation dynamics to change over time. Most of the evidence is available for the US. Hall et al. (2009) use a time-varying model, which arguably corrects for the specification bias (due to incorrect functional forms, omitted variables, and measurement errors) inherent to fixed-coefficient estimation. They conclude that the lagged inflation term in the ‘hybrid’ version turns out to be insignificant. In a similar vein, Cogley and Sbordone (2008) claim that inflation persistence in the NKPC arises due to variation in the long-run trend component of inflation, which can be attributed to monetary policy shifts. Once log-linearization around the time-varying inflation trend is taken (in their two-step VAR estimation), the ‘pure’ forward-looking NKPC explains the US inflation dynamics fairly well. Zhang and Kim (2008) find with inflation survey data (and recursive GMM estimation) that forward-looking behavior played a smaller role during the high and volatile inflation regime before 1981 than in the period of moderate inflation afterwards. Kang et al. (2009) employ an unobserved component model for inflation with Markov switching parameters, confirming the claim that inflation persistence does indeed change across policy regimes. They find a break around the collapse of Bretton Woods in the early 1970s and another around 1981 with the Volcker disinflation. Cogley et al. (2010) obtain similar results using VAR with drifting coefficients and stochastic volatility (unlike most other studies they analyze the inflation gap, measured as the difference between inflation and its trend). On the contrary, Stock and Watson (2007) argue, on the basis of an unobserved component model with stochastic volatility, that the US inflation persistence has not changed for decades. D’Agostino et al. (2011) provide evidence that explicit modeling of structural changes in inflation dynamics (within a time-varying VAR framework) can improve the accuracy of inflation forecasts.

The evidence on changes in inflation dynamics in other economies, especially within the NKPC framework, is less abundant. There are numerous studies initiated by the ESCB Inflation Persistence Network¹, but they mainly use micro data and do not explicitly test the NKPC. There are only a few papers tracking the issue of overall inflation dynamics. Benati (2008) uses data for several developed inflation-targeting countries (Canada, New Zealand, Sweden, Switzerland, and the UK) and the euro area and concludes that inflation persistence decreased almost to zero once credible monetary regimes had been implemented and, therefore, that inflation persistence is not structural. Hondroyannis 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 (using the present-value formulation of the model in a rolling-window regression). He finds that the explanatory power of the model varies substantially across the underlying monetary regimes, 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.

Research focused on inflation dynamics and NKPC estimation in CE countries has been gradually expanding. However, the issue of possible structural changes, which seems to be highly relevant in this case, has not been explicitly tackled yet. The time-invariant estimates provide rather ambiguous evidence on the fit of the NKPC. Arlt et al. (2005) reject the validity of the pure NKPC for the Czech economy using cointegration-based tests. Franta et al. (2007) conclude that inflation in three CE countries is more persistent than in the EMU and that the NKPC proposed in GG is not consistent with the data for any of the countries analyzed. Plašil (2011) estimates the NKPC for the Czech Republic by making use of advances in the area of optimal instrument selection, time series factor analysis, and the GMM bootstrap. He finds some support for the hybrid NKPC. Vašíček (2011) estimates a hybrid NKPC augmented for open economies for four CE countries. He confirms higher persistence of inflation in CE countries and finds that the common measures of the marginal cost perform worse than the output gap and that external rather than internal factors seem to drive inflation. Mihailov et al. (2011b) test a small economy NKPC proposed in Galí and Monacelli (2005) using data from twelve new EU member states. Although they find rather mixed evidence on the importance of external factors, the fit of this model is better for the NMSs than for the developed OECD economies (Mihailov et al., 2011a). Basarac et al. (2011) estimate the NKPC for a panel of nine new EU countries and obtain a measure of expected

¹ http://www.ecb.int/home/html/researcher_ipn.en.html

inflation directly from consumer surveys using a probability method. They confirm that inflation in these countries is very persistent. Hondroyiannis et al. (2008) provide some evidence for a group of seven new EU member states based on a time-varying model. Rather surprisingly, they find that the inflation persistence in these countries is practically nonexistent (and therefore similar to the euro area), which contradicts practically all the country-specific, though time-invariant, evidence. Moreover, their panel estimation for a 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.

3. Model and Estimation Strategy

3.1 Closed and Open-economy Hybrid NKPC

In our empirical analysis we start with the seminal hybrid NKPC model laid out in GG:

$$\pi_t = \gamma_f E_t \pi_{t+1} + \gamma_b \pi_{t-1} + \lambda s_t + \varepsilon_t, \quad (3.1)$$

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. 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 all the CE countries can be classified as small open economies, we also consider an NKPC model in the spirit of Galí and Monacelli (2005), which accounts 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 can be viewed as an extension of their approach to the hybrid NKPC and time-varying framework. In line with the open-economy model, we now assume that *CPI* inflation can be expressed as:

$$\pi_t = \pi_{H,t} + \alpha \Delta TT_t, \quad (3.2)$$

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 the openness of the economy. Analogously to (3.1), the dynamics of domestic inflation are given by:³

$$\pi_{H,t} = \gamma_f E_t \pi_{H,t+1} + \gamma_b \pi_{H,t-1} + \lambda_\alpha s_t. \quad (3.3)$$

Plugging (3.3) into (3.2) and making use of the fact that $\pi_{H,t} = \pi_t - \alpha \Delta TT_t$, we get:

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

After some rearranging we obtain a hybrid open-economy NKPC model of the form:

$$\pi_t = \gamma_f E_t \pi_{t+1} + \gamma_b \pi_{t-1} + \lambda_\alpha s_t + \alpha \{ \Delta TT_t - \gamma_f E_t \Delta TT_{t+1} - \gamma_b \Delta TT_{t-1} \}. \quad (3.4)$$

To motivate economic interpretation of the term in curly brackets in (3.4), it is useful first to consider the two extreme cases when either γ_f or γ_b is equal to one.⁴ If $\gamma_f = 1$, the bracketed term becomes $(\Delta TT_t - E_t \Delta TT_{t+1})$ and model (3.4) 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

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 use identical symbols for the forward and backward-looking terms, although they are not necessarily equal to their closed-economy counterparts. We leave out the error term for expositional ease.

⁴ Although we do not impose the restriction $\gamma_f + \gamma_b = 1$ a priori, the results usually show close-to-convexity properties.

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, as implied by $\gamma_b = 1$, the bracketed 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 the 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, with a slight simplification, the linear⁵ combination 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 than 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 level of goods and services prices or as the relative price of imports in terms of exports.

3.2 Econometric Framework

Models (3.1) and (3.4) 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 (3.1) into the following form⁶:

⁵ If one restricts the coefficients to sum to 1, it is also a convex combination with the straight interpretation of a weighted average.

⁶ For the sake of brevity, we only describe our estimation strategy for the basic NKPC model. The open-economy version is estimated analogously. We make explicit reference to the open-economy model only if this is necessary to avoid confusion.

$$\pi_t = \gamma_f \pi_{t+1} + \gamma_b \pi_{t-1} + \lambda s_t + e_t, \quad (e_t \equiv \varepsilon_t - \gamma_f \vartheta_{t+1}). \quad (3.5)$$

To obtain time-invariant parameter estimates in model (3.5) one usually resorts to GMM techniques. Since a GMM methodology with time-varying coefficients has not yet been fully developed,⁷ we broadly stick 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, we run the OLS regression⁸ of endogenous variables on a set of instruments that are uncorrelated with the error term in (3.5) and store the standardized residuals. In the second step, the standardized residuals are added as additional regressors into (3.5) 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), in particular for inflation in CE countries. Applying methods that ignore possible variation in the volatility of the error term may lead to serious bias of the estimated time-varying coefficients.

To address these issues we slightly modify the 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 in the second step. 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

⁷ See Partouche (2007) for a valuable breakthrough in this area.

⁸ Kim (2006) assumes that the relation between endogenous variables and instruments is time-invariant. Kim (2008) also considers other alternatives. Notably, one can also assume that the relation between endogenous variables and instruments is time-varying. For reasons that will become clear later we do not adopt this approach here.

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). 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:

$$y_t = a + Z_t\delta + \epsilon_t \quad \epsilon \sim N(0, \sigma^2), \quad (3.6)$$

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 $M = [M_1, M_2, \dots, M_K]$, 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:

$$\hat{y}_t^{BMA} = \sum_{k=1}^K \hat{y}_{t,k} p(M_k|y, Z), \quad (3.7)$$

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:

$$p(M_k|y, Z) = \frac{p(y|M_k, Z)p(M_k)}{p(y|Z)} = \frac{p(y|M_k, Z)p(M_k)}{\sum_{s=1}^K p(y|M_s, Z)p(M_s)}, \quad (3.8)$$

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).⁹ Before running BMA, the researcher needs to specify the model universe

⁹ 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.

(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 (3.5)¹⁰ 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.

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).¹¹

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 (3.5) 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):

¹⁰ 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.

¹¹ 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 2^{20} .

$$\pi_t = c_t' \kappa + x_t' \alpha_t + \psi_t, \quad \psi_t \sim N(0, \sigma_t^2) \quad (3.9)$$

$$\alpha_{t+1} = \alpha_t + u_t, \quad u_t \sim N(0, \Sigma) \quad (3.10)$$

$$\sigma_t^2 = \gamma \exp(h_t) \quad (3.11)$$

$$h_{t+1} = \rho h_t + \eta_t, \quad \eta_t \sim N(0, \sigma_\eta^2), \quad (3.12)$$

where $c_t \equiv (v_{t,\pi}^*, v_{t,gap}^*)'$ is a vector of the endogeneity correction terms, $x_t \equiv (\pi_{t+1}, \pi_{t-1}, s_t)'$ is a vector containing key model covariates, κ is a vector of constant parameters, and $\alpha_t \equiv (\gamma_{f,t}, \gamma_{b,t}, \lambda_t)'$ 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 changes and/or structural breaks in the coefficients. Disturbances in (3.9), 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 (3.9)-(3.12) 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.¹² To obtain the results, we drew $M = 55,000$ samples from the posterior distribution and discarded the first 5,000 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. In all the estimated models the inefficiency factors were usually quite low (well below 50). Occasionally, however, they reached values close to 200 for some coefficients (close to 100 for tighter priors). Nevertheless, this still implies that we get about $M/200 = 250$ uncorrelated samples, which is considered enough for posterior inference (see Nakajima, 2011).

As indicated above, one might also be interested in the structural parameters of the NKPC model. Note that their direct estimation leads to a system of equations which are highly non-linear in parameters. Since under quite mild conditions there exists one-to-one mapping between the reduced-form coefficients and the structural parameters, we

¹² 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 used for the estimation.

avoid direct estimation of the structural parameters and instead use a non-linear solver to obtain their value from the estimated reduced-form coefficients.¹³

3.3 Data

Our dataset combines time series taken from several data sources (ECB, Eurostat, OECD, and IMF). They were all 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 1995 Q1 (CZ) and 1996 Q1 in Hungary and Poland, respectively, 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 Outlook¹⁴ rather than the commonly used unit labor cost (labor share of income). 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 instrument set includes (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.¹⁵ 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 note 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

¹³ We fixed the subjective factor β to 0.99.

¹⁴ It seems to correspond by and large to the output gap obtained by the HP filter.

¹⁵ We resort to this rather simplistic definition due to the limited availability of other interest rate data in the given period.

1999 Q1–2010 Q4 and re-estimated models (3.1) and (3.4).¹⁶ Given that the overall results remained largely identical we report the outcomes for the longer time span only.

4. Results

4.1 Benchmark NKPC (GG, 1999; GGL, 2001) with Time-varying Parameters

Czech Republic

Figure 1 presents the estimated time-varying reduced-form coefficients for the Czech Republic. In general terms we can observe that Czech inflation is mainly a forward-looking process. While the coefficient γ_f oscillated between 0.6 and 0.7 between 1995 and 2005, we can observe a slight upward tendency since 2004, reaching the value of 0.8 recently. The increase is particularly pronounced since the onset of the global recession in 2008, which is consistent with recent evidence for the euro area (Koop and Onorante, 2011). On the other hand, the backward-looking term γ_b has decreased over time, from 0.3 to less than 0.2. Moreover, since 2003 the estimates of γ_b are insignificant, with the exception of the first quarter of 2008, when inflation jumped up due to the combined effect of increased food and energy prices and a hike in value added tax.¹⁷ A possible explanation for the increase in the backward-looking parameter in response to the increase in VAT might be the following: an increase in VAT increases the volatility of inflation, which translates into higher uncertainty about the future path of inflation. Thus, as the formation of inflation expectations becomes more complicated, both firms and households pay higher attention to past inflation rather than to possibly biased forecasts. On the other hand, it seems that inflation expectations were firmly anchored and the effect of this shock was rather time limited.

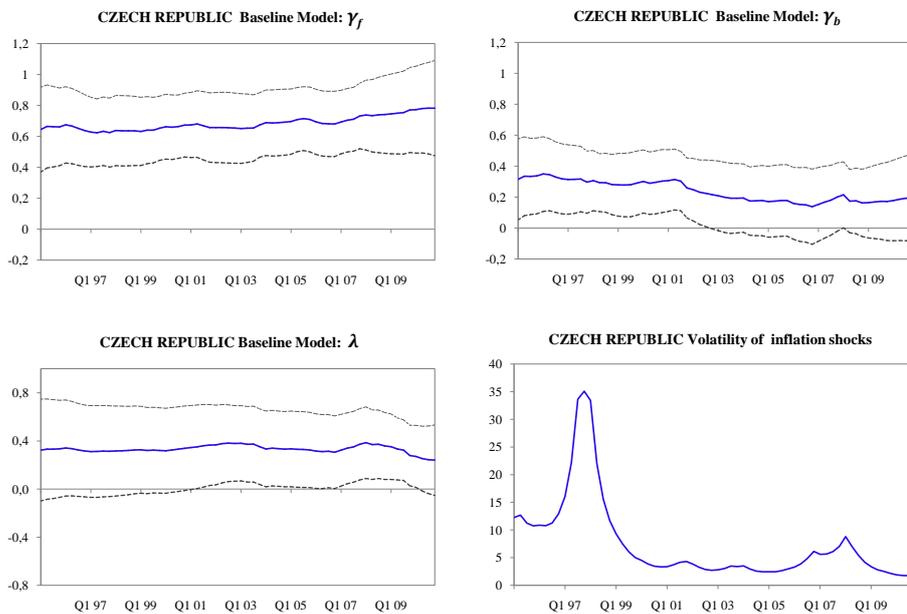
Interestingly, we do not observe any peak or change in trend around 1998, when inflation targeting was adopted by the Czech National Bank. However, several years after, a clear decline in the value of γ_b appears. The decrease of inflation persistence started in the last quarter 2001, and by the beginning of 2003 the coefficient γ_b had dropped by 0.1. This decrease appeared after the inflation rate had slumped significantly below the inflation target from its previous values of between 4% and 6%. This disinflation appeared shortly after the Czech National Bank changed its approach to inflation target setting. In particular, the CNB decided to move from periodic setting of targets for the

¹⁶ Other variables were HP filtered, if necessary, to achieve stationarity.

¹⁷ The estimated impact of lagged inflation is in accordance with Babetskii et al. (2007), who estimated the inflation persistence on disaggregated data at roughly between 0.2 and 0.3 depending on the time range included in the sample. Their results for the aggregate CPI were slightly larger, but still below 0.5. From this perspective, it seems that either the results from the time-varying model do not suffer from aggregation bias, or the micro data on inflation persistence suggest that inflation persistence could have disappeared.

end of the year in terms of the net inflation rate, to continuous targeting of headline inflation within a predefined target range. 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 on the inflation dynamics 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 1: Czech Republic: Reduced-form Coefficients



The coefficient λ , measuring the impact of real economic activity on the inflation rate, is insignificant up to 2001. This seems to suggest that the effects of the large disinflation in the early years of the transition dominated the effects of real economic activity. Since then, the coefficient has been positive and significant, with the exception of the very last quarters of the sample.

The last subplot shows the estimated volatility of inflation shocks. The first of the two conspicuous peaks in volatility may be associated with the depreciation of the Czech koruna following the currency crisis in mid-1997 as well as administrative changes in regulated prices. The second peak may be linked to an increase in import prices and value added tax in 2007/2008. Clearly, these policy shocks were short lived and do not seem to have affected the properties of the inflation dynamics.

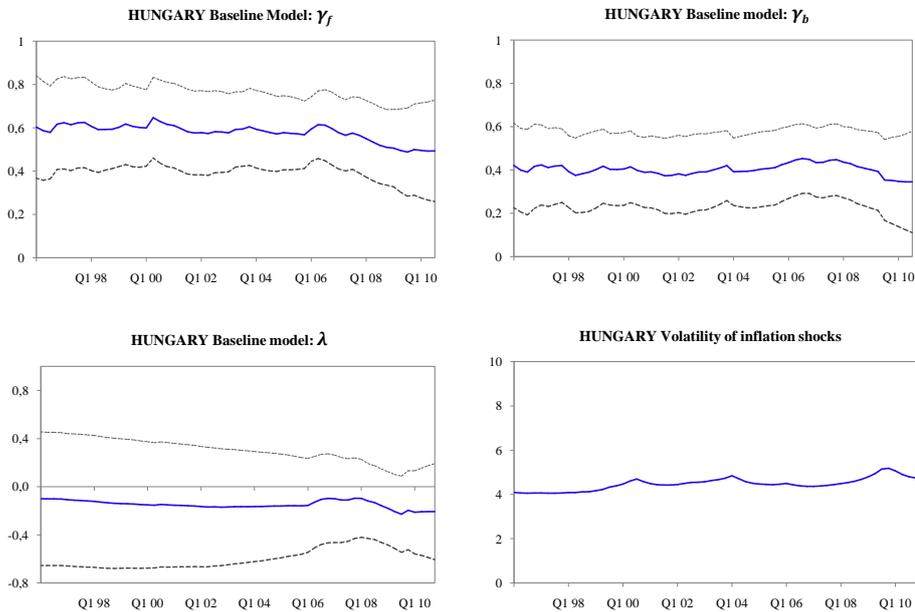
To sum up, the evidence from the time-varying coefficients suggests that the characteristics of the inflation process in the Czech Republic have converged to those in developed countries. A predominantly forward-looking nature of the inflation process is commonly reported in recent studies on the US and large EU countries (Hondroyiannis et al., 2009; Cogley and Sbordone, 2008; Benati, 2008). These studies argue that inflation turns mainly forward-looking once the inflation rate has stabilized under a credible monetary policy regime. The timing of the gradual decrease in the backward-looking term suggests that it was jointly caused by a low inflation rate and a simultaneous switch to a more transparent inflation-targeting regime that anchored inflation expectations at low levels.

Hungary

The inflation dynamics in Hungary show a rather different picture to those in the Czech Republic. First, the forward-looking term γ_f is relatively stable over time, with only a slight decrease since 2007. The backward-looking term γ_b does not decrease and is significant over the whole time span. Hence, inflation persistence still seems to be an important phenomenon in Hungary. Our results with a significant forward-looking term correspond to Menyhért (2008), who provides the first evidence of a significant forward-looking term in Hungary. However, it must also be noted that the coefficient λ is negative, though insignificant, on the entire sample, with weak evidence of an increase since 2007. This result either can be explained by poor reliability of the output gap estimate for Hungary, or may reflect the fact that the output gap is not a driving factor of inflation, as suggested by the NKPC model. In addition, we tested alternative domestic forcing variables such as the unit labor cost and the unemployment rate with very similar results.

The stability of the coefficients is somewhat surprising, as, at least formally, monetary policy in Hungary has changed significantly since 1996. Inflation targeting was adopted in 2001¹⁸ and at the same time the crawling band was replaced by the ‘shadow’ ERM II regime of a fixed exchange rate with a fluctuation band of +/-15% around the central parity against the euro. However, despite the announcement of the shadow ERM II exchange rate regime, several pro-inflationary depreciation periods followed. The most significant one in terms of its effect on inflation occurred in 2004, when inflation increased from 3% to 7%. Nevertheless, the formal changes in monetary policy did not lead to lower inflation persistence and the volatility of inflation also remained high. This seems to be at odds with evidence for advanced countries (Benati, 2008) where implementation of inflation targeting led to the significant decrease in inflation persistence. Arguably, this might be partially attributed to role of the exchange rate in monetary

¹⁸ 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 in 2008.

Figure 2: Hungary: Reduced-form Coefficients

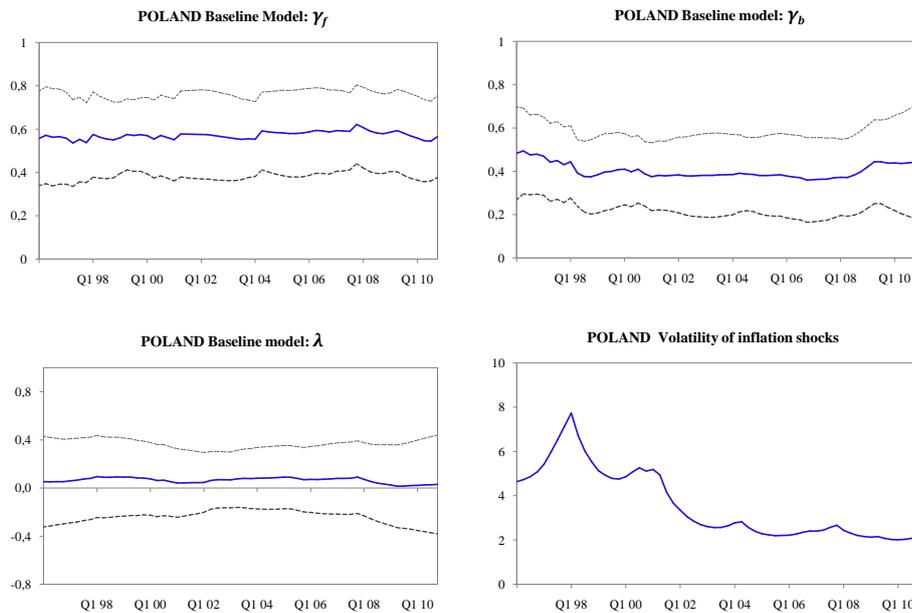
policy transmission in Hungary (Vonnak, 2008). This is also supported by the fact that the exchange rate was identified as one of the most important factors of inflation expectations (see the Appendix). Hence, despite the adoption of inflation targeting and the diminishing effects of transition, we cannot identify any important changes either in the parameters of the NKPC or in the volatility of shocks.

Poland

Poland adopted explicit inflation targeting in 1999. Its inflation persistence had already decreased prior to this date, as indicated by the backward-looking term γ_b , which moved from values close to 0.5 to below 0.4. Between 2001 and 2009 the coefficient γ_b stayed below 0.4 with no tendency to move and the forward-looking term γ_f remained stable as well, with just small variations between 0.55 and 0.60. The overall dynamics of inflation were stable, with both the forward and backward-looking components playing a significant role. The observed stabilization of the driving factors of the inflation dynamics can be linked to stabilization of inflation expectations: Lyziak (2003) and Orłowski (2010) document that inflation expectations became anchored to the target path about two years after inflation targeting was adopted by the National Bank of Poland, that is, in 2001/2002. Correspondingly, our results suggest that the volatility of inflation shocks decreased sharply, and since 2001 inflation in Poland has been characterized by

a stable monetary policy regime, stable coefficients of the NKPC, and low volatility of inflation shocks. However, the estimate of λ is close to zero (never exceeds 0.1) and never significant. Hence, the dynamics of inflation, when the closed-economy specification of the NKPC is considered, are not driven by (the estimate of) the output gap. Again, this points to some problems with finding a good proxy for firms' marginal cost or, potentially, to empirical failure of the model.

Figure 3: Poland: Reduced-form Coefficients



Overall Assessment of Results

In general, we find evidence that the forward-looking inflation term is more important than the backward-looking one. This implies that inflation expectations play a substantial 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, so that the central banks do not need to rely on interest rate changes only.

When the time averages of the estimated time-varying coefficients are considered (Table 1), the coefficient for expected inflation γ_f is significant at two standard deviations¹⁹ in all three countries. The backward-looking term is lower and significant at one standard

¹⁹ The standard deviation is again calculated as the time average of its time-varying counterpart.

deviation only. The coefficients γ_f and γ_b for Poland and Hungary are similar, with γ_f close to 0.55 and γ_b slightly over 0.4. The results for the Czech Republic are somewhat different, suggesting higher importance of inflation expectations for the overall inflation dynamics, with the time average of the estimated coefficient γ_f at 0.68. Correspondingly, the role of the backward-looking term is lower, as the associated coefficient γ_b is equal to 0.24. Hence, the Czech inflation persistence is roughly one half that in other countries.²⁰

Table 1: Time Averages of Estimated Coefficients (Reduced-form NKPC)

	Czech Republic	Poland	Hungary
γ_f	0.68 (0.23)	0.57 (0.19)	0.58 (0.19)
γ_b	0.24 (0.22)	0.40 (0.18)	0.40 (0.18)
λ	0.33 (0.33)	0.06 (0.30)	-0.15 (0.45)

The impact of the output gap on inflation is rarely significant over the entire sample. In fact, only in the case of the Czech Republic is the estimated coefficient λ significant at one standard deviation. This result may have a number of explanations. First, the Phillips curve might flatten out at lower levels of inflation, as the relationship between output or unemployment and inflation is likely to be non-linear. This idea already appears in the original article by Phillips (1958) and has been acknowledged by a number of authors (an overview can be found in Stock and Watson, 2010). Second, the slope of the Phillips curve might depend on the size of the output gap: in normal times without recessions and with only mild output gaps, the relationship implied by the Phillips curve is small, but when larger recessions occur, the curve steepens (Stock and Watson, 2010). All the countries in our sample were characterized by relatively low volatility of growth rates, and although the output gap was negative in a number of periods, output growth remained positive or decreased by few percentage points below zero (with the exception of the current crisis). Third, and perhaps most importantly within the context of the CE countries represented in our sample, the low λ may be associated with factors specific to transition countries (such as the fading impact of changes in regulated prices) that cause shifts of the Phillips curve rather than movement along it. Since the volatility in the time-varying λ is relatively small and no systematic correlations with either inflation or the output gap appear, the hypothesis of shifts of the Phillips curve in response to supply shocks seems to be the most likely. However, the first-step BMA results suggest that

²⁰ Note that the time averages can be interpreted as being a result of a 2SLS model with stochastic volatility, where the first-step equation is replaced by a BMA model.

real factors are often relevant for inflation forecasts (inflation expectations), though the relationship may be more complicated than the standard NKPC suggests.

Our time-average results are in line with existing studies that estimate the time-invariant NKPC (Arlt et al., 2005; Franta et al., 2007; Basarac et al., 2011; Danišková and Fidrmuc, 2011; Vašíček, 2011) and confirm the hybrid nature of the NKPC. However, when time-varying responses are allowed, the estimated inflation persistence is somewhat smaller.²¹ It is arguable whether the omission of potential changes in the inflation process implies some upward bias in the backward-looking term; nevertheless, these results are in line with Hondroyiannis et al. (2009). In summary, we show that inflation dynamics differ substantially between the three CE countries and that imposing slope homogeneity in a panel of a rather heterogeneous group of CE countries might not be appropriate. Our findings also confirm the rather ambiguous evidence on the importance of domestic forcing variables. Mihailov et al. (2011a) and Vašíček (2011) provide some evidence that external factors may be more important as inflation forcing variables. In the following section, we extend this evidence to the time-varying framework.

4.2 Open-economy NKPC with Time-varying Parameters

All three countries under study are small open economies highly integrated with international markets, in particular the euro area. The CE countries liberalized their foreign trade in the early 1990s and their integration has increased further since EU accession in 2004. As a result, a large proportion of domestic production is aimed at foreign markets and a major share of both intermediate and final products is imported. Therefore, domestic consumer inflation is probably affected (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. From this perspective the model without terms of trade is likely to be misspecified.

The estimated coefficients of the time-varying open-economy NKPC are shown in Figures 4–6. A key observation is that the dynamics of the reduced-form coefficients remain almost untouched. Second, the coefficient α has certain similarities in terms of dynamics to λ , even though it is usually insignificant.²²

²¹ Additionally, Vašíček (2011) reports estimates of the hybrid NKPC in all countries under study and finds that the forward-looking term dominates the inflation process but lagged inflation is still important. Also, his estimates of λ are close to zero for Poland, often negative for Hungary, and about 0.2 for the Czech Republic, but with high variation across different variables representing economic activity.

²² Given that the bracketed term in equation (3.4) is rather complicated, we checked the robustness of our results using the simple deviation of the terms of trade from the HP-filtered trend, but the results were qualitatively similar to those presented in Figures 4–6.

Figure 4: Czech Republic, Open-economy NKPC Coefficients

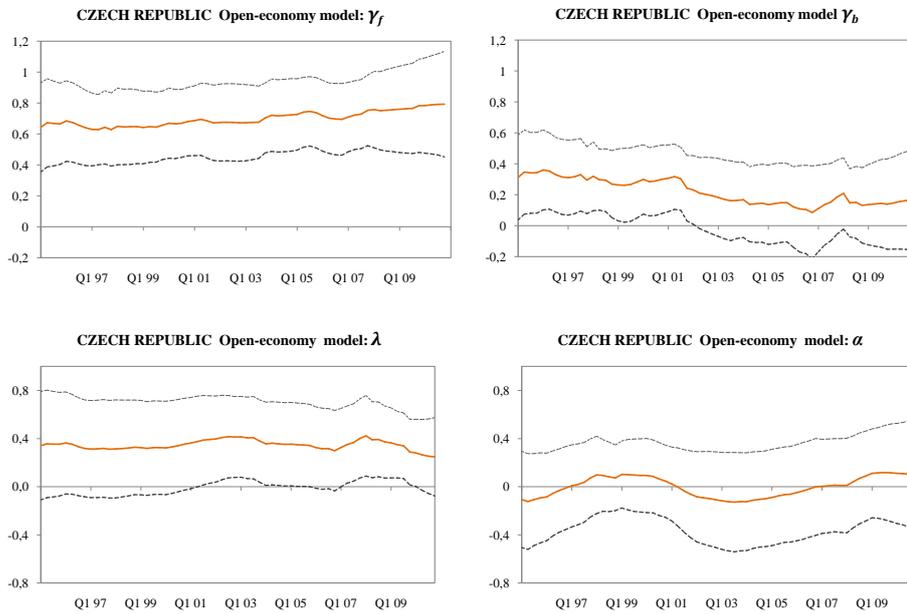


Figure 5: Hungary, Open-economy NKPC Coefficients

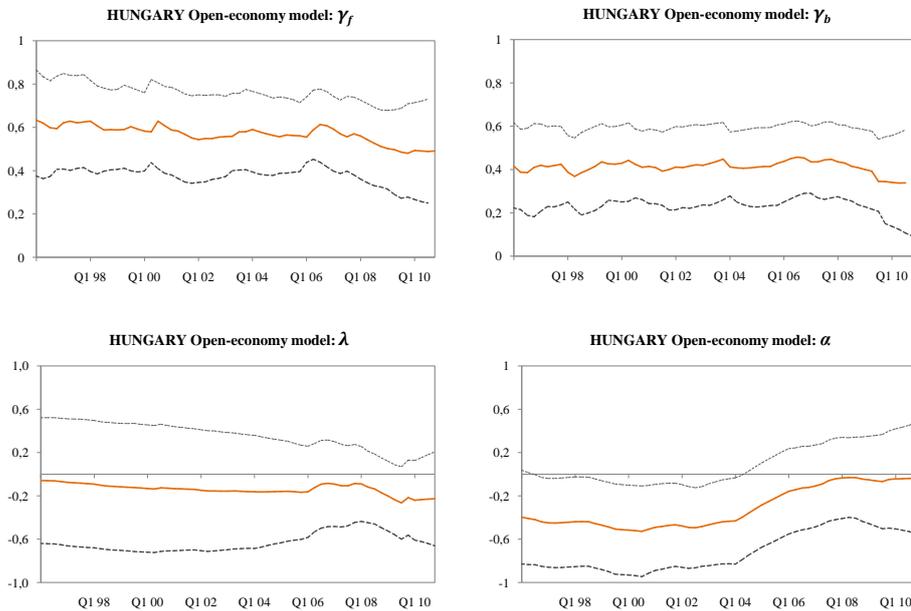
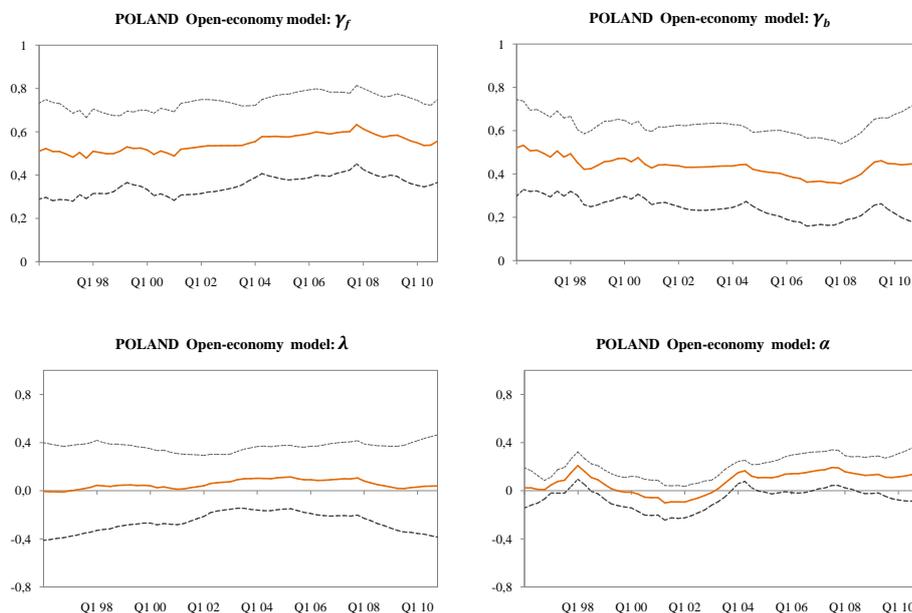


Figure 6: Poland, Open-economy NKPC Coefficients

The presented results suggest that the open-economy NKPC improved only marginally on the benchmark NKPC. The terms of trade seem to be a driving factor of inflation only in Poland, where the coefficient α is significant in several periods following large depreciations (1997Q4–1998Q2, 2003Q4–2004Q3) and again at the onset of the late-2000s recession (2007Q1–2008Q2). In the Czech Republic the terms of trade are not significant and the information from foreign prices seems to be already well reflected in domestic inflation expectations themselves.

As far as Hungary is concerned, the results for the open-economy NKPC are puzzling in a similar manner as for the closed economy model. The estimated coefficient α is negative and significant until 2003, despite the fact that negative values are not allowed by the theory. Note that these rather baffling results are in line with the findings of Mihailov et al. (2011b). Later on, the coefficient α approaches zero and inflation is driven only by its expected and lagged values.

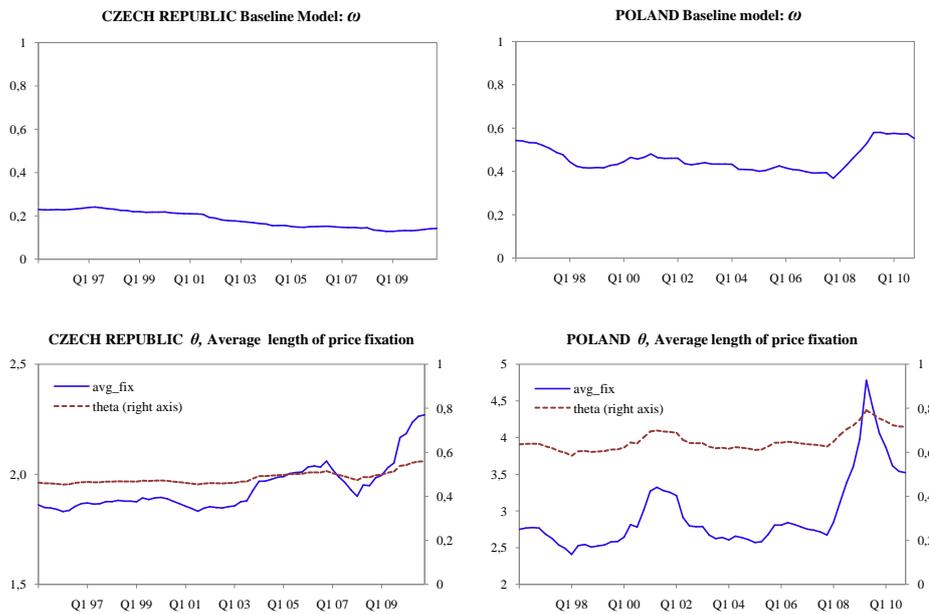
The overall insignificance of the additional term tracking external sources of inflation may seem rather counterintuitive for small open economies. An intuitive explanation is that the factors affecting the terms of trade are already reflected in inflation expectations. For instance, if domestic firms engage in foreign trade, their inflation expectations are influenced by the foreign price level as well as by the exchange rate.

4.3 Are the NKPC Structural Coefficients Truly Structural in CE Countries?

There has been some controversy in the existing empirical work over the structural coefficients in the NKPC model. While they are typically reported in papers based on the time-invariant framework, time-varying studies usually do not go that far. Indeed, 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 in CE countries. Consequently, we use the estimates of the reduced-form coefficients to obtain a sequence of structural coefficients corresponding to the benchmark hybrid NKPC, 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 mentioned in subsection 3.2, the structural coefficients were derived under the assumption of a fixed β equal to 0.99.²³

The results for the Czech Republic and Poland are reported in Figure 7. For Hungary, the reduced-form coefficient of the output gap λ is negative on the whole sample, which prevents us from obtaining structural parameters. It is also important to note, that since estimates of λ are often insignificant, one has to read the results with a great caution.

Figure 7: Structural Parameters, Baseline Model



²³ We do not estimate the structural coefficient from the open-economy NKPC because *i*) mapping between reduced-form and structural coefficients is more complicated in this case, *ii*) while the coefficient of the open-economy component is mostly insignificant, the other coefficients are largely similar to the benchmark case.

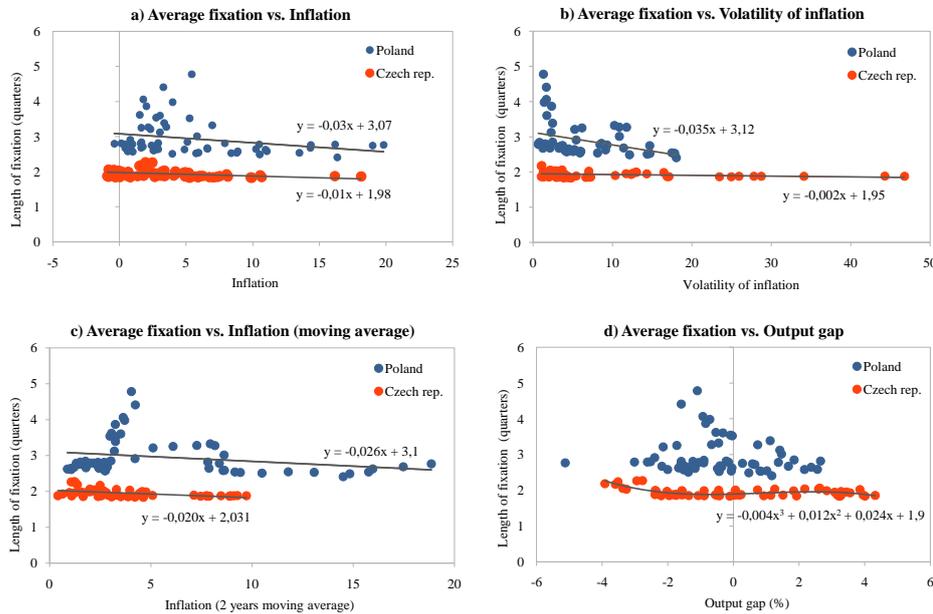
The structural coefficients for the Czech NKPC are depicted in the left panel of Figure 7. Parameter θ was stable until 2003 and has been slowly increasing since then. Correspondingly, the average length of price fixation co-moves. This result may be associated with decreased volatility of inflation, which in turn translates into longer periods when prices do not change. On average, the length of fixation was 1.94 quarters over the sample. The share of the ‘rule of thumb’ firms ω decreased by 10 percentage points from 24% to 14% during the transition, especially after 2001/2002. The decrease corresponds 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 role of the administered prices (which are typically set in a backward-looking manner) in overall inflation decreased. Third, forward-looking price setting is arguably subject to learning. Therefore, a decreasing share of backward-looking price setters may signal that firms are becoming more sophisticated and forming their expectations in a rational (i.e., forward-looking) rather than adaptive (i.e., backward-looking) fashion.

As far as Poland is concerned, the overall stability of the structural coefficients ends at the onset of the late-2000s recession (Figure 7, right panel). All the structural coefficients rose sharply between 2007 Q4 and 2009 Q2, when a peak occurred and the trajectories of the structural coefficients reversed. These dynamics reflect a decrease in the backward-looking term γ_b , which rose due to oil and food prices. However, by the end of our sample period (2010 Q4) the values of the structural coefficients had not returned to their pre-recession levels. The average share of backward-looking (‘rule of thumb’) firms is 46% and the average time for which prices remain fixed is 2.94 quarters. The increase in both structural parameters θ and ω at the end of the sample is clearly linked to the late-2000s recession. A concurrent upward shift in structural parameter θ , capturing price rigidity, and a deep slump in output suggest the existence of downward price rigidities. Nevertheless, the validity of these results needs to be corroborated with microeconomic data.

To assess whether the derived time path in θ or the average length of fixation can be linked to economic intuition, we present scatter plots between the average length of fixation and key macroeconomic variables: inflation, average inflation, the volatility of inflation, and the output gap (Figure 8a–d)²⁴. Economic intuition says that in a situation of higher and more volatile inflation it becomes more complicated for economic agents to distinguish changes in relative prices from changes in the overall price level. Given this hypothesis, a negative relationship should exist between the average length of fixation and the inflation rate or the volatility of inflation (a discussion of this issue can be found in Taylor, 1999). Figure 8a) shows that a negative relationship is indeed observable in the data, as the negative slope is significant for both the Czech Republic

²⁴ In the case of the share of backward-looking firms ω , we argued above that its variation is related rather to institutional changes.

Figure 8: Relation between Average Price Fixation and Macroeconomic Variables



and Poland. There is also a significant negative relationship between the average length of price fixation and the volatility of inflation (measured as the 2-year rolling standard deviation).²⁵

The potential negative relationship between the average length of fixation and the size of the output gap may be related to the presence of downward price rigidities. Figure 8d) shows that especially for the Czech Republic the periods with the largest negative output gaps are truly also the periods with the longest length of fixation. Non-linear regression of the third order supports this claim, as all the estimated coefficients are highly significant and the R-square of the regression is almost 50%. In the case of Poland, evidence of a significant relationship between the size of the output gap and the average length cannot be observed using this simple approach. Nevertheless, even in Poland the periods with the longest price fixation are connected with negative output gaps.

²⁵ In addition, a negative relationship was identified with respect to the moving average of inflation and the estimated stochastic volatility of the residuals.

5. Conclusions

This paper analyzed the dynamics of inflation through the lens of the New Keynesian Phillips curve nested within a time-varying framework. Although originally, the NKPC was proposed as a structural model 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. We aimed to shed some light on this issue by estimating a standard hybrid version of the NKPC and its open-economy counterpart for the CE countries.

The changes in inflation dynamics are usually linked to monetary policy actions. In particular, the recent decrease in inflation persistence was commonly related either to a more aggressive reaction of central banks or to the anchoring of inflation expectations to the long-term inflation target. In the period under study, the countries in Central Europe went through a unique 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 drive down inflation persistence.

In general, we found evidence that the forward-looking inflation term is more important than the backward-looking one. This implies that inflation expectations play a substantial role and are (at least partially) anchored in the CE countries. In this respect, our results favor the hybrid NKPC over specifications without a forward-looking term. However, the nature of the inflation process differs considerably across the selected CE countries. In particular, inflation is substantially less persistent in the Czech Republic than in Hungary and Poland. The almost negligible inflation persistence in the Czech Republic implies that lower inflation can be achieved by successful anchoring of inflation expectations and does not need to be accompanied by output or employment losses. The estimated volatility of inflation shocks decreased quickly a few years after the adoption of inflation targeting both in the Czech Republic and in Poland, whereas it remains stable in Hungary.

The NKPC model postulates that inflation dynamics are determined not only by inflation persistence and inflation expectations, but also by real activity. However, finding the proper forcing variable for inflation in the CE countries has proved to be a nontrivial task. This fact seems to be related to factors specific to transition countries, such as a diminishing impact of administered prices, gradually increasing labor productivity, trade integration with the EU, and higher vulnerability of the CE countries to shocks on financial markets in the 1990s. All these factors lead to a weaker link between the output gap and inflation.

Lastly, we found some evidence that the ‘structural’ coefficients are not stable across time as is commonly believed. We showed that the average time for which prices remain fixed is negatively correlated with both the level and the volatility of inflation. The share of backward-looking price setters is changing smoothly, with a predominantly downward-sloping trend. Unlike in the case of the average fixation, the reasons should be sought in long-term determinants such as increasing competition, decreasing administered prices, and the learning capacity of price setters.

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Appendix: BMA results (first step)²⁶

Figure A1: Czech Republic

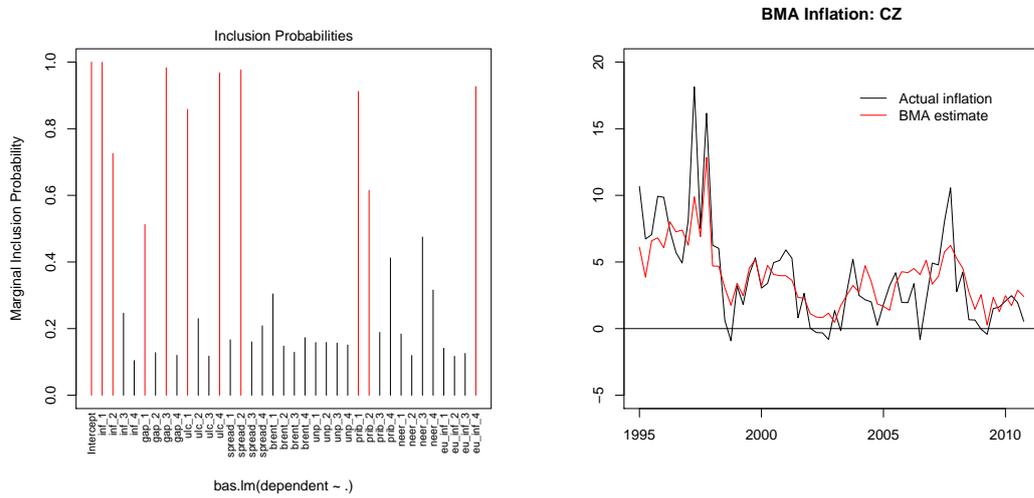
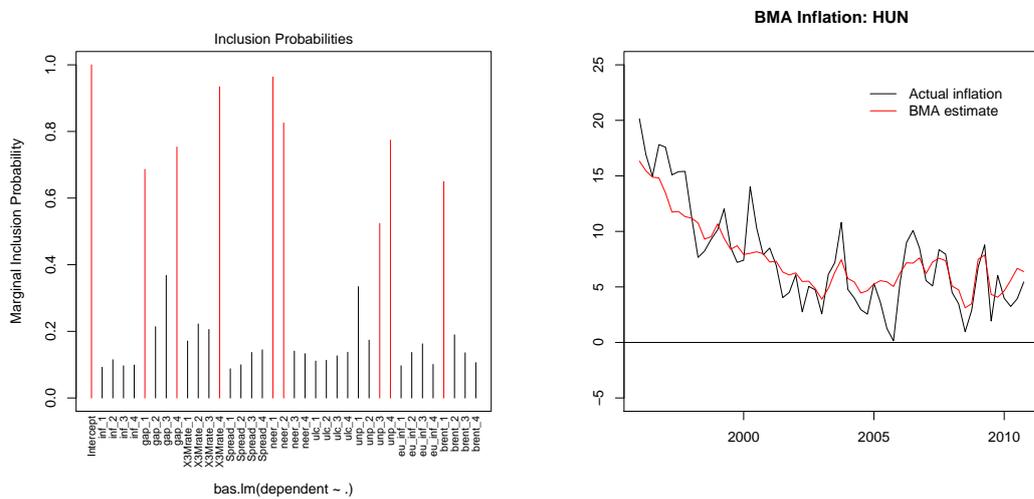
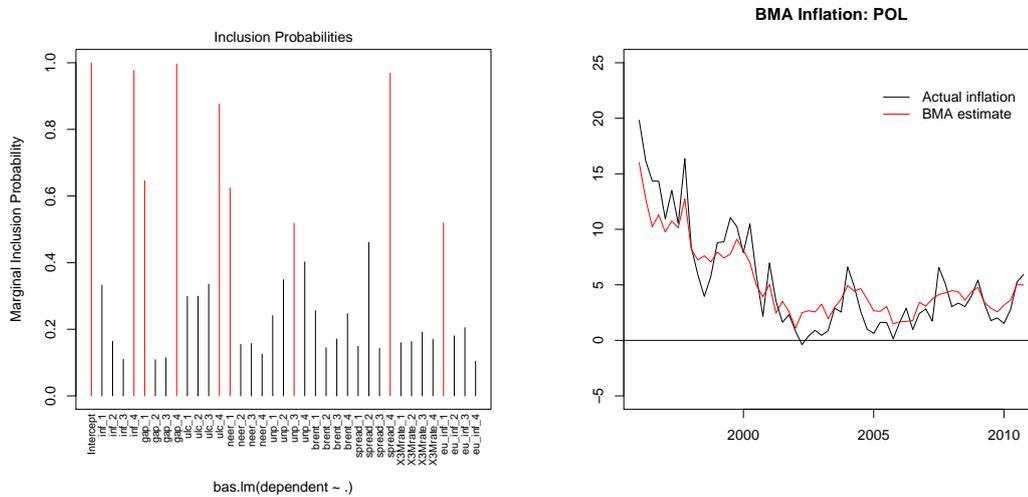


Figure A2: Hungary



²⁶ Red bars indicate posterior inclusion probability higher than 0.5.

Figure A3: Poland



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