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The Hybrid Phillips Curve: Empirical Evidence from Transition Economies*

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Abstract
In this paper we estimate the hybrid New Keynesian Phillips curve for nine transition economies and examine its ability to explain inflation dynamics. Special emphasis has been made on obtaining a measure of expected inflation directly from consumer surveys via the probability method, as opposed to most similar studies, which employ various proxy or instrumental variables for expected inflation. Unlike similar studies that employ the Generalized Method of Moments in evaluating the hybrid Phillips curve, here we use a dynamic fixed effects (DFE) model, as suggested by recent advances in the estimation of nonstationary heterogeneous dynamic panel models. This empirical investigation leads to the conclusion that there does exist a cointegration relation between inflation, expected inflation, and the output gap (as a proxy for real marginal cost). The long-run coefficients for both independent variables are positive and statistically significant. Moreover, based on the error correction model evaluated, one arrives at a conclusion that the error correction term is statistically significant and of appropriate sign, pointing to a 15 percent quarterly imbalance correction. Furthermore, our results are robust to a variety of dynamic panel estimation procedures.

1. Introduction

Important advances have emerged recently in the modeling of inflation dynamics. Specifically, the Phillips curve has evolved significantly since Phillips' seminal 1958 paper[1] and its empirical validity has been re-examined and reviewed in a number of papers. So, in the modern Phillips curve specification, Galí and Gertler (1999) on US data and Galí, Gertler, and López -Salido (2001, henceforth GGLS) on euro area data demonstrated the role of both forward-looking and backward-looking behavior in the inflation process.

The most widely accepted model of the Phillips curve in modern macroeconomics is the hybrid New Keynesian Phillips curve (NKPC). It is derived from a New Keynesian model characterized by monopolistic competition and short-run price rigidity and it is hybrid in the sense that it contains past inflation. This assumption of backward-looking price-setting leads to inertia in inflation behavior. The baseline model was developed by Galí and Gertler (1999), who estimated a hybrid variant of the NKPC that relates inflation to expected inflation, lagged inflation, and real marginal cost.

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[1] A. W. Phillips wrote an article in 1958 entitled "The Relationship Between Unemployment and the Rate of Change of Money Wage Rates in the United Kingdom, 1861–1957".

In this paper, a hybrid New Keynesian Phillips curve is empirically tested for nine European transition economies (the Czech Republic, Estonia, Latvia, Lithuania, Hungary, Poland, Romania, Slovenia, and Slovakia) applying the panel data method on quarterly data for the period 2002q2 to 2009q2. In quantifying consumers' inflation expectations, the probability method is applied. Milestone papers concerning the NKPC usually employ official, publicly available expected inflation measures: Paloviita (2006) applies OECD forecasts for the euro area; Paloviita and Mayes (2005) use real-time information on expectations; and Zhang, Osborn, and Kim (2006) apply the Michigan Survey, the Greenbook and the Survey of Professional Forecasters for the US. As opposed to that, institutions conducting consumer surveys in transition countries do not publicly offer such indicators on a monthly/quarterly basis. For this reason, similar NKPC analyses for developing European countries use actual inflation realizations or other instrumental variables to account for expected inflation (Dabušinskas and Kulikov, 2007; Hondroyiannis et al., 2008; Vašíček, 2010).

Our extensions to the previous research are several. First, instead of imposing rational expectations, here we employ an up-to-date methodology proposed by Forsells and Kenny (2004) to extract an expected inflation measure from consumer surveys in the transition countries of interest. Second, our empirical approach differs from that of previous authors. Most empirical studies of the hybrid NKPC make use of the Generalized Method of Moments (GMM). One of the central findings from datasets with a reasonably large T is that the assumption of homogeneity of slope parameters is often inappropriate. The GMM approach can produce inconsistent and potentially very misleading estimates of the average parameter values in dynamic panel models unless the slope coefficients are in fact identical (Pesaran and Smith, 1995). We avoid this problem by using a pooled mean-group (PMG) estimator for dynamic heterogeneous panel data. Generally, GMM estimators can be used in the case of stationary panel data (Arellano and Bond, 1991; Blundell and Bond, 1998), while PMG can be used in the case of nonstationary regressors (Pesaran, Shin, and Smith, 1999). PMG outperforms other estimators in situations of large T and heterogeneous panel data (Pesaran, Shin, and Smith, 1999). Third, based on a large and recent dataset we demonstrate a coherent multi-country analysis that enables certain common conclusions and international analogy.

The remainder of this paper is organized as follows. The relevant literature is briefly discussed in the next section, which starts with a brief account of the development of theoretical and empirical modeling of the hybrid New Keynesian Phillips curve and then addresses some important empirical issues and general aspects of the estimation procedure. The third section provides an explanation of all the variables used in the model. Special attention is given to quantifying consumers' inflation expectations (obtained by surveys) via the probability method and to a thorough discussion of the panel data methodology. The fourth section analyzes and interprets the estimation results of the hybrid NKPC for the observed transition countries based on a dynamic panel model. Finally, we conclude that the hybrid NKPC with the output gap as an explanatory variable performs in accordance with the theory in all the observed countries.

2. The Hybrid Phillips Curve: Theory and Evidence

A voluminous literature demonstrating the role of the relationship between inflation and real activity and estimating different Phillips curve specifications has been produced over the past 50 years. In that sense, the empirical implementation of the forward-looking New Keynesian Phillips curve has been the subject of notable controversy. To analyze how well the New Keynesian Phillips curve captures inflation inertia, Galí and Gertler (1999) develop and estimate a model of inflation that allows for a fraction of firms that use a backward-looking rule to set prices. It is called a hybrid variant of the New Keynesian Phillips curve. Consequently, they extend the basic Calvo model (1983).

Therefore, Galí and Gertler (1999) assume that, as in Calvo's model, each firm has a probability $1-\theta$ of being able to reset its price in any given period, independently of the time elapsed since its most recent price adjustment. Thus, a fraction θ of firms, which we refer to as "forward-looking", keep their prices unchanged in any given period.Unlike Calvo, from those firms able to adjust prices in a given period, only a fraction $1-\varpi$ set prices optimally. The remaining firms, measured by ω, which we refer to as "backward-looking", use a simple rule of thumb instead, that is, they set prices equal to the average of newly adjusted prices last period plus an adjustment of expected inflation, based on lagged inflation π_{t-1}.

The recent advancement in the NKPC model has been developed from earlier sticky price models in the spirit of the staggered contracts models developed by Taylor (1980) and Calvo (1983) and the quadratic price adjustment cost model of Rotemberg (1982). Galí and Gertler augment this New Keynesian Phillips curve with a backward-looking element. Accordingly, current inflation is expressed as a function of expected future inflation, lagged inflation, and real marginal costs. The hybrid Phillips curve is given by:

$$\pi_t = \lambda mc_t + \gamma_f E_t\left\{\pi_{t+1}\right\} + \gamma_b \pi_{t-1} + \varepsilon_t \tag{1}$$

where

$$\lambda = (1-\omega)(1-\theta)(1-\beta\theta)\varphi^{-1} \tag{2}$$

$$\gamma_f = \beta\theta\varphi^{-1} \tag{3}$$

$$\gamma_b = \omega\varphi^{-1} \tag{4}$$

with $\varphi = \theta + \omega\left[1-\theta(1-\beta)\right]$ and error term ε_t. In this specification, mc_t is (log) real marginal cost and all the coefficients are explicit functions of three model structural parameters: θ, which measures the degree of price stickiness, ω, which is the degree of "backwardness" in price setting, and the subjective discount factor β.[2]

There has been an extensive discussion in the literature about the correct proxy for real marginal costs [3] as the relevant real sector driving variable. One can employ the output gap (commonly used in Phillips curve regressions) or the labor income share, as proposed by Galí and Gertler (1999), who emphasize that under certain conditions[4] there is an approximate log linear relationship between the output

[2] See also Galí and Gertler (1999) for a discussion of some of the issues involved.

[3] See, for example, Galí and Gertler (1999), Galí et al. (2001), and Sbordone (2005).

gap and marginal cost. Furthermore, the authors emphasize that conventional measures of the output gap are likely to be ridden with error, primarily due to the unobservability of the natural rate of output. Hence, they suggested the use of the labor income share. On the other hand, Rudd and Whelan[5] (2005) do not find the labor income share to be statistically significant, although it does have the correct positive sign.

Although there is a growing recognition in the literature that the New Keynesian Phillips curve is a widely used tool, the empirical performance of the model has often been the object of criticism.[6]

Dabušinskas and Kulikov (2007) present an empirical analysis of the inflation process in Estonia, Latvia, and Lithuania within the framework of the NKPC with an open economy extension. Their analysis covers the years 1995 to 2005. The empirical results suggest that the inflation process in these countries primarily depends on inflation expectations and past inflation rates. They also find that the real marginal cost measure plays a limited role in determining inflation dynamics. Franta et al. (2007) analyze two country groups – the new EU member states (NMS)[7] and the current euro area countries. The authors analyze statistical and structural measures of inflation persistence. According to their findings, inflation persistence in the NMS group is comparable to that in the euro area countries. Franta et al. stress that the superiority of time-varying mean models suggests that expectations have been an important source of inflation persistence in the NMS. However, their estimates of the hybrid New Phillips curve indicate that inflation behavior is still more backward--looking in the NMS than in the current euro area countries.

Hondroyiannis et al. (2008) provide evidence on the NKPC for the euro area and a group of seven new member countries[8] that joined the European Union (EU) in

[4] Galí and Gertler (1999, p. 199) claim that: "In the standard sticky price framework without variable capital (e.g., Rotemberg and Woodford, 1997), there is an approximate proportionate relation between marginal cost and output. With variable capital the relation is no longer proportionate. Simulations suggest, however, that the relationship remains very close to proportionate."

[5] Rudd and Whelan (2005) employ both the output gap and the labor share of income as driving variables. Their estimates of the effects of the driving variables are different from Galí and Gertler's. They find that the coefficients on the output gap are insignificant and have the "wrong" sign and do not find the labor income share to be statistically significant either. Furthermore, Henzel and Wollmershäuser (2008) claim that real marginal costs are a linear function of the output gap under the assumption that labor market frictions exist, but do not vary over time.

[6] An early critical assessment can be found in Fuhrer and Moore (1995). Mankiw (2000, p. 24) argues that the NKPC is not consistent with the standard stylized facts about the dynamic effects of monetary policy, according to which monetary shocks have a delayed and gradual effect on inflation. Furthermore, Mankiw and Reis (2002) propose a model to replace the NKPC by explaining the dynamic effects of aggregate demand on output and the price level. The baseline of the model is that some price setters set their prices based on old decisions and old information. They called this a sticky-information model to contrast it with the sticky-price model. Rudd and Whelan (2005) and Linde (2005) criticize some aspects of NKPC estimation and suggest that Galí and Gertler's (1999) results are the product of specification bias or suspect estimation methods. Nevertheless, GGLS (2005) re-assert that the NKPC is robust to a variety of estimation procedures, including GMM estimation of the closed form and non-linear instrumental variables.

[7] The NMS included in the analysis are the Czech Republic, Hungary, Poland, and Slovakia. The current euro area countries are Belgium, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal, and Spain.

[8] Their analysis is based on quarterly data for the period 1995:1 to 2005:4 for seven new members of the EU: the Czech Republic, Hungary, Latvia, Lithuania, Poland, Slovakia, and Slovenia.

or after 2004. They employ two alternative methods of estimation: the GMM and time-varying coefficient (TVC) estimation techniques. Using the GMM, they confirm the results typically found in the literature for the euro area, namely that lagged inflation has a positive and significant coefficient in the NKPC framework. This result based on the hybrid Phillips curve applies to both the euro area and the group of seven new EU countries. When applying TVC estimation, the results support the New Keynesian model of a purely forward-looking Phillips curve for both groups under consideration. Vašíček (2010) estimates an open economy Philips curve. The main finding is that inflation not only is driven by backward persistency but also comprises a forward-looking component. The forward-looking component dominates in some countries (Bulgaria, Cyprus, and Slovenia), the backward-looking component is predominant in others (Hungary, Lithuania, Poland, Romania, and Slovakia), and both components are of similar magnitude in others (the Czech Republic, Estonia, Latvia, and Malta). Mihailov et al. (2010a) estimate the NKPC based on the model by Galí and Monacelli (2005) for twelve NMS.[9] They extended the standard New Keynesian Phillips curve to an open economy setting. By employing GMM, the authors emphasize the importance of external (terms of trade) and internal drivers [10] of CPI inflation. The domestic and external inflation drivers are jointly significant in about half of the NMS, as opposed to just one country (UK) in their sample of OECD countries (Mihailov et al., 2010b). Their results indicate that the inflation process in four of the larger countries tends to be dominated by domestic variables, while in five of the smaller ones it is mostly affected by external variables.

3. Data and Methodology

The hybrid new Keynesian Phillips curve in this paper is given as:

$$\pi_{it} = \lambda mc_{it} + \gamma_f E_{it}\left\{\pi_{i,t+4}\right\} + \gamma_b \pi_{i,t-1} + \varepsilon_{it} \tag{5}$$

Galí and Gertler (1999) use the percentage change in the GDP deflator for π_t. Our empirical framework differs from that benchmark. We employ an alternative measure of inflation, π_t, expressed as the year-on-year (y-o-y) percentage change in the HICP for all the observed countries. Vašíček (2009) argues that although producer price inflation or value added inflation (proxied typically by the GDP deflator) is the appropriate measure of inflation for the Galí and Gertler model, economic agents and monetary authorities unquestionably perceive rather CPI inflation. Vašíček (2009) also emphasizes that CPI inflation is especially relevant in small open economies that import a substantial part of their consumption basket.[11]

[9] Poland, Hungary, the Czech Republic, Slovakia, Slovenia, Estonia, Latvia, Lithuania, Bulgaria, Romania, Cyprus, and Malta.

[10] For example, Stavrev (2006), applying a generalized dynamic factor model framework, decomposes inflation in the NMS into common and country-specific components. The analysis suggests that a significant part of headline inflation in the NMS is driven by common factors, such as price level convergence and EU integration. However, country-specific factors have also played a role in the inflation process and are related to the financial conditions, pass-through from foreign prices, and demand-supply situation in each country.

As a proxy for real marginal cost, we use a measure of the output gap (mc_t) obtained by extracting short-run fluctuations from the long-run GDP trend via the Hodrick-Prescott (HP) filter.[12]

The model used here comprises the expected inflation rate $E_{it}\left\{\pi_{i,t+4}\right\}$ derived from consumer surveys. The questionnaire in those surveys refers to consumers' perceptions and expectations about aggregate price level changes in the previous and following 12 months (4 quarters), so the indicators of expected inflation obtained from them are to be compared with y-o-y inflation rates (European Commission, 2007). To extract the expected inflation measure from consumer surveys in the countries of interest we employ the probability method as proposed by Forsells and Kenny (2004).[13]

The initial impulse for quantifying consumers' inflation expectations (obtained by surveys) via the probability method was given by Theil (1952). Since then, significant improvements and modifications have been made in applying the method by Carlson and Parkin (1975) as well as Berk (1999). Continuing to the Carlson and Parkin (CP) methodology, in this paper the approach used by Forsells and Kenny (2004) is applied. The full derivation of the probability method can be found in Nardo (2003) and Sabrowski (2009). Here we present only the basic assumptions and relations.

The main assumption behind the CP probability method is that respondents answer the questionnaire according to a probability density function related to their inflation perceptions/expectations. Moreover, the probability density function may vary over time and between respondents. Hence, the shares of respondents providing each particular answer to the survey question can be interpreted as maximum likelihood estimates (probabilities) of areas under the density function of inflation perceptions/expectations (Forsells and Kenny, 2004). The choice of the distribution of perceived/expected inflation in most research studies comes down to the normal distribution, as supported by the Central Limit Theorem. Assuming that consumers' inflation perceptions/expectations at time *t* for *N*surveyed consumers are *i.i.d.* random variables, then the distribution of the sum of these variables must be asymptotically normal (Dias et al., 2009).

The necessary condition to apply the probability method in quantifying consumers' inflation expectations is to have access to time series data on two consumer survey questions: one concerning past price changes (inflation perceptions) and one concerning future price changes (inflation expectations) (see the *Appendix* for a more detailed elaboration of the method).

All data is obtained from Eurostat and expressed in quarterly frequencies. Since the consumer survey data is not available until 2001 for some countries,

[11] Accordingly, Vašíček (2009) analyzes four NMS countries and uses the hybrid version of the NKPC as developed in Galí and Gertler and Galí, Gertler, and Lopez-Salido as a benchmark. He extends the empirical framework to include additional (potentially) inflation-driving variables and suggests that the inflation dynamics of small open economies can be affected by external variables unrelated to domestic firms' price setting.

[12] Constant λ (lambda) is set to 1600, as suggested for quarterly data (Hodrick and Prescott, 1997).

[13] In applying consumer surveys to quantify inflation perceptions and expectations, a number of different methods are available (balance statistic, nonlinear regression method or probability method). However, it is widely accepted that the probability method results in the best fit (Nardo, 2003).

the dataset has been aligned to the time span 2002Q2–2009Q2. All variables were seasonally adjusted using the TRAMO/SEATS method. A graphical presentation of the average values of π_t, mc_t, and $E_{it}\left\{\pi_{i,t+4}\right\}$ is given in *Figure 2 in the Appendix*. It is clear that all the observed variables exhibit a common trend. In the first three years all the variables display rather stable levels. As from 2005 they share a common upward trend, while the emergence of the global crisis at the start of 2008 fosters a sharp downturn in all the observed variables.[14]

To estimate the model given by equation (5), a dynamic panel model has to be used. Most dynamic panel data estimators have good properties for datasets with a large number of cross-section units and a small time span (T). For larger T as in this case, the GMM estimators proposed by Anderson and Hsiao (AH), Arellano and Bond (AB) (1991), and Blundell and Bond (BB) (1998) can produce inconsistent and biased estimates.

All these estimators impose the assumption of sample homogeneity. One of the central findings from datasets with a reasonably large T is that the assumption of slope parameter homogeneity is often inappropriate. Therefore, a separate model has to be estimated for each cross section. Additionally, dynamic panels with large T are usually nonstationary. This point has been made by Pesaran and Smith (1995) and Pesaran, Shin, and Smith (1999). There are several approaches to the estimation of nonstationary heterogeneous dynamic panel models. For example, the dynamic fixed effects (DFE) model pools the time series of all cross-sections and only the intercepts are allowed to differ across groups. On the other hand, the mean group method (MG) estimates separate regressions for each group and averages the coefficients over groups. Pesaran, Shin, and Smith (1999) have proposed a pooled mean group (PMG) estimator that combines pooling and averaging. This intermediate estimator allows the intercept, short-run coefficient, and error variances to differ across the groups but constrains the long-run coefficients to be equal across groups. Taking into account that the observed nine emerging economies are different with respect to monetary policy and exchange rate regimes, these three heterogeneous dynamic panel models are considered here.

The hybrid new Keynesian Phillips curve can be written as an autoregressive distributive lag ARDL (1,0,0)

$$\pi_{it} = \gamma_{bi}\pi_{i,t-1} + \gamma_{fi}E_{it}\left\{\pi_{i,t+4}\right\} + \lambda_i mc_{it} + \varepsilon_{it}, i = 1,2,...,N, t = 1,2,...,T \tag{6}$$

If the variables in (6) are I(1) and cointegrated, then the error term is an I(0) process for all i.

The error correction reparametrization of (6) is

$$\Delta\pi_{it} = \varphi_i\left(\pi_{i,t-1} - \varphi_{1i}E_{it}\left\{\pi_{i,t+4}\right\} - \varphi_{2i}mc_{it}\right), \tag{7}$$

where

$$\varphi_i = -\left(1-\gamma_{bi}\right), \varphi_{1i} = \frac{\gamma_{fi}}{1-\gamma_{bi}}, \varphi_{2i} = \frac{\lambda_i}{1-\gamma_{bi}} \tag{8}$$

[14] It is worthwhile to mention that many empirical papers dealing with euro area consumers' inflation expectations and perceptions demonstrate their upward bias after the introduction of the euro. For a very detailed and thorough discussion of the phenomenon, see European Commission (2009).

Parameter φ_i is the error-correcting speed of adjustment. If $\varphi_i = 0$, then there would be no evidence of a long-run relationship. This parameter is expected to be significantly negative under the prior assumption that the variables show a return to long-run equilibrium.

4. Results

Prior to any econometric modeling, all variables were tested for stationarity. Univariate unit root tests in general have low power. In order to improve the test power, five different panel unit root tests are analyzed. Unlike the other tests performed, Pesaran's covariate-augmented Dickey-Fuller (CADF) test allows for cross--section dependence. The results of the unit root tests (Levin, Lin & Chu, Im, Pesaran and Shin, ADF-Fisher and PP-Fisher, and Pesaran's CADF) are presented in *Table 1.*

Unit root tests give somewhat mixed results for π_{it} and mc_{it}. Pesaran's CADF test does not reject the nonstationarity hypothesis for any of the observed variables. Hence it cannot be concluded that all the series of interest are stationary. Accordingly, three alternative nonstationarity dynamic models for heterogeneous panels are estimated: PMG, MG, and DFE. Prior to that, panel cointegration tests can be employed to identify the long-run equilibrium process. The Westerlund (2007) and Persyn and Westerlund (2008) heterogeneous panel cointegration test is employed, allowing for cross-section dependence (see *Table 2*).

Therefore, there does exist a cointegration relationship between π_{it}, mc_{it}, and $E_{it}\{\pi_{i,t+4}\}$. Further on, the Hausman test of long-run homogeneity of coefficients is employed to determine which estimator is more appropriate. The MG estimator provides consistent estimates of the mean of the long-run coefficients, but these will be inefficient if slope homogeneity holds. If the long-run slope coefficients are homogeneous, the PMG and DFE estimators are consistent and efficient (Pesaran, Shin, and Smith, 1999) – see *Table 3*.

On the basis of Hausman test it is not possible to reject the hypothesis of poolability of long-run coefficients (p-value 0.9151 for the PMG estimator and 0.9115 for the DFE estimator). It can be concluded that the PMG and DFE estimators are efficient under the null hypothesis, and are preferred to the MG estimator. The PMG estimator is preferred to the DFE estimator because it allows for short-run coefficient heterogeneity. The long-run coefficients for mc_{it} and $E_{it}\{\pi_{i,t+4}\}$ are positive and statistically significant. The average value of the error correction coefficient is -0.15, implying that equilibrium is reached in about 6 quarters. All three independent variables ($\pi_{i,t-1}, mc_{it}$, and $E_{it}\{\pi_{i,t+4}\}$) are statistically significant (γ_b, γ_f, and λ from *Table 3*). The PMG estimator allows for short-run heterogeneity, so it is possible to estimate separate short-run coefficients for each country (see *Table 4*).

To investigate the robustness of the results, several additional estimation procedures (which imply slope homogeneity) were employed: Least Squares Dummy Corrected (LSDVc), Arellano and Bond one step (GMM1), Arellano and Bond two step (GMM2), Arellano and Bond two step where variable mc_{it} is treated as an endogenous variable (GMM2e1), and Arellano and Bond two step where variable $E_{it}\{\pi_{i,t+4}\}$ is

Table 1 Panel Unit Roots Tests

Variables		Method	Prob.*	Obs
π_{it}	const	Levin. Lin & Chu t*	0.2327	207
		Im. Pesaran and Shin W-stat	0.0014	207
		ADF-Fisher Chi-square	0.0019	207
		PP-Fisher Chi-square	0.7133	224
		Pesaran's CADF	0.347	234
	const and trend	Levin. Lin & Chu t*	1.0000	198
		Im. Pesaran and Shin W-stat	0.0004	198
		ADF-Fisher Chi-square	0.0002	198
		PP-Fisher Chi-square	0.9562	252
		Pesaran's CADF	0.591	234
	no const no trend	Levin. Lin & Chu t*	0.1319	201
		Im. Pesaran and Shin W-stat		
		ADF-Fisher Chi-square	0.3439	201
		PP-Fisher Chi-square	0.1803	224
		Pesaran's CADF		
$E_{it}\left\{\pi_{i,t+4}\right\}$	const	Levin. Lin & Chu t*	0.0000	249
		Im. Pesaran and Shin W-stat	0.0000	249
		ADF-Fisher Chi-square	0.0000	249
		PP-Fisher Chi-square	0.0000	252
		Pesaran's CADF	0.063	234
	const and trend	Levin. Lin & Chu t*	0.0000	249
		Im. Pesaran and Shin W-stat	0.0000	249
		ADF-Fisher Chi-square	0.0000	249
		PP-Fisher Chi-square	0.0000	252
		Pesaran's CADF	0.085	225
	no const no trend	Levin. Lin & Chu t*	0.0000	248
		Im. Pesaran and Shin W-stat		
		ADF-Fisher Chi-square	0.0000	248
		PP-Fisher Chi-square	0.0000	252
		Pesaran's CADF		
mc_{it}	const	Levin. Lin & Chu t*	0.0000	229
		Im. Pesaran and Shin W-stat	0.0000	229
		ADF-Fisher Chi-square	0.0000	229
		PP-Fisher Chi-square	0.9798	252
		Pesaran's CADF	0.670	234
	const and trend	Levin. Lin & Chu t*	1.0000	198
		Im. Pesaran amd Shin W-stat	0.0044	232
		ADF-Fisher Chi-square	0.0003	232
		PP-Fisher Chi-square	0.9562	252
		Pesaran's CADF	0.834	234
	no const no trend	Levin. Lin & Chu t*	0.8087	234
		Im. Pesaran and Shin W-stat		
		ADF-Fisher Chi-square	0.0000	234
		PP-Fisher Chi-square	0.0365	252
		PP-Fisher Chi-square	0.0054	252
		Pesaran's CADF		234

Notes: *Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality Im, Pesaran and Shin; ADF-Fisher and PP-Fisher test- Null Hypotesis: Unit Root (Individual Unit Root process) Levin, Lin & Chu Test- Null Hypothesis: Unit Rott (common

Unit Root process). Automatic lag lenght selection based on Modified Scwarz Criterion and Barlett kernel.

Source: authors´ calculations

Table 2 Panel Cointegration Tests

Statistic[a]	*P*-value	Robust *P*-value[b]
Gt	0.001	0.010
Ga	0.007	0.000
Pt	0.001	0.020
Pa	0.006	0.030

Notes:[a] H_0: series are not cointegrated

[b] Robust in presentence of cross-section dependence

Source: authors´ calculations

Table 3 PMG and DFE Estimation Results

Explanatory Variables	PMG	DFE
Long run mc_{it}	1.3971*** (0.2101)[a]	1.4599*** (0.0521)
Long run $E_{it}\{\pi_{i,t+4}\}$	0.2404*** (0.0078)	0.2202** (0.01428)
Error correction coefficient $\varphi_{.}$	-0.15375*** (0.0127)	-0.1395*** (0.0194)
γ_b	0.8462*** (0.0127)[b]	0.8604*** (0.0197)
λ	0.2148*** (0.0368)	0.2037*** (0.0171)
γ_f	0.0369*** (0.0124)	0.0307*** (0.0114)
Number of observations	252	252
Number of groups	9	9
Hausman test for poolability of countries (*p* value)	0.9151	0.9115

Notes: *, **, *** indicates significance at 10%, 5% and 1% significance level.

[a] Numbers in the brackets are standard errors for full PMG and DFE.

[b] The standard error for γ_b, λ and γ_f are approximated by using the Delta method.

Source: authors´ calculations

treated as an endogenous variable (GMM2e2). GMM1, GMM2, and GMM2e are useful for examining the potential endogeneity of the explanatory variables. All the GMM estimators can control for potential problems due to the presence of endogenous regressors, by using their lagged values in level or in difference as instruments. Endogeneity is tested by the Sargan test (see *Table 5*).

Despite all the mentioned restrictions of the other estimators, their results are not significantly different from the results of the PMG estimates. Furthermore, theoretically it is plausible that mc_{it} influences inflation and vice versa. With the purpose of evaluating the potential endogeneity of the output gap and expected inflation, instrumental variables are used (the second lag of the output gap and the second lag

Table 4 Full PMG Estimation Results

Country	φ_i	γ_{bi}	λ_i	γ_{fi}
Czech Republic	-0.2081*** (0.0536)[a]	0.7919*** (0.0536)[b]	0.2907*** (0.0669)	0.0500*** (0.0206)
Estonia	-0.1661*** (0.0314)	0.8339*** (0.0314)	0.2320*** (0.0323)	0.0399*** (0.0150)
Latvia	-0.1487*** (0.0277)	0.8513*** (0.0277)	0.2077*** (0.0296)	0.0357*** (0.0136)
Lithuania	-0.1207*** (0.0206)	0.8793*** (0.0206)	0.1686*** (0.0265)	0.0290*** (0.0090)
Hungary	-0.1168* (0.0638)	0.8832* (0.0638)	0.1631* (0.0877)	0.0280 (0.0175)
Poland	-0.2165*** (0.0572)	0.7835*** (0.0572)	0.3024*** (0.0753)	0.0520** (0.0223)
Romania	-0.1329*** (0.0330)	0.8671*** (0.0330)	0.1856*** (0.0476)	0.0319** (0.0142)
Slovenia	-0.1605*** (0.0375)	0.8395*** (0.0375)	0.2242*** (0.0440)	0.0385** (0.0154)
Slovak Republic	-0.1131** (0.0528)	0.8869** (0.0528)	0.1580** (0.0728)	0.0271* (0.0156)

Notes: *, **, *** indicates significance at 10%, 5% and 1% significance level.
[a] Numbers in the brackets are standard errors for full PMG.
[b] The standard error for γ_{bi} , λ_i and γ_{fi} are approximated by using the Delta method.

Source: authors´ calculations

Table 5 Robustness Check Results

Explanatory Variables	LSDVc[e]	GMM1	GMM2	GMM2e1[c]	GMM2e2[d]
$\pi_{i,t-1}$	0.8810*** (0.0217)[a]	0.8722*** (0.0521)[b]	0.9001*** (0.0664)	0.9323*** (0.0673)	0.8211*** (0.1504)
$E_{it}\left\{\pi_{i,t+4}\right\}$	0.0271** (0.0114)	0.02805** (0.01428)	0.0239* (0.0143)	0.0208*** (0.0150)	0.0380* (0.0225)
mc_{it}	0.2020*** (0.0195)	0.2148*** (0.0194)	0.2114*** (0.0205)	0.2082*** (0.0200)	0.2190*** (0.0356)
Number of observations	243	243	243	243	243
Number of groups	9	9	9	9	9
R^2	0.9422			-	-
Sargan test (*p*-value)		-	0.9993	1	1
m1 test (*p*-value)		0.0526	0.0362	0.0352	0.0719
m2 test (*p*-value)		0.3582	0.4039	0.3938	0.6576

Notes: *, **, *** indicates significance at 10%, 5% and 1% significance level.
[a] Numbers in the brackets are standard errors for LSDVc,GMM2,GMM2e1,GMM2e2.
[b] Numbers in the brackets are robust standard errors,GMM1.
[c] Variable mc_{it} is treated as endogenous variable.
[d] Variable $E_{it}\left\{\pi_{i,t+4}\right\}$ is treated as endogenous variable.
[e] Bias correction initialized by Arellano and Bond estimator.

Source: authors´ calculations

of expected inflation). In the GMM2e1 model mc_{it} is treated as an endogenous variable and in GMM2e2 expected inflation is treated as an endogenous variable. The results did not significantly change.

Overall, it can be concluded that the results obtained are robust across different estimation procedures and that there is no serious problem of endogeneity in the proposed model.

Taken as a whole, the results obtained from the research conducted are in line with the NKPC research performed to date and are supportive of the hybrid NKPC for the nine observed transition economies.

5. Conclusion

Recent papers conclude that the NKPC is a good approximation of inflation dynamics in both the US and Europe. Thus, this paper reviews the ability of the model to fit the data, the importance of forward-looking and backward-looking behavior in price-setting, and whether the output gap can be regarded as a good indicator of transition countries' inflationary pressures. Most empirical studies of the NKPC employ official, publicly available expected inflation measures or proxy variables to account for future inflation expectations. By contrast, taking into consideration recent advances in modeling expected inflation, the probability CP method was employed here to assess expected inflation on the basis of consumer survey responses.

Furthermore, to estimate the hybrid NKPC model, several approaches for the estimation of dynamic panel models are employed. All variables were tested for stationarity and five different panel unit root tests are analyzed. Since the unit root tests give somewhat mixed results, we considered three alternative nonstationarity dynamic models for heterogeneous panels: the dynamic fixed effects (DFE) model, the mean group (MG) method, and the pooled mean group (PMG) estimator. Prior to that, a Westerlund heterogeneous panel cointegration test is employed to identify the long--run equilibrium process, allowing for cross-section dependence. Therefore, there does exist a cointegration relationship between π_{it}, mc_{it}, and $E_{it}\left\{\pi_{i,t+4}\right\}$. Further on, the Hausman test of long-run homogeneity of coefficients is employed to determine which estimator is more appropriate. It can be concluded that the PMG estimator is preferred because it allows for short-run coefficient heterogeneity. The long-run coefficients for mc_{it} and $E_{it}\left\{\pi_{i,t+4}\right\}$ are positive and statistically significant. The average value of the error correction coefficient is -0.15, implying that equilibrium is reached in about 6 quarters. All three independent variables (π_{it}, mc_{it}, and $E_{it}\left\{\pi_{i,t+4}\right\}$) are statistically significant (γ_b, γ_f, and λ) and have the correct positive sign. The results of the error correction model suggest strong inertial behavior of inflation due to a relatively large and statistically significant coefficient of lagged inflation. Explicitly, lagged inflation is statistically significant and the estimated coefficient rises from 0.84 to 0.86 (see *Table 3*) in the PMG and DFE estimation results. The existence of backwardness implies that inflation dynamics are determined not merely by forward--looking behavior and real marginal costs, but also by past inflation rates. These results apply to all the observed countries and are in line with the euro area results presented by GGLS (2001), Dabušinskas and Kulikov (2007), Hondroyiannis et al.

(2008), and Vašíček (2010). The output gap is also statistically significant for all models and the coefficient is around 0.21 for both the PMG and DFE models (see *Table 3*). In spite of the consequential historical disruptions and other economic specificities (changes in political and economic systems) of transition economies, they are no exception when estimating the hybrid NKPC. Therefore, we can conclude that the hybrid NKPC with the output gap as an explanatory variable performs in accordance with the theory in all the observed countries.

To investigate the robustness of the results, several additional estimation procedures (which imply slope homogeneity) were employed and their results are not significantly different from the results of the PMG estimates. Overall, it can be concluded that the results obtained are robust across different estimation procedures and that there is no serious problem of endogeneity in the proposed model.

APPENDIX A

Consumer Survey Questions on Perceived and Expected Inflation

Q5 How do you think that consumer prices have developed over the last 12 months? They have:

1. risen a lot (++)
2. risen moderately (+)
3. risen slightly (=)
4. stayed about the same (-)
5. fallen (--)

N. don't know

Q6 By comparison with the past 12 months, how do you expect that consumer prices will develop in the next 12 months? They will:

1. increase more rapidly (++)
2. increase at the same rate (+)
3. increase at a slower rate (=)
4. stay about the same (-)
5. fall (--)

N. don't know

APPENDIX B

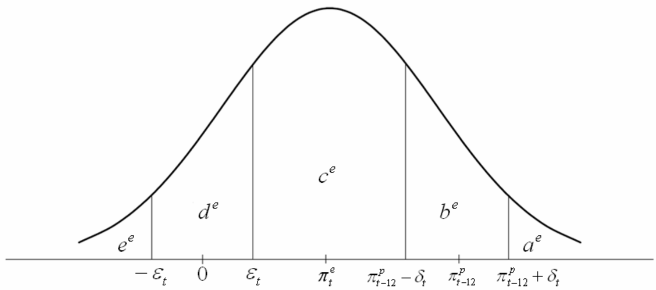
Expected Inflation Derivation

Once one has the data on consumers' responses, the aggregate expected inflation probability density function can be presented graphically as in *Figure 1*.

Let a^e, b^e, c^e, d^e, and e^e be the fractions of respondents indicating that prices will increase more rapidly, increase at the same rate, increase at a slower rate, stay about the same or fall in the following 12 months (respectively).

Hence it can be seen that the CP method rests on two sensitivity intervals. Firstly, it is assumed that the respondents declare in the questionnaire that prices

Figure 1 Inflation Expectations Probability Distribution



will change (rise/fall) in period t only if the expected inflation rate π_t^e (but surveyed 12 months before period t) is at least ε_t units from zero.

Secondly, it is assumed that the respondents condition their qualitative assessments on their perception of past inflation. Therefore, they indicate that prices will rise faster/slower only if the expected inflation rate is outside the interval

$$\left[\pi_{t-12}^p - \delta_t, \pi_{t-12}^p + \delta_t\right]$$

centered around the perceived inflation rate π_{t-12}^p.

Denoting the cumulative distribution function of the expected rate of inflation by F, *Figure 1* can be expressed mathematically through the following set of equations:

$$a^e = P\left(\pi_t^e > \pi_{t-12}^p + \delta_t\right) = 1 - F\left(\pi_{t-12}^p + \delta_t\right)$$

$$b^e = P\left(\pi_{t-12}^p - \delta_t < \pi_t^e < \pi_{t-12}^p + \delta_t\right) = F\left(\pi_{t-12}^p + \delta_t\right) - F\left(\pi_{t-12}^p - \delta_t\right)$$

$$c^e = P\left(\varepsilon_t < \pi_t^e < \pi_{t-12}^p - \delta_t\right) = F\left(\pi_{t-12}^p - \delta_t\right) - F\left(\varepsilon_t\right)$$

$$d^e = P\left(-\varepsilon_t < \pi_t^e < \varepsilon_t\right) = F\left(\varepsilon_t\right) - F\left(-\varepsilon_t\right)$$

$$e^e = P\left(\pi_t^e < -\varepsilon_t\right) = F\left(-\varepsilon_t\right)$$

Standardizing the normal probability density function and solving the above five equations enables derivation of the explicit expression for expected inflation (see Łyziak and Stanisławska, 2006, for a more detailed elaboration on quantifying methods).

$$\pi^e = \pi^p \cdot \frac{\left(C^e + D^e\right)}{C^e + D^e - \left(A^e + B^e\right)}$$

where $A^e = Nz^{-1}\left(1 - a^e\right)$, $B^e = Nz^{-1}\left(1 - a^e - b^e\right)$, $C^e = Nz^{-1}\left(1 - a^e - b^e - c^e\right)$, $D^e = Nz^{-1}(e^e)$, and Nz is the cumulative standardized normal distribution function.

Many empirical papers use the official inflation rate most recently available to the respondents as a proxy for perceived inflation π^p (for instance Łyziak, 2003).

Nevertheless, it seems unreasonable to assume *a priori* that consumers perceive price changes accurately due to the possible signal extraction problem (Lucas, 1976). Consequently, perceived inflation is also quantified from consumer survey results using the CP probability method.

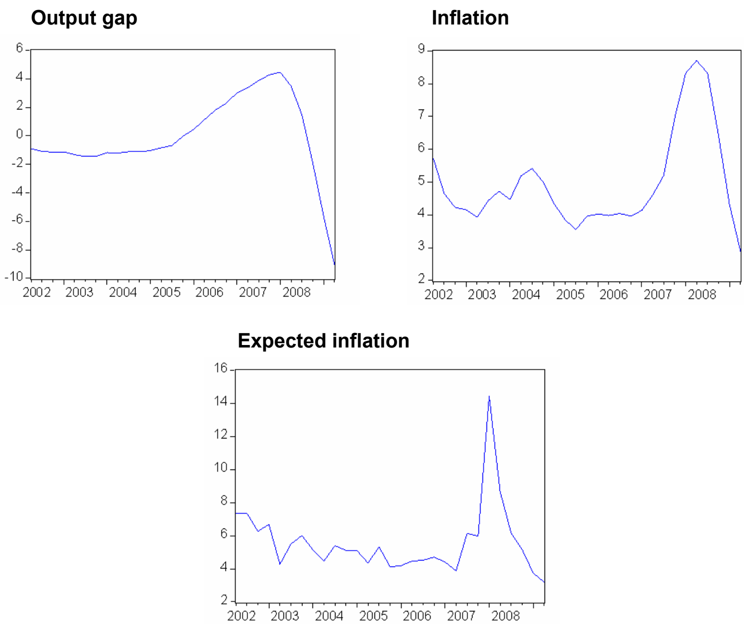
Let a^e, b^e, c^e, d^e, and e^e be the fractions of respondents indicating that prices have risen a lot, risen moderately, risen slightly or fallen in the last 12 months (respectively). Then, again using the probabilistic method, the expression for perceived inflation can be derived as follows:

$$\pi^p = -s \cdot \frac{A^p + B^p}{A^p - B^p}$$

where $A^p = Nz^{-1}\left(1 - a^p - b^p - c^p\right)$, $B^p = Nz^{-1}\left(e^p\right)$, and s is a scaling factor (derived under the assumption that the average value of the perceived inflation rate equals the average value of the actual inflation rate) (see Łyziak and Stanisławska, 2006, for a full derivation of the perceived inflation and scaling factor expressions). Thus, the expression for obtaining the scaling factor in period T is given as follows:

$$s_T = \frac{-\sum_{t=1}^{T} \pi_t}{\sum_{t=1}^{T} \left(\frac{A_t^p + B_t^p}{A_t^p - B_t^p} \right)}$$

Figure 2 Average Values of Observed Variables for Nine Countries of Interest



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