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Inflation convergence in Central and Eastern Europe with a view to adopting the euro

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Abstract

In this paper we consider inflation rate differentials between seven Central and Eastern Countries (CEECs) and the Eurozone. We focus explicitly upon a group of CEECs given that although they are already member states, they are currently not part of the Economic and Monetary Union (EMU) and must fulfil the Maastricht convergence criteria before being able to adopt the euro. However, this group of countries does not have an opt-out clause and so must eventually adopt the single currency. Hence, considering divergence in inflation rates between each country and the Eurozone is important in that evidence of persistent differences may increase the chance of asymmetric inflationary shocks. Furthermore, once a country joins the Eurozone the operation of a country specific monetary policy is no longer an option. We explicitly test for convergence in the inflation rate differentials, incorporating non-linearities in the autoregressive parameters, fractional integration with endogenous structural changes, and also consider club convergence analysis for the CEECs over the period 1997 to 2011 based on monthly data. Our empirical findings suggest that the majority of countries experience non-linearities in the inflation rate differential, however there is only evidence of a persistent difference in three out of the seven countries. Complementary to this analysis we apply the Phillips and Sul (2007) test for club convergence and find that there is evidence that most of the CEECs converge to a common steady state.

Keywords: Central and Eastern Europe , euro adoption, inflation convergence, non-linearities.

JEL classification: E31, E32, C22

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1. Introduction and background

Enlargement of the euro area is probably one of the most prominent topics within the European Union (EU) agenda. This is not surprising after the problems that Greece has experienced (and is still undergoing) and the questioning of their preparation in 1999, for adopting the euro back in 2001. Arguably, if Greece still had flexibility in their exchange rate policy then perhaps their current debt crisis could have been avoided. During 2011 the Eurozone has experienced unprecedented turmoil, which raises further questions about whether the euro area has ever been an optimum currency area. The focus of this paper is to address whether or not it is a good idea to encourage more member states to adopt the common currency, in particular the Central and Eastern European Countries (CEECs). Most of the CEECs which are already member states still have to fulfill the Maastricht convergence criteria so as to be able to join the euro area.[1] However, fulfilling the Maastricht nominal convergence criteria may not be enough to grant a future free of turbulence within euroland.

Given the commitment from the European Central Bank (ECB) for price stability and the current target to tackle inflation,[2] losing monetary policy may be especially problematic if the countries face so-called asymmetric shocks. That is, shocks affecting different countries in a different manner, and hence, causing a problem of synchronisation of income, inflation and unemployment rates, which potentially will require different policy responses. This is particularly important in a monetary union, since it implies losing the possibility of intervention in the exchange rate market to depreciate the currency, or the option of financing deficits by monetary expansions.

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[1] Details on the criteria can be found in www.ecb.int/ecb/history/enlargement/html/faqenlarge.en.html#l4.

[2] The primary objective of the ECB's monetary policy is to maintain price stability. The ECB aims at inflation rates of below, but close to, 2% over the medium term.

Hence, in this paper we investigate whether a common monetary policy decided by the ECB would be appropriate to the new EU countries from Central and Eastern Europe. The process of transition from communism towards that of a free market has been intense during the last 20 years, following a series of structural and political reforms. However, whether this process has facilitated conditions favourable to economic convergence is open to debate. Currently, it is unknown whether their inflation rates have converged to the same cycle and level as that of the Eurozone. Hence, applying the same monetary policy to an area where different countries have different inflation rates, may be detrimental for some economies. Whereas the inflation rates of Germany, Luxemburg and the Netherlands may be relatively low and stable, Central and Eastern European countries still face high and volatile inflation rates, due to expasionary fiscal policies, aiming at boosting the process of convergence in GDP per capita. Consequently, it is questionable whether a central institution with the monopoly of supplying money (i.e. the European Central Bank ECB) and which decides a uniform monetary policy is beneficial for all countries.

Although the optimum currency areas theory establishes the necessary conditions for the success of a monetary union (see Mundell, 1961), in this paper we focus on the possibility of asymmetric shocks and their effects upon inflation. Mundell (1961) showed the importance of facing symmetric macro shocks in a currency union composed of different countries or regions. It is also clear that the degree of factor productivity within the EU is quite unequal (see Table 1), and, hence, analysing the (a)symmetry of shocks to inflation becomes of paramount importance when deciding about the appropriateness of adopting the European single currency. GDP per capita, which is a proxy for labour productivity, of all of the CEECs is well below the EU-27 average with Bulgaria and Romania being the worst cases. Hence, it is debateable what would happen if those countries with lower income per capita embraced expansionary policies to improve the productivity of their production factors. Lehmann and Muravyev (2009) examine CEECs labour markets in comparison to that of the EU and reveal differences in terms of labour market policies and economic performance, which may hinder a uniform response to economic shocks. In addition to this, Figure 2 shows the budget balance as a percentage of the GDP for our target countries. Some countries such as the Czech Republic, Hungary and Poland have been running constant deficits during the sample period, and some others have been more affected by the financial crisis, 2007 onwards. In this situation, an autonomous monetary policy may be beneficial in order to correct the effects of these asymmetric shocks on the public deficit.

Inflation expectations are a key macroeconomic variable when deciding the appropriate monetary policy to adopt.[3] This is the base of the Lucas critique; central banks need to enhance credibility. Persistent differences in inflation rates within the monetary union may affect real interest rates, thus, creating important disparities in inflation expectations within the Union, thereby increasing the likelihood of asymmetric inflationary shocks (Busetti et al., 2007). Also, price differentials within an integrated monetary area can be seen as differences in external competitiveness, and hence is a useful way to test for asymmetric shocks, since exchange rate policy is no longer available to depreciate the currency and encourage exports. Nevertheless, there are other ways to test for the possibility of asymmetric shocks than using inflation differentials, such as business cycle synchronisation (see, Cuestas et al. 2011, amongst many others). However, given that the Maastricht criteria clearly establishes the importance of the inflation convergence criterion, and that the ECB medium term inflation target is clearly

[3] Taylor and McNabb (2007) showed the importance of individuals' expectations and business confidence in predicting the economic cycle. More recently, Gelper and Croux (2010) considered the role of a European Economic Sentiment Indicator (ESI) in forecasting economic activity and find that the ESI is a useful barometer of the economy. Hence, proper management of expectations becomes of paramount importance in economic policy.

defined,[4] we believe that inflation convergence is arguably the most appropriate and compelling means of assessing preparation to adopt the single currency.

The sample of countries considered in this paper consists of CEECs which are member states but not part of the Economic Monetary Union (EMU), i.e. Bulgaria, the Czech Republic, Hungary, Latvia, Lithuania, Poland and Romania. Most of these countries joined in 2004, with the exception of Bulgaria and Romania which joined in 2007, and none of them joined with an opt-out clause. This means that eventually they all need to fulfil the Maastricht criteria and adopt the single European currency. One of the Maastricht criteria is that the candidate country should be a member of the Exchange Rate Mechanism II (ERM II) for at least two years before adopting the euro. Only two countries, to date, the Baltic States Latvia and Lithuania, are in the ERM II.

The econometric techniques applied in this paper are related to the analysis of inflation convergence between these countries and the Eurozone. Assessing this hypothesis will allow us to provide valuable insights into the appropriateness of a centralised monetary policy, with no possibility of devaluations. More specifically, we test for the existence of unit roots in the inflation differentials for each country with respect to the Eurozone, accounting for the possibility of non-linearities in the data generation processes (DGPs), which may affect the speed of convergence, and also take into account the possibility of fractional integration incorporating breaks. Finally, we employ the recently developed club convergence tests (Phillips and Sul, 2007), to explore the robustness and gain additional insights from the analysis.

The remainder of the paper is organised as follows. In the next section we summarise the convergence hypothesis definition used in this paper and the most recent

[4] The Treaty on the Functioning of the European Union, Article 127 (1) establishes that "Without prejudice to the objective of price stability", the Euro-system shall also "support the general economic policies in the Union with a view to contributing to the achievement of the objectives of the Union". These include inter alia "full employment" and "balanced economic growth", however, the ECB has focussed its monetary policy decisions on inflation control since its creation.

contributions on the empirical testing of this hypothesis within the EU. Section 3 describes the econometric techniques applied in the paper, whilst in sections 4 and 5 we summarise the results and provide concluding remarks, respectively.

2. The convergence hypothesis

There are several definitions of economic convergence within the literature the most popular of which are the sigma-convergence (SC) by Barro and Sala-i-Martin (1991) and the long run convergence by Bernard and Durlauf (1995).

Barro and Sala-i-Martin (1991) base their SC definition on the assumption that over time the differentials of income per head between two countries should decrease. Basically, and applied to inflation differentials, SC will imply that:

$$y_t = \sigma y_{t-1} + \varepsilon_t \ , \ \text{with } \sigma < 1 \tag{1}$$

where $y_t = \pi_{i,t} - \pi_{euro,t}$, and with $\pi_{i,t}$ and $\pi_{euro,t}$ as the inflation rates of country *i* and the Eurozone respectively.

In a similar fashion, the Bernard and Durlauf (1995) definition of convergence implies that a set of income per capita converge if the long-term forecasts of the these variables are equal at a fixed time conditional on a set of available information Ω. Applied to the case of inflation convergence we have,

$$\lim_{k\to\infty} E(\pi_{i,t+k} - \pi_{euro,t+k} \,|\Omega_t) = 0 \tag{2}$$

The popularity of these definitions of convergence is related to their ease of empirical testing. Both definitions can be empirically analysed by means of tests for the order of integration of y_t and by performing a cointegration test on the vector $\left(\pi_{i,t} - \pi_{euro,t}\right)$. Hence, the hypothesis of convergence will be accepted if the variables are stationary and revert to zero.

There are a number of authors who have tested whether there is evidence in favour of the convergence hypothesis within Europe, and the EMU in particular. An important contribution, close to the date of the creation of the euro, is Kočenda and Papell (1997) who generally find results which are supportive of the convergence hypothesis for a number of EU member states and other industrialised countries, by means of unit root testing. Similar results are found by Camarero et al. (2000) who analysed the hypothesis of convergence for a number of peripheral EU countries, i.e. Italy, Spain and the UK, against Germany and the European average. Nevertheless, more recent contributions cast doubt on the convergence hypothesis in Europe. For instance, Holmes (2002) and Weber and Beck (2005) found that at the end of the period analysed, 1972-1999 and 1991-2004, respectively, the dispersion in inflation rates had not decreased. An interesting recent contribution is Lopez and Papell (2011) who find evidence of different levels of persistence in inflation differentials within the EMU. In particular, they find that there is an increase in convergence of inflation rates within the EMU after the creation of the euro, and some mild dispersion in the inflation rates of peripheral countries towards the end of the period considered (1999m1-2006m12).

Turning to studies focussing on CEECs, to the best of our knowledge, only a few contributions have analysed the inflation convergence hypothesis. Kočenda (2001) analysed macroeconomic convergence in this area focussing on several key variables, i.e. real industrial output, money aggregate (M1), producer and consumer prices, and nominal and real interest rate spreads. However, the results for inflation rates are mixed, and depend on the groups of countries analysed. A recent contribution by Spiru (2008), analyses the convergence hypothesis for this group of countries. Applying unit root tests for panel data based upon linear DGPs, she finds supportive evidence towards the convergence hypothesis against the Eurozone for Cyprus, Estonia and Slovenia (which have already adopted the euro), and Latvia and Poland. She finds evidence of non-linearities by means of applying linearity tests which are based upon the assumption of stationary residuals. Hence, Spiru's (2008) paper is an important starting point for understanding inflation convergence with the EMU.[5] Finally, Staehr (2010) finds evidence supporting the hypothesis of price convergence within the ten new EU countries from central and eastern Europe.

In this paper we take into account the possibility of non-linearities in the DGPs, by, firstly, testing for the presence of non-linearities in inflation differentials by means of linearity tests when the order of integration is unknown, and, secondly, by incorporating those non-linearities into the analysis by means of unit root tests for non-linear models, fractional integration, structural breaks and club convergence analysis.

3. Econometric methodology

In order to empirically test for convergence between pairs of variables, it is common to apply tests for the order of integration of the differential between the variables. In this paper we apply a group of tests which we consider are appropriate given the expected DGPs of our target variable. Initially, we conduct tests for non-linearities followed by the appropriate unit root test over the inflation differential between each country and the Eurozone (details on the data are provided in Section 4), depending upon whether there is underlying non-linearity in the DGP; then, fractional integration and structural breaks are considered, and finally we examine the issue of club convergence.

3.1 Non-linearities

In the literature on applied macroeconomics and mean reversion, there is an important debate on the power of the tests when the DGP is not properly specified in the auxiliary regressions. For instance, the existence of non-acknowledged non-linearities in the DGP has been reported as a source of power problems in traditional unit root tests (e.g. Kapetanios et al., 2003). Hence, this situation may increase the likelihood of committing

[5] Cuestas and Harrison (2010) also test for inflation persistence in the CEECs. However, the authors do not provide a comparison with the EU or Eurozone.

Type II Errors, which implies a bias towards not rejecting the null hypothesis when it is false. The existence of non-linearities can be justified economically for our inflation differentials. The speed of mean reversion or convergence may depend on the size of the initial deviation. For greater deviations, the monetary authorities may apply measures in order to control the inflation rate. However, for small shocks, which have only mild effects on the inflation rate, the monetary authorities may decide that it is not worth applying any contractionary monetary policy. Such instances would potentially yield non-linearities.

The most obvious approach to analyse this point is to test whether the process follows a linear or a non-linear process. However, traditional linearity tests such as the Granger and Teräsvirta (1993), Teräsvirta (1994) and Luukkonen et al. (1988) tests, are based upon the assumption that the variables are I(0), i.e. stationary. This is especially problematic in our framework, since the order of the integration is unknown. Thus, in a recent contribution, Harvey et al. (2008) propose a linearity test which can be applied either to I(0) or I(1) processes. These authors propose a Wald test when the order of integration is unknown, which is a weighted average of the Wald tests for the null of linearity when the variable is known to have a unit root and when it is known to be stationary I(0). Let's suppose that y_t is a stationary I(0) process. To test for the null of linearity we need to specify the following auxiliary regression:

$$y_t = \beta_0 + \beta_1 y_{t-1} + \beta_2 y_{t-1}^2 + \beta_3 y_{t-1}^3 + \varepsilon_t \ . \qquad (3)$$

Under the null hypothesis of linearity we have $H_0\colon \beta_2 = \beta_3 = 0$, and the alternative of nonlinearity, $H_1\colon \beta_2 \neq 0 \ and/or \ \beta_3 \neq 0$. The Wald test for testing these hypotheses is given by:

$$W_0 = {(RSS_R - RSS_U)}\Big/{RSS_U} \ , \qquad (4)$$

where RSS_R and RSS_U denote the residual sum of squares of the restricted, imposing H_0, and the unrestricted regression for equation (3), respectively. The W_0 test follows the standard $\chi^2(2)$ distribution, see Harvey et al. (2008). However, if the variable y_t is nonstationary I(1), the auxiliary regression for the test becomes:

$$\Delta y_t = \alpha_1 \Delta y_{t-1} + \alpha_2 (\Delta y_{t-1})^2 + \alpha_3 (\Delta y_{t-1})^3 + \varepsilon_t \ . \tag{5}$$

Under the null hypothesis of linearity in (5) we have $H_0 \colon \alpha_2 = \alpha_3 = 0$, against the alternative of a nonlinear process of $H_1 \colon \alpha_2 \neq 0 \ and/or \ \alpha_3 \neq 0$. Similarly to W_0, the Wald test for testing these hypotheses is given by:

$$W_1 = {(RSS_R - RSS_U)}\Big/{RSS_U} \ , \tag{6}$$

where RSS_R and RSS_U denote the residual sum of squares of the restricted, imposing H_0, and the unrestricted regression for equation (5), respectively. The W_1 test also follows the standard $\chi^2(2)$ distribution, see Harvey et al. (2008). Hence, the weighted averaged Wald test when the order of integration is unknown can be written as:

$$W_\lambda = (1 - \lambda) W_0 + \lambda W_1 \underset{d}{\rightarrow} \chi^2(2) \ , \tag{7}$$

where λ converges in probability to 1 when the variable is I(1) and to 0 when the process is stationary. According to Harvey et al. (2008), λ should be chosen as a combination of unit root and stationarity tests statistics.[6]

3.2 Unit root tests

Depending on whether it is possible to reject the null of linearity, we apply linear unit root tests, i.e. ADF tests, or non-linear unit root tests, in this case, following Sollis (2009). Sollis proposes a unit root test which takes into account the possibility of an autoregressive parameter, and hence the speed of mean reversion, dependent on the size of the deviations. This test is based upon the approach of Kapetanios et al. (2003), who

[6] See Harvey et al. (2008) for more details about λ.

propose a unit root test against the alternative of a globally stationary exponential smooth transition autoregression (ESTAR) model. The innovation of Sollis' (2009) test is related to the fact that ESTAR functions only allow controlling for absolute deviations of the shocks from equilibrium, regardless of the sign of the shock, i.e. symmetry. However, Sollis (2009) incorporates in his test the possibility of analysing the existence of asymmetric effects, which means that negative shocks may have different effects, in absolute magnitude, than positive shocks. This is particularly relevant for the purpose of our analysis. It is well known that an increase in the inflation rate is more difficult to tackle than a reduction below the target. Hence, we expect that the speed of mean reversion would differ depending on the sign, not only the size, of the shock.

Sollis' (2009) test is based upon the following asymmetric ESTAR (AESTAR) model:

$$\Delta y_t = G_t(\gamma_1, y_{t-1})\{S_t(\gamma_2, y_{t-1})\rho_1 + \left(1 - S_t(\gamma_2, y_{t-1})\right)\rho_2\}y_{t-1} + \varepsilon_t, \tag{8}$$

where

$$G_t(\gamma_1, y_{t-1}) = 1 - exp\left(-\gamma_1(y_{t-1}^2)\right), \text{ with } \gamma_1 \geq 0 \tag{9}$$

and

$$S_t(\gamma_2, y_{t-1}) = \left\{1 + exp\left(-\gamma_2\left(y_{t-1}\right)\right)\right\}^{-1}, \text{ with } \gamma_2 \geq 0. \tag{10}$$

Hence, the null hypothesis of a unit root can be specified as $H_0\colon \gamma_1 = 0$. However, under the null hypothesis, γ_1, ρ_1 and ρ_2, cannot be identified. Sollis (2009), by means of Taylor approximations, proposes testing for unit roots in this nonlinear framework using the following auxiliary equation:

$$\Delta y_t = \beta_1 y_{t-1}^3 + \beta_2 y_{t-1}^4 + error. \tag{11}$$

Thus, testing for unit roots in model (11) implies testing $H_0\colon \beta_1 = \beta_2 = 0$. Note that equation (11) may also incorporate lags of the dependent variable to control for autocorrelated residuals. Another innovation of Sollis' (2009) approach is that, once the null hypothesis of a unit root has been rejected, the null hypothesis of symmetric ESTAR versus the alternative of AESTAR can be tested. That is, it allows us to test for different effects, in absolute value, of positive and negative shocks on the variable. In this case, testing for the null hypothesis of symmetric ESTAR implies testing $H_0\colon \beta_2 = 0$, by means of standard hypotheses tests.

In order to explore the robustness of the analysis, we also analyse the convergence hypothesis by means of fractional integration tests. It is important to bear in mind that long memory processes, which need long periods of time to revert to equilibrium after a shock, may be wrongly classified as I(1) processes by conventional unit root tests. This is because the aforementioned unit root tests classify the variables as I(d), where d is only allowed to be 0 or 1. Fractional integration tests break the dichotomy of d, since this parameter is allowed to take any real value. Thus, it may be 0, 1, but also any real value between 0 and 1 or even above 1. Hence, if d is between 0 and 0.5, the variable is stationary and mean reverting, whereas if d belongs to the interval [0.5, 1) the variable is non-stationary, but still mean reverting. If $d \geq 1$, the variable is then non-stationary and non-mean-reverting. This has important implications for our analysis, since the degree of persistence is then determined by the estimation of d. Fractionally integrated or I(d) models can be specified as:

$$(1-L)^d y_t = u_t, \quad t = 1, \dots, T \tag{12}$$

where u_t is a covariance stationary I(0) process, whose spectral density function is positive and finite at the zero frequency. In this paper we apply several methods based on parametric, semiparametric and non-parametric techniques. Thus, we first employ Whittle estimates of d based on the frequency domain (Dahlhaus, 1989) along with a Lagrange Multiplier (LM) testing procedure developed by Robinson (1994). This latter method is very general in the sense that it allows us to test any real value of d, including the stationary (d < 0.5) and nonstationary hypotheses (d ≥ 0.5) with no need of prior differentiation of the series. Several semiparametric methods (Robinson, 1995a,b; Abadir et al., 2007; etc.) will also be conducted in the paper. In case of Robinson (1995b) the method is multivariate and thus, it permits us to test the null that all the *d* parameters are the same, which will give us some insights about the degree of homogeneity of persistence of shocks on the variables. Finally, the possibility of structural breaks in the context of I(d) models is also considered. This last point is particularly important in our framework. As aforementioned, this group of countries have undergone a number of deep structural reforms during the transition process from communism to market economies, as well as for preparation for EU membership. In addition, some events such as the fact of joining the EU, the creation of the euro, or the 2008-2011 financial crisis, may have also affected the speed of convergence (or divergence) in their inflation rates with respect to the Eurozone.

3.3 Club convergence

Finally, in order to test whether our target countries converge to a common inflation rate we apply the Phillips and Sul (2007) club convergence procedure. These authors develop a technique to test the hypothesis of convergence amongst countries, which allows us to group the countries ($i = 1,2,\dots,N$) into convergence clubs or clusters. According to Phillips and Sul (2007), any panel of individuals, countries, or regions, can be decomposed into a common term, μ_t, and an idiosyncratic component, δ_{it}:

$$Y_t = \{y_{1t}, y_{2t}, \dots, y_{Nt}\}' = \mu_t \delta_{it} \qquad \forall\, i, t. \tag{13}$$

To measure the distance of each country of the panel from the common component, Phillips and Sul (2007) propose the squared average transition differential H_1/H_t where:

$$H_t = \frac{1}{N}\sum_{i=1}^{N}\left(\hat{h}_{it} - 1\right)^2 \tag{14}$$

and

$$h_{it} = \delta_{it} \Big/ \tfrac{1}{N}\sum_{i=1}^{N}\delta_{it} \qquad (15)$$

is a measure of δ_{it} relative to the panel average, and therefore, the transition of country *i* relative to the panel mean. To identify the idiosyncratic component δ_{it}, the authors propose the following semiparametric model,

$$\delta_{it} = \delta_t + \left\{ \sigma_i \xi_{it} \big/ L(t) t^{\alpha} \right\}, \qquad (16)$$

where $\xi_{it} \sim iid(0,1)$ for all i, L(t) is a time dependent variable and α is the speed of adjustment. Accordingly, δ_{it} converges to δ_t for any positive value of α. The null hypothesis $H_0: \delta_i = \delta \; and \; \alpha \geq 0$ is tested against the alternative hypothesis $H_1: \delta_i \neq \delta, \forall \; \alpha \prec 0$. Testing for the null is based upon the following auxiliary regression:

$$\log(H_1/H_t) - 2\log L(t) = \hat{c} + \hat{b}\log(t) + u_t \qquad (17)$$

where $\log L(t) = \log(t+1)$. The fitted value of $\log(t)$ is $\hat{b} = 2\hat{\alpha}$ where $\hat{\alpha}$ is the estimated value of α under the null hypothesis. Rejection of the null for the whole panel does not imply there is not convergence, since it is possible to test, by means of an algorithm, whether there are clubs/clusters of convergence. That is, the procedure identifies different convergence clubs if it is not possible to identify convergence to a common component for all the countries analysed.

4. Empirical Evidence

4.1 The data

The inflation differentials are computed as the difference between the inter-annual inflation rate of the country and the inter-annual inflation rate of the Eurozone. The data has been downloaded from *Eurostat* and are based on harmonised Consumer Price Indices (CPIs). For all countries we have used monthly observations from 1997:1 to 2011:7, except Bulgaria, whose sample starts in 1997:12.

The plots of the inflation differential versus that of the Eurozone are displayed in Figure 2. In general, it is possible to observe a clear convergence pattern in the inflation rate differentials. Most countries suffered from periods of high inflation at the beginning of the sample, with Bulgaria and Romania being the worst cases. We also see a significant drop in the inflation differential for most countries at the beginning of 2003, due probably to increased inflation in the euro area caused by the "rounding up" of prices in euros, during the euro cash changeover (European Central Bank, 2003). Also, we observe a significant drop in the inflation rates of the Baltic States, where the 2010 financial downturn was more damaging for their aggregate demand than the other countries. This was preceded by a sudden rise in the inflation rates of the latter countries during 2008 and the beginning of 2009, which was mainly caused by food prices and housing expenses. In general it would appear that there is evidence of co-movement in the inflation rate differential with respect to that of the euro area, which may be an indication of a lack of asymmetric shocks affecting the inflation rates of these countries.

4.2 Results

The results of the Harvey et al. (2008), Sollis (2009) and the ADF tests are presented in Table 2. All the tests have been applied to the raw data, without any deterministic component in the auxiliary regressions. The reason for this is that allowing for a constant will imply that, if the null is rejected, the inflation series will show a constant gap with respect to the inflation rate of the Eurozone. In such a case, concluding that there is evidence of convergence will not imply that the same monetary policy should be applied to both.

First, we start by testing the hypothesis of linearity of the inflation differentials for each country. According to the second column of Table 2, for only two countries, i.e. Hungary and Lithuania, the null of linearity cannot be rejected. For the rest of the countries, the Harvey et al. (2008) test w_λ indicates evidence in favour of non-linear models. Hence, for Bulgaria, the Czech Republic, Latvia, Poland and Romania we apply the Sollis (2009) unit root test for non-linear AESTAR models, whereas for Hungary and Lithuania we apply the ADF test.

According to the results reported in the last three columns of Table 2, the null of a unit root cannot be rejected for three of our target countries; Bulgaria, Latvia and Poland. For the rest of the countries, the results indicate that the inflation differentials are non-linear and globally stationary and, in particular for the Czech Republic and Romania, that shocks have asymmetric effects on the variable. The latter finding means that shocks with a different sign but of equal magnitude will have different effects, in absolute terms, on the target variable.

Next we examine the possibility of fractional integration. As earlier mentioned unit root methods have the inconvenience that they have extremely low power if the true underlying process is I(d) with d different from 0 or 1. Table 3 displays the estimates of d for each individual series. The first two columns refer to the Whittle estimates of d, displaying also the 95% confidence band of the non-rejection values of d using Robinson's (1994) parametric approach, first assuming that the error term u_t is white noise and then allowing for autocorrelation by adopting the nonparametric method of Bloomfield (1973).[7] The last two columns refer to the semiparametric Whittle method of Robinson (1995b) generalized later by Abadir et al. (2007). We present the results here for three bandwidth numbers, m = 5, 13 ($\cong T^{0.5}$) and 20.

The first thing we observe in Table 3 is that there is very little evidence of mean reversion in the series examined. Thus, we only obtain an estimate of d significantly below 1 in the case of Bulgaria for the model with white noise u_t. For the remaining

[7] This method produces autocorrelations decaying exponentially as in the AR(MA) cases.

cases, we cannot reject the null of I(1) behaviour or, if it is rejected, it is in favour of higher degrees of integration.

In addition, we present, in Table 4, the results of the Robinson (1995a) log-periodogram test for fractional integration. In Panel (a) we report the results of the test for each country's inflation rate and that of the Eurozone. The reason for applying the test to each individual country's inflation rate is to analyse how (dis)similar the order of integration is across countries. Although the unit root tests reported some cases whereby the unit root was rejected, it was not possible to infer anything about how fast or slow the series would revert to equilibrium after a shock. In the second column of Panel (a) we report the estimated order of integration. Interestingly the euro area's inflation rate is quite close to the unit root, whereas Bulgaria's inflation rate seems to be stationary. In order to test whether shocks have similar effects on the inflation rates, we test for the equality of the *d* parameters. According to this F-test (which is reported in the note to Panel (a)), not surprisingly, the hypothesis of equal orders of integration is rejected. In Panel (b) we apply the F-test to pairs consisting of each country and the euro area, to highlight those countries' inflation rates with the same order of integration than the Eurozone's inflation rate. The hypothesis of equality of *d* cannot be rejected for the Czech Republic, Latvia, Lithuania and Poland, implying that Bulgaria, Hungary and Romania's inflation rates *d* are not similar to the *d* of the euro area's rate of inflation.

These results do not pose any contradiction with respect to our findings relating to the unit root tests. The unit root tests provide analysis of whether the inflation differentials tend to converge to zero after a shock, whilst with the fractional integration approach, we test whether the inflation rates react in a similar way after a shock. Hence these results have important policy implications. Although Hungary and Romania's inflation differentials appear to be stationary according to the unit root analysis, the results in Table 4 indicate that inflationary shocks experienced by these two countries tend to disappear faster than in the euro area. The cases of Latvia and Poland also deserve some comment. There was no evidence against the null of a unit root in the results reported in Table 2. However, the results presented in Table 4 Panel (b) indicate similar order of integration in their inflation rates than that in the euro area. Hence, although there is no evidence in favour of the convergence hypothesis, shocks tend to have similar effects on the inflation rates in all three, i.e. Latvia, Poland and the Eurozone.

Still in the context of fractional integration, the possibility of breaks in the data is also examined. This is a relevant issue since it has been argued by many authors that fractional integration might be an artificial artifact generated by the presence of breaks in the data (see, e.g., Cheung, 1993; Diebold and Inoue, 2001; Giraitis et al., 2001; Mikosch and Starica, 2004; Granger and Hyung , 2004; etc.). Table 5 displays for each series the number of breaks, along with the estimates of the break dates and the fractional differencing parameters for each subsample using the procedure developed by Gil-Alana (2008). This method is based on minimising the residuals sum of squares for different subsamples assuming that the break dates are endogenously determined by the model.[8] The results suggest that there are no breaks in the cases of Latvia and Romania; a single break in case of Bulgaria, Hungary and Poland, and two breaks are detected for Lithuania and the Czech Republic. Once more the results indicate little evidence of mean reversion, and although some estimates are found to be below unity the unit root null cannot be rejected. Interestingly, the breaks are quite close in time across countries, i.e. around the date of the creation of the euro and close to the end of the period, probably caused by the financial crisis. Also, it is worth mentioning that none of the breaks seem to be related to joining the EU or the ERM II (for the case of Lithuania). In all cases, it appears that the

[8] It uses a grid of values for the fractional differencing parameters and for the break dates.

creation of the euro generated a higher degree of dispersion between the Eurozone and our target countries. Furthermore, the years of the financial crisis have decreased slightly the speed of mean reversion. The latter phenomenon can be justified by the fast drop in the inflation differential with respect to the Eurozone, just after the initial shock in 2007.

Finally, we test, by means of the Phillips and Sul (2007) club convergence test, if the series of inflation of the CEECs candidates to adopt the euro, tend to converge to a common steady state. The null hypothesis is hence convergence to a common steady state. This is done by comparing the t-statistic of the $\log(t)$ coefficient in the auxiliary regression (17) with the critical value -1.65, for different groups of countries. In our case the t-statistic is -1.62, which is greater than the critical value, when Bulgaria is excluded; hence we cannot reject the null hypothesis that this group of countries, with the exception of Bulgaria, form a convergence club.

The results obtained highlight important policy implications for the future of the Eurozone, and for these countries. Bulgaria is a clear candidate to wait longer before adopting the euro, this is perhaps not surprising given it only became a member state in 2007. The results point against the convergence hypothesis and the order of integration of Bulgaria's inflation rate is much lower than the euro area. This does not mean that Bulgaria is not doing things right, but for their own self interest they should have the possibility to accommodate differently to inflationary shocks than the euro area. The Czech Republic is probably one of the most clear cases of similarity of inflationary shocks with the euro area, along with Lithuania, which basically implies that losing their monetary policy will not, in principle, pose major problems in case of asymmetric macro shocks. Hungary and Romania are interesting case studies. Both countries inflation rates have seemed to converge to the inflation rate of the Eurozone, however, according to the results of Table 4, there is still some danger of hazardous effects of asymmetric shocks.

Latvia and Poland also seem to be similar to the euro area in the way they react to inflationary shocks, although there is not statistical evidence in favour of the convergence hypothesis. This is a positive sign though for their future within an enlarged euro area.

5. Conclusions

Focusing upon a group of CEECs which at some point will have to adopt the single currency is of policy relevance given that these countries do not have an opt-out clause and so will eventually relinquish control of monetary policy. If there is evidence of persistence in the inflation rate differential between a country and the Eurozone then this may lead to asymmetric macro shocks which could be difficult to deal with if there are large underlying differences in this key macro indicator between a specific country and the Eurozone. Whilst three out of the seven countries show persistence in their inflation rate differential to the Euro, employing fractional integration tests reveals that there are differences in the speed of adjustment in the inflation rates. Further tests reveal that the CEECs inflation rates converge to a common steady state. Out of the seven CEECs our findings imply that Bulgaria should delay adoption of the euro and there is evidence that Hungary and Romania may be vulnerable to asymmetric shocks.

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TABLE 1: GDP PER CAPITA IN PURCHASING POWER STANDARDS, INDEX EU-27 = 100

	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009	2010
Euro area (17 countries)	113	113	112	112	111	110	109	109	109	109	108	108	108
Belgium	123	123	126	124	125	124	121	120	118	116	115	117	119
Bulgaria	**27**	**28**	**28**	**30**	**32**	**34**	**35**	**37**	**38**	**40**	**43**	**44**	**44**
Czech Republic	**73**	**72**	**71**	**73**	**73**	**77**	**78**	**79**	**80**	**83**	**84**	**84**	**82**
Denmark	132	131	132	128	128	124	126	124	124	123	123	121	125
Germany	122	121	118	116	115	116	115	116	115	116	115	115	117
Estonia	42	43	45	46	50	55	57	62	66	70	69	64	64
Ireland	122	127	132	134	139	142	143	145	146	148	133	128	127
Greece	83	83	84	86	90	93	94	91	92	90	92	93	88
Spain	95	96	97	98	100	101	101	102	105	105	103	103	100
France	115	115	115	115	115	111	110	110	108	108	106	107	107
Italy	120	118	118	118	112	111	107	105	105	104	104	104	100
Cyprus	86	87	88	90	88	88	89	90	91	92	97	98	97
Latvia	**36**	**36**	**36**	**38**	**41**	**43**	**46**	**48**	**51**	**56**	**56**	**52**	**52**
Lithuania	**40**	**39**	**40**	**42**	**44**	**49**	**51**	**53**	**56**	**59**	**61**	**55**	**58**
Luxembourg	218	238	245	234	240	248	253	254	270	275	278	267	274
Hungary	**54**	**54**	**54**	**58**	**61**	**63**	**63**	**63**	**63**	**62**	**64**	**64**	**63**
Malta	80	81	85	79	81	80	78	78	76	76	78	80	83
Netherlands	129	131	134	134	133	129	129	131	131	132	133	131	133
Austria	132	132	132	126	127	128	128	125	126	124	124	125	126
Poland	**48**	**49**	**48**	**48**	**48**	**49**	**51**	**51**	**52**	**54**	**56**	**61**	**62**
Portugal	79	81	81	80	80	79	77	79	79	79	78	80	81
Romania	**27**	**26**	**26**	**28**	**29**	**31**	**34**	**35**	**38**	**42**	**47**	**46**	**45**
Slovenia	79	81	80	80	82	84	87	87	88	88	91	88	86
Slovakia	52	51	50	52	54	55	57	60	63	68	72	73	74
Finland	114	115	117	115	115	113	116	114	114	118	118	114	116
Sweden	123	126	128	122	122	124	126	122	123	125	123	119	123
United Kingdom	118	118	119	120	121	122	124	122	120	116	114	113	114

Source: *Eurostat*. Bold text denotes CEECs.

TABLE 2: LINEARITY AND UNIT ROOT TESTS RESULTS

Country	**W_λ**	**Sollis**	**Asymmetry**	**ADF**
Bulgaria	74.866**	2.087	–	–
Czech Republic	46.041**	11.102**	2.630**	–
Hungary	0.469	–	–	-3.100**
Latvia	14.042**	1.382	–	–
Lithuania	1.105	–	–	-1.714*
Poland	39.926**	2.059	–	–
Romania	74.214**	19.572**	5.124**	–

Note: The symbols * and ** indicate rejection of the null hypothesis at the 5 and 10% respectively. The lag length for the unit root tests has been obtained by means of the Modified Akaike Information Criterion proposed by Ng and Perron (1995). The critical values are as follows:

	$\chi^2(2)$	**Sollis**	**t-statistic**	**ADF**
5%	5.990	4.886	1.960	-1.942
10%	4.600	4.009	1.645	-1.615

TABLE 3: ESTIMATES OF THE FRACTIONAL DIFFERENCING PARAMETER

Country	PARAMETRIC	NON-PARAMETRIC	SEMI-PARAMETRIC		
	Robinson, 1994 White noise u_t	Robinson, 1994 Autocorrelated u_t	m = 5	m = 13	m = 20
Bulgaria	0.804* (0.681, 0.958)	0.679 (0.402, 1.094)	0.723	1.384**	1.303**
Czech Republic	1.143** (1.049, 1.270)	1.100 (0.920, 1.325)	0.898	1.179	1.477**
Hungary	1.006 (0.913, 1.130)	0.970 (0.798, 1.201)	0.705	1.038	1.341**
Latvia	1.096** (1.019, 1.195)	1.243** (1.063, 1.469)	0.671	1.431**	1.488**
Lithuania	0.940 (0.854, 1.054)	0.920 (0.760, 1.121)	1.116	1.342**	1.227**
Poland	1.002 (0.908, 1.129)	0.919 (0.774, 1.139)	1.048	1.182	1.493**
Romania	1.413** (1.296, 1.554)	1.394** (1.111, 1.822)	0.910	0.784	1.087

Note: The values in parenthesis in the second and third column refer to the 95% confidence band of the non-rejection values of *d* using Robinson's (1994) tests. The symbols * means evidence of mean reversion (i.e., *d*<1) and ** indicate rejection of the null hypothesis of *d*=1 in favour of the alternative of *d*>1. For the 3rd, 4rd and 5th columns the 95% confidence intervals corresponding to the I(1) hypothesis are respectively (0.632, 1.367), (0.771, 1.228) and (0.816, 1.184).

TABLE 4: ROBINSON (1995B) FRACTIONAL INTEGRATION TESTS

PANEL (A): Estimation of *d* for inflation rates

Country	Estimated *d*	Standard error	p-value
Bulgaria	0.444	0.690	0.00
Czech Republic	1.113	0.690	0.00
Hungary	0.829	0.690	0.00
Latvia	1.195	0.690	0.00
Lithuania	1.025	0.690	0.00
Poland	0.885	0.690	0.00
Romania	0.713	0.690	0.00
Eurozone	**0.911**	**0.690**	**0.00**

Note: F-tests for equality of *d* coefficients; F(7,728) = 11.818, Prob > F = 0.0000.

PANEL (B): Tests for equality of *d* coefficients for inflation rates vs the Eurozone

Pair	F	p-value
Bulgaria	35.944	0.000
Czech Republic	0.401	0.527
Hungary	8.040	0.005
Latvia	0.123	0.725
Lithuania	1.939	0.165
Poland	1.242	0.266
Romania	22.501	0.000

TABLE 5: FRACTIONAL INTEGRATION AND BREAKS

	No. of breaks	Break dates	d_1	d_2	d_3
Bulgaria	1	2000m7	0.798 (0.579, 1.119)	1.247^{**} (1.093, 1.453)	–
Czech Republic	2	1999m1 & 2008m1	0.927 (0.607, 1.353)	1.212^{**} (1.068, 1.415)	0.828 (0.645, 1.111)
Hungary	1	2007m1	1.049 (0.948, 1.186)	0.963 (0.772, 1.237)	–
Latvia	0	–	1.096^{**} (1.019, 1.195)	–	–
Lithuania	2	1999m10 & 2009m1	0.805 (0.579, 1.200)	1.085 (0.985, 1.230)	0.911 (0.709, 1.215)
Poland	1	2000m7	0.951 (0.771, 1.233)	0.959 (0.844, 1.112)	–
Romania	0	–	1.413^{**} (1.296, 1.554)	–	–

Note: ** indicate rejection of the null hypothesis of $d = 1$ in favour of the alternative of $d > 1$. d_1, d_2 and d_3 show the order of integration for each of the period(s) before the break(s) in the series.

FIGURE 1: BUDGET BALANCE IN % OF GDP, SEASONALLY ADJUSTED, QUARTERLY

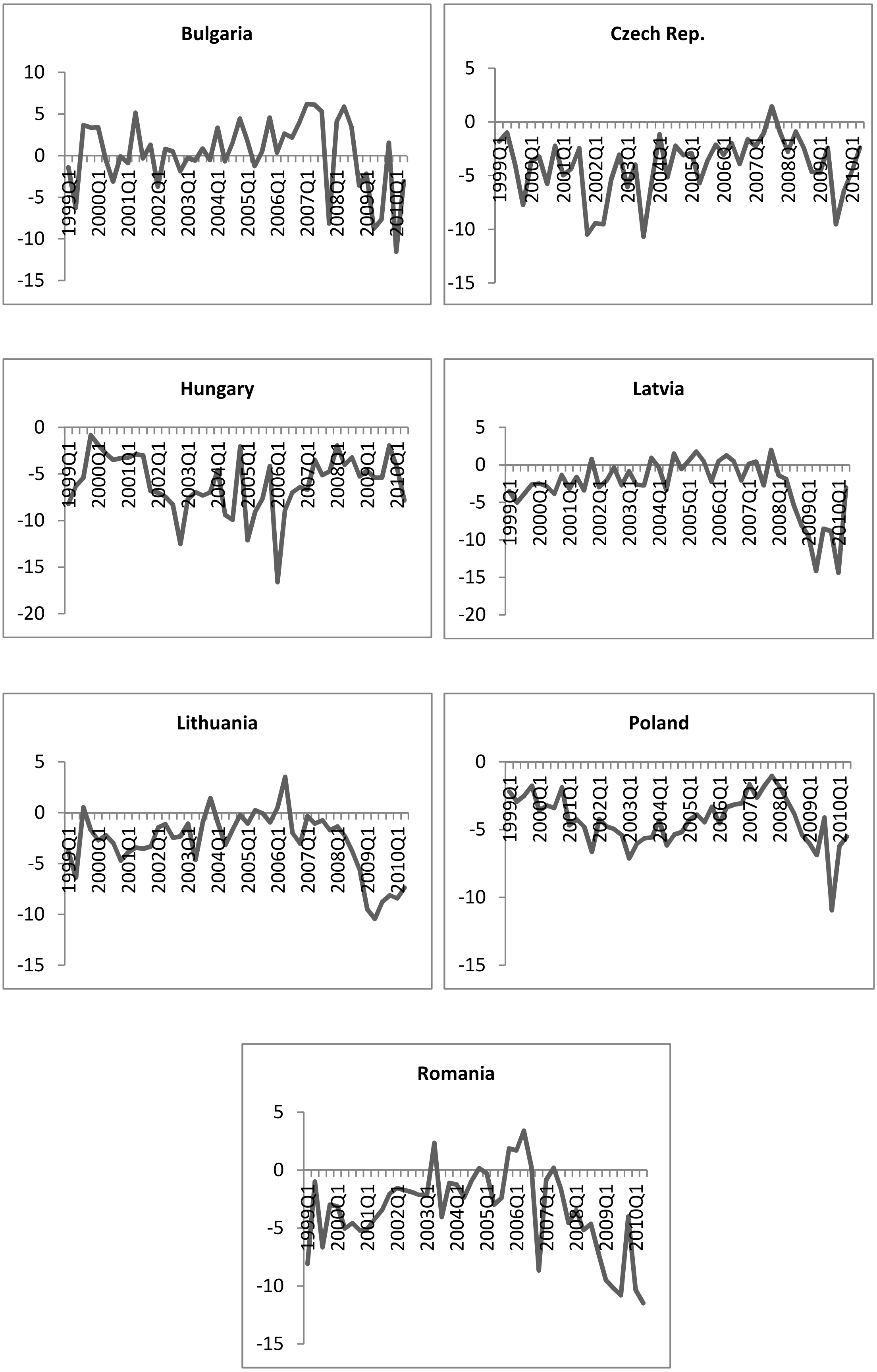
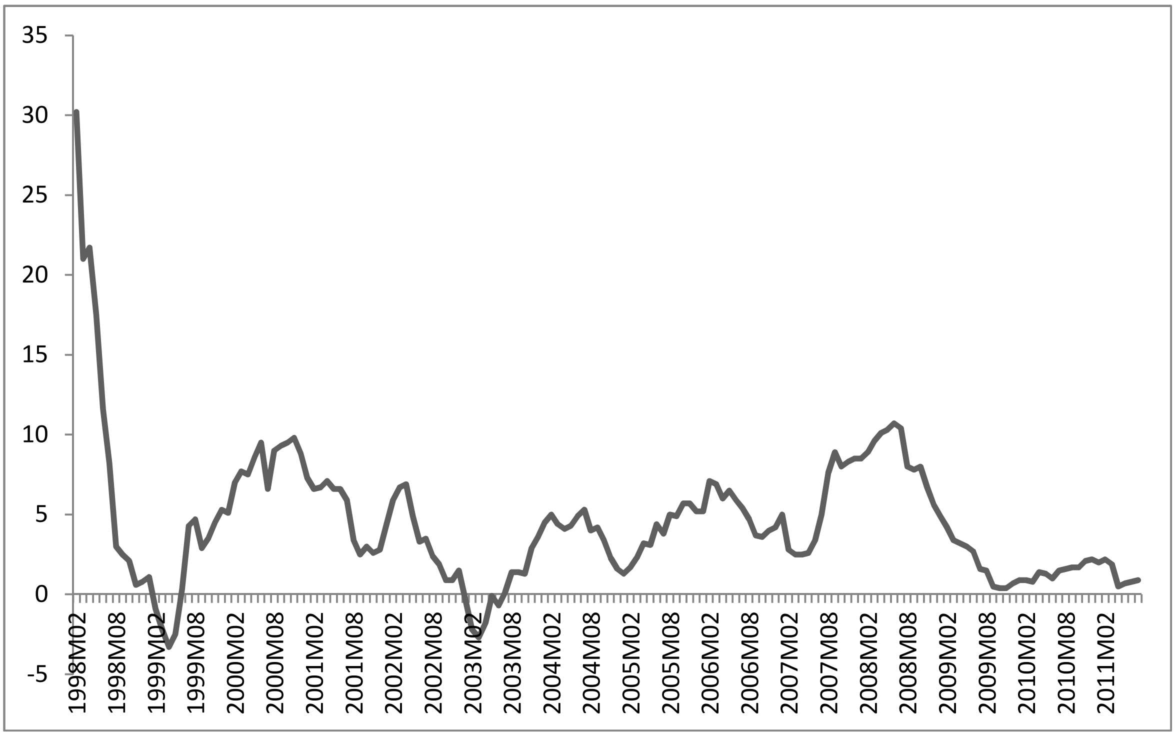
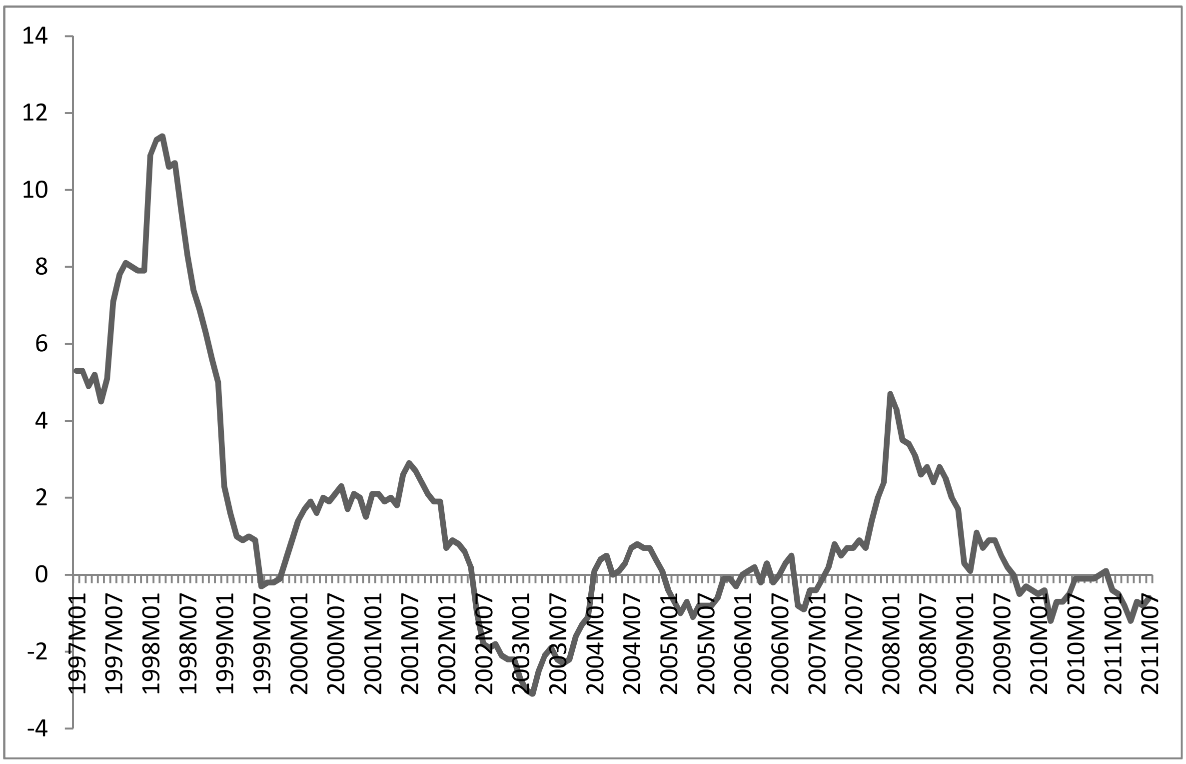


FIGURE 2: INFLATION DIFFERENTIALS

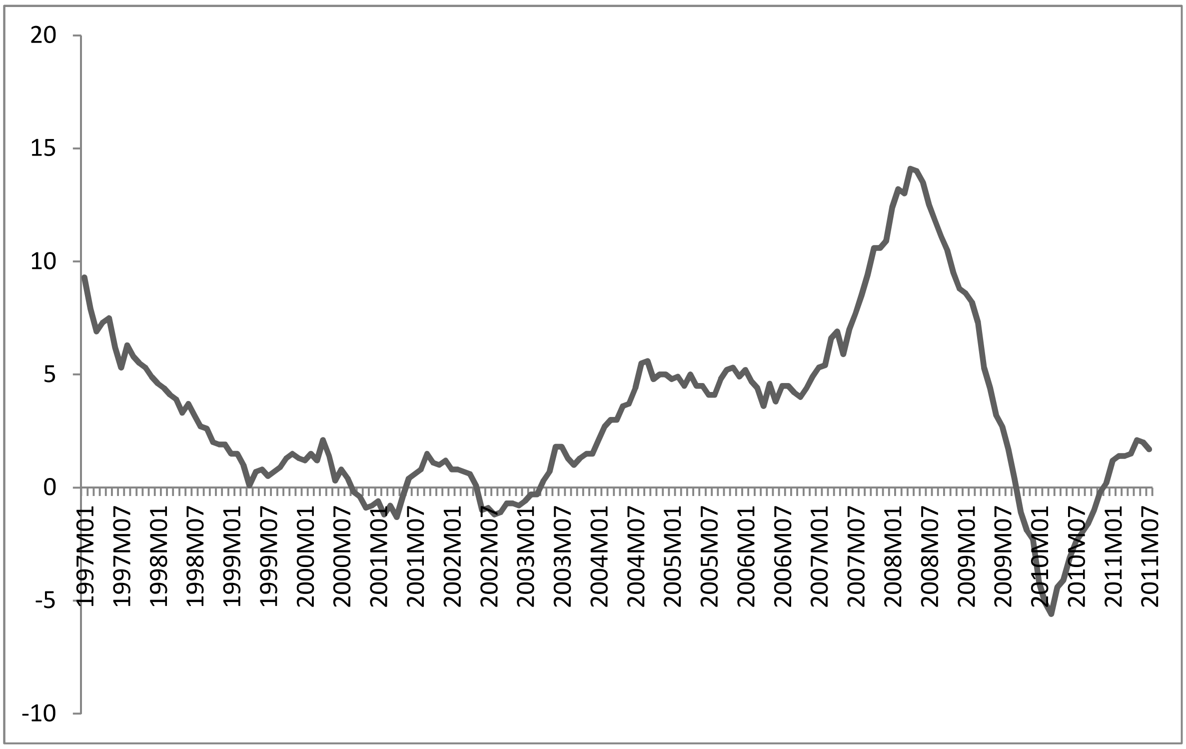
BULGARIA



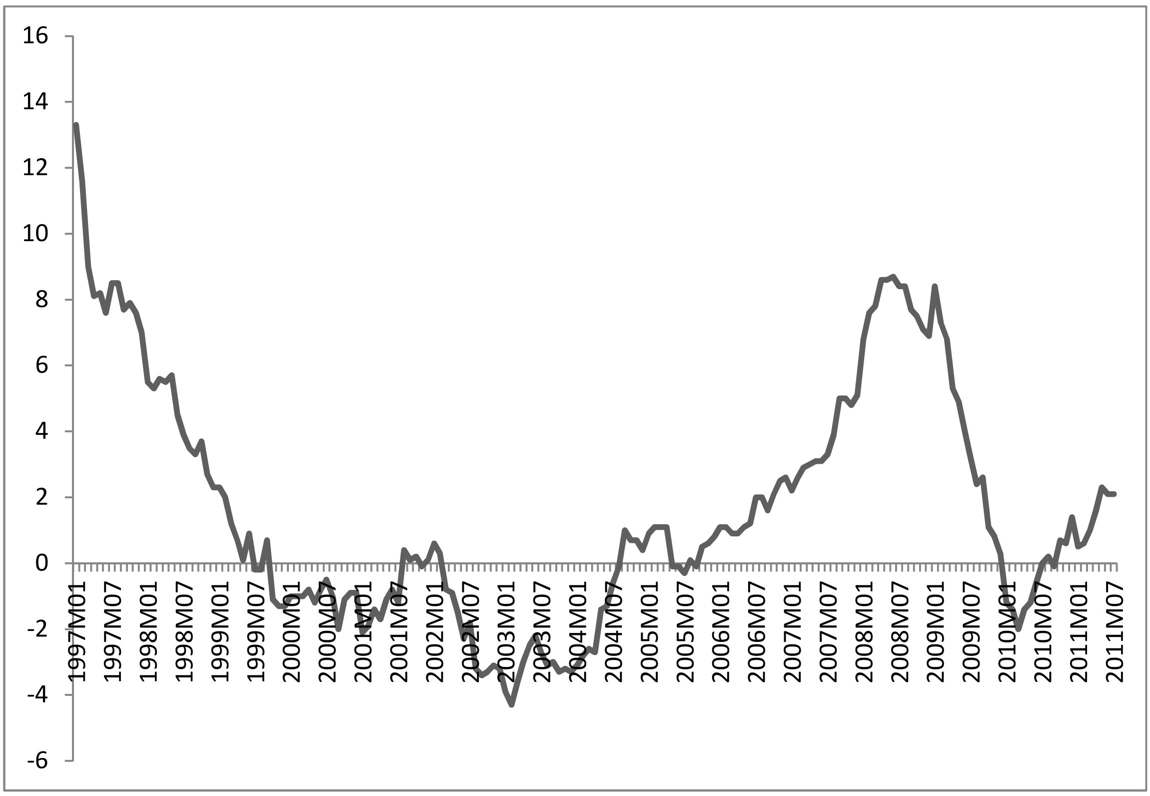
CZECH REPUBLIC



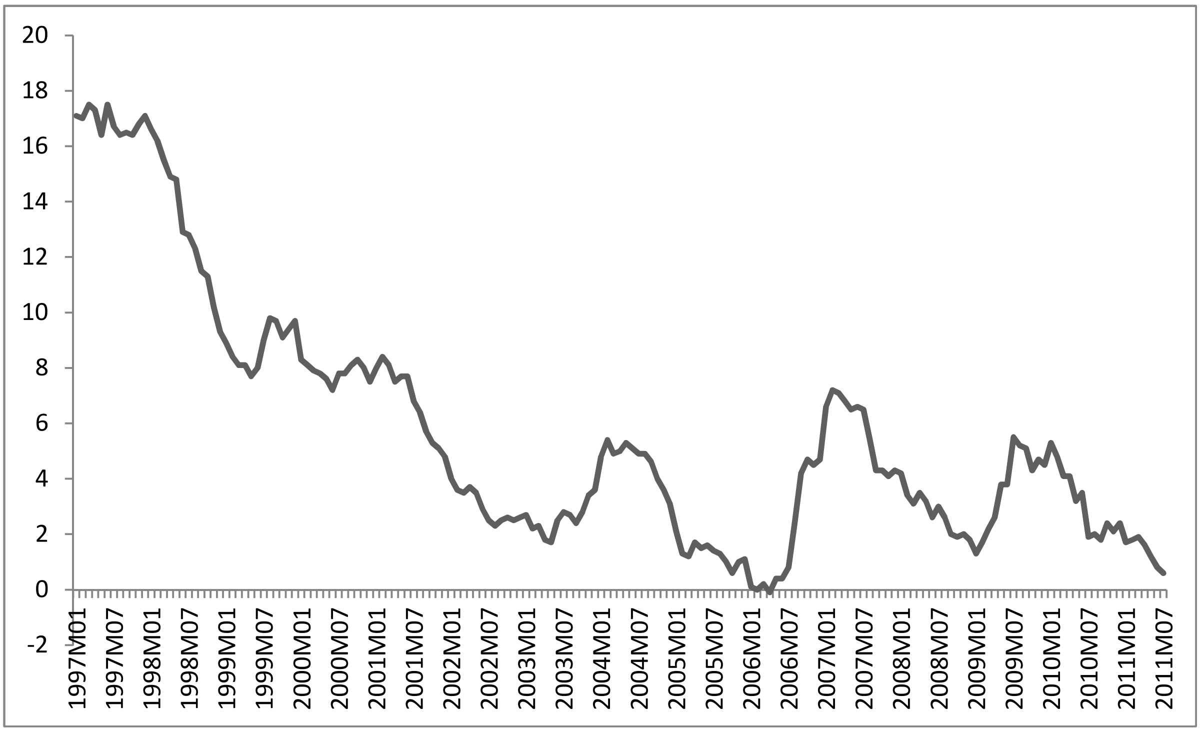
<u>LATVIA</u>



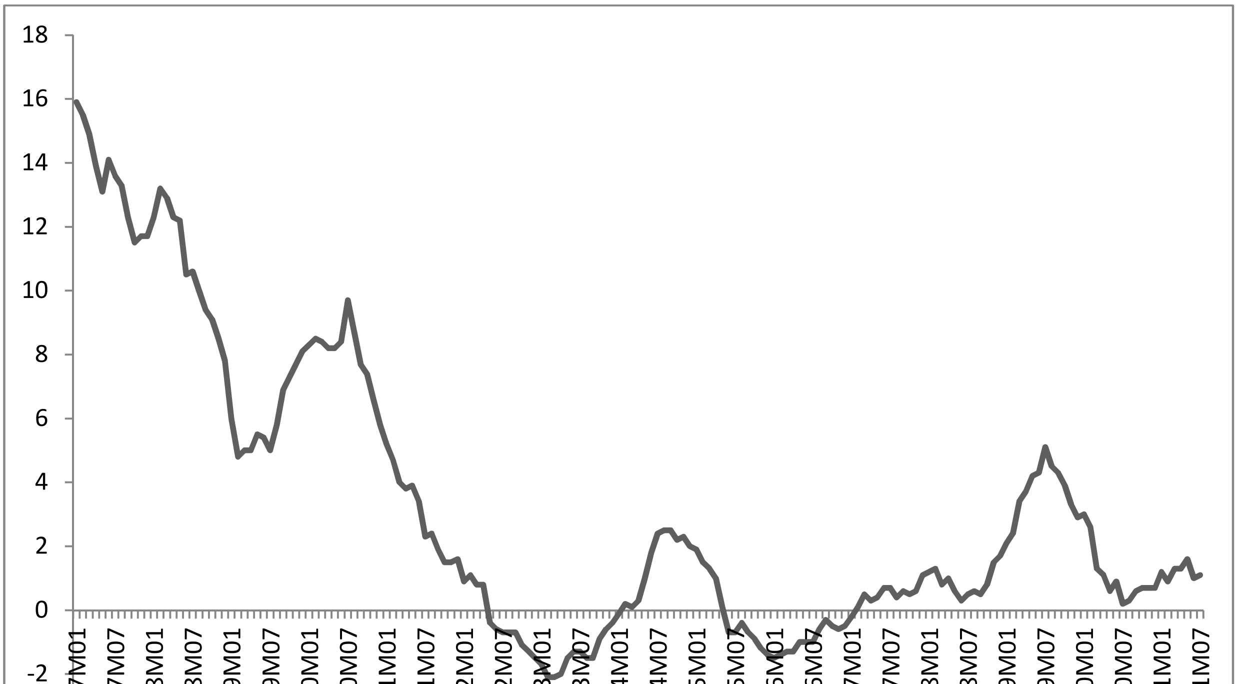
<u>LITHUANIA</u>



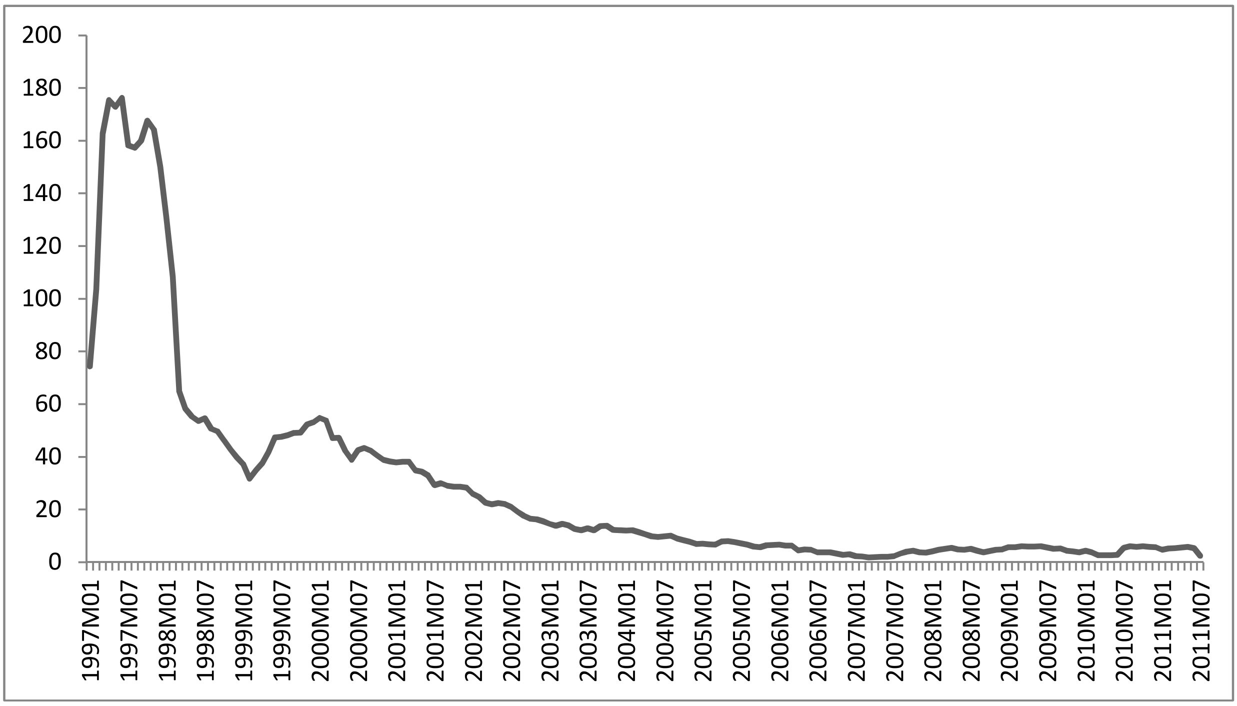
HUNGARY



POLAND



ROMANIA



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