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Inflation convergence in Central and Eastern Europe vs the Eurozone: non-linearities and long memory

Juan Carlos Cuestas*
Department of Economics
University of Sheffield
UK

Luis A. Gil-Alana
University of Navarra
Spain

Karl Taylor
Department of Economics
University of Sheffield
UK

Abstract

In this paper we consider inflation rate differentials between seven Central and Eastern European Countries (CEECs) and the Eurozone. We 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 2015 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 some 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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*Corresponding author: email address j.cuestas@sheffield.ac.uk, Department of Economics, University of Sheffield, 9 Mappin Street, Sheffield, South Yorkshire, S1 4DT, United Kingdom. Tel: +441142223410. Fax: + 441142223458

1. Introduction

In the wake of economic crises in some Eurozone countries in recent years, the merits of other countries being expected to join in the common currency is subject to renewed scrutiny. The focus of this paper is to test whether there is inflation convergence of the new member states from central and eastern Europe with the rest of the Economic and Monetary Union (EMU). We focus predominantly on those countries which have yet to join the Eurozone, but also include Latvia and Lithuania, given its recent adoption of the euro.[1] With this analysis we may be able to shed some light on the debate of whether or not it is a good idea to encourage more member states to adopt the common currency, in terms of the consequences of asymmetric shocks and the loss of monetary control to accommodate to them, in the Central and Eastern European Countries (CEECs). The issue of whether applying the same monetary policy to an area where different countries have different inflation rates may be detrimental for some economies is debatable. This argument is backed up by a number of authors such as Brissimis and Skotida (2008), who found that it is important that the European Central Bank (ECB) takes into account national characteristics when deciding monetary policy. This does not go against the common finding (see Lim and McNelis, 2007, amongst many others) that monetary policy should focus on inflation targeting.

Given the commitment from the ECB for price stability and the current target to control 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

[1] Latvia and Lithuania’s adoption of the euro occurred outside the sample period we focus upon, see section 4.1.

[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.

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. Most of the CEECs which are already member states still have to fulfil the Maastricht convergence criteria so as to be able to join the euro area.[3] Although only the first criterion focuses explicitly upon inflation control, the rest of the criteria have a direct link with the evolution of inflation expectations and, hence, inflation rate. For this reason, the focus of this paper is inflation convergence.

The sample of countries considered in this paper consists of CEECs which are member states but not part of the euro area, i.e. Bulgaria, the Czech Republic, Hungary, Poland and Romania. Latvia and Lithuania have also been included in the analysis, which may serve for comparison purposes. These countries are an interesting case study since, during the period analysed, they have been preparing for euro adoption.[4] Most of these countries joined the EU 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, when that happens, they are expected to adopt the single European currency.

In this paper we analyse the hypothesis 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. The process of transition from planned economies 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

[3] Details on the criteria can be found in http://epp.eurostat.ec.europa.eu/statistics_explained/index.php/Glossary:Maastricht_criteria.

[4] Latvia adopted the euro on January, 1st, 2014, whereas Lithuania did on January 1st, 2015.

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, we test for inflation convergence between each CEECs and the Eurozone by means of analysing the existence of unit roots in the inflation differentials for each country. We account for the possibility of non-linearities in the data generation processes (DGPs), which may affect the speed of convergence, and also take consider the possibility of fractional integration with potential breaks. Finally, we employ the recently developed club convergence test (Phillips and Sul, 2007), to explore the robustness of the results and gain additional insights from the analysis. In particular, we test whether these countries' inflation rates have been driven by a common trend during the period analysed.

The remainder of the paper is organised as follows. In the next section we summarise the recent developments in the process of European integration, paying particular attention to inflation convergence. 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. Background and literature review

Although the theory of optimum currency areas 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 differentials 'vs' the Eurozone. Mundell (1961) showed the importance of facing symmetric macro shocks in a currency union composed of different countries or regions.

In a recent contribution, Herz and Hohberger (2013) find that joining the monetary union exacerbates the vulnerability of the country to productivity shocks and increases the volatility of the real exchange rate and the current account. This may have an impact in the normal evolution of a country's inflation rate. Regarding the CEECs, Lehmann and Muravyev (2009) examine their labour markets in comparison to that of the Western Europe and reveal differences in terms of labour market policies and economic performance, which may hinder a uniform response to economic shocks.

Inflation expectations are a key macroeconomic variable when deciding the appropriate monetary policy to adopt.[5] This is the base of the Lucas critique; central banks need to enhance credibility. This is particularly relevant if a central bank wants to avoid the negative slope of the short run Phillips curve, when aiming to reduce inflation. 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, significant inflation rate differentials within an integrated monetary area can be seen as divergences in the path of 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, (see, Belke et al., 2013, and Cuestas et al. 2015, amongst others).

Given that the Maastricht criteria highlight the importance of controlling inflation, and that the ECB medium term inflation target is clearly defined,[6] we believe that inflation convergence is arguably the most appropriate and compelling means of assessing preparation to adopt the single currency.[7] However, fulfilling the inflation and the exchange rate criteria may pose problems. Given the rapid economic growth enjoyed in these countries during the transition process, the Balassa Samuelson effect may have

[5] 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.

[6] 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".

[7] Lewis and Staehr (2010) also show that enlargement of the EU may affect the reference inflation rate of the Maastricht criteria.

helped their real exchange rates to appreciate during this period (see Lewis, 2009 amongst others). This phenomenon is related to the Dynamic Penn Effect, whereby income and relative prices measured in common currency tend to share a common trend. The Balassa Samuelson and Dynamic Penn effects imply that the inflation rates of the CEECs and the Eurozone cannot converge, as long as the GDP growth rates of the CEECs are greater than that of the Eurozone. Hence, the countries need to decide how to deal with this real appreciation. The candidate country could fix the exchange rate and let the prices level increase, paying the price of higher inflation rates, or alternatively they could target inflation and let the nominal exchange rate absorb the pressure towards appreciation. The empirical literature, however, suggests that for the CEECs the Balassa Samuelson effect plays a very limited role in the apparent appreciation of the real exchange rate (Égert et al., 2003).

There are a number of authors who have tested whether there is evidence in favour of the inflation 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. 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. Gregoriou et al. (2011) claim that non-linearities are an important feature to consider in the DGPs for inflation differentials, and found more evidence towards the convergence hypothesis for the euro-12 countries, once non-linearities are incorporated in the analysis. Another interesting recent contribution is Lopez and Papell (2012) 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 inflation convergence, to the best of our knowledge, only a few contributions have analysed this issue. 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. His results are actually conditional to the individual countries' monetary policies and differences in economic development, which explain different results between countries. See also Kočenda et al. (2006). 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, Slovenia, 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.[8] Finally, Staehr (2010) finds evidence supporting the hypothesis of price convergence within the ten new EU countries from central and eastern Europe.

We expand on this analysis by focusing on a long time period 1997 through to the latest data available in 2015, incorporating tests for non-linearities in inflation differentials and if this hypothesis is not rejected, non-linearities are included in the formulation of the auxiliary regression for unit root testing. We argue the importance of accounting for potential non-linearities in inflation rate differentials below in section 3.1.

[8] 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.

3. Empirical methods

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 $y_t = \pi_{i,t} - \pi_{euro,t}$, with $\pi_{i,t}$ and $\pi_{euro,t}$ denoting 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,

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

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 differentials are stationary or if they revert to zero.

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 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 larger 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. Also, the speed of adjustment towards equilibrium may depend on the sign of the shocks. We tests for these possibilities.

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). See appendix for further details.

Hence, in order to test for the convergence hypothesis using tests for the order of integration of the variables, we first try to gain some insights on whether or not the variables follow a linear or non-linear process.

3.2 Unit root tests

Our interest lies on analysing the existence of unit roots (no-convergence) in the inflation differentials, 'vs' the hypothesis of stationarity (convergence). Hence, 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 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 control 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 may be the case that an increase in the inflation rate is more difficult to tackle than a reduction below the target, or vice versa. In such a scenario, we would expect that the speed of mean reversion would differ depending on the sign, not only the size or magnitude, of the shock.

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 an integer, typically 0 or 1. Fractional integration tests break the dichotomy of d equals to 1 or 0, 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 ≥ 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 magnitude of d.

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, 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 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

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. With this approach we should gain some insights into the existence of commonalities between our target countries, in terms of their inflation rates' evolution. Full details on the empirical methodologies adopted in 3.1 to 3.3 are provided in the appendix.

4. Empirical Evidence

4.1 The data

The inflation differentials are computed as the difference between the inter-annual inflation rate of each of the CEECs 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 2015:5, except Bulgaria, whose sample starts in 1997:12.

[Figure 1 about here]

The plots of the inflation differentials versus that of the Eurozone are displayed in Figure 1. 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. Some countries performed a rapid increase in prices while they maintained fixed exchange rates (the Baltic States and Bulgaria), whereas other countries showed much lower inflation rates and appreciating nominal exchange rates (Poland and the Czech Republic). Also, we observe a significant increase in the Baltic States' inflation rates in 2008 and the beginning of 2009, which was mainly caused by food prices and housing expenses. This episode preceded a drop in the inflation rates, due to the financial downturn, which was more damaging for their aggregate demand than in other countries. In general the plots suggest that there is evidence of co-movement in the inflation rate differentials with respect to that of the Eurozone, which may be an indication of lack of asymmetric shocks affecting the inflation rates of these countries. From Figure 1, it can be seen that although there is more or less a clear pattern towards a zero differential, we observe that the line does not cross the zero line frequently. This implies that still there is a gap between these countries' inflation rates and that of the Eurozone. Hence, whether this is actually convergence and an equilibrium reached needs to be formally tested. The results from this analysis are summarised in the next section.

4.2 Results

As aforementioned, we aim to test the hypothesis of inflation convergence we proceed as follows; firstly, in order to choose a linear unit root test or a non-linear one, we first test the null of linearity, by means of the Harvery et al. (2008) test. Secondly, if the test rejects the null of linearity, then we apply the Sollis (2009) test, and if the test cannot reject the null of linearity we apply the ADF test. After that, in order to gain additional flexibility in the analysis of the order of integration we apply fractional integration techniques, including structural breaks. Finally, we test for common trends in the CEECs, by means of applying the Phillips and Sul (2007) test. The results of the Harvey et al. (2008), Sollis (2009) and ADF tests are presented in Table 1. 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.

[Table 1 about here]

First, we start by testing the hypothesis of linearity of the inflation differentials for each country. According to the second column of Table 1, 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 1, the null of a unit root cannot be rejected for three of our target countries; Bulgaria, Latvia and Romania. For the rest of the countries, the results indicate that the inflation differentials are non-linear and globally stationary, implying convergence between each of the aforementioned countries and the Eurozone and, in particular for the Czech Republic, where shocks are found to have asymmetric effects. The latter finding means that shocks with a different sign but of equal magnitude will have a different impact, in absolute terms, on the target variable. The implications such asymmetric effects stemming from shocks are of policy relevance. Monetary authorities would have to act with caution and different strength when attempting to decrease the inflation rate, than when the aim is to increase it. This has been observed to create difficulties in the application of expansionary or contractionary monetary policies.

[Table 2 about here]

Next we examine the possibility of fractional integration. As mentioned earlier 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 2 displays the estimates of d for each individual series. Recall that if d is less than 0.5, the variable is stationary and mean reverting; if it is greater than 0.5 but strictly smaller than 1, the variable is non-stationary but mean reverting, and if greater than or equal to 1, the variable is not stationary and non-mean reverting. The first two columns in Table 2 refer to the Whittle estimates of d, displaying 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).[9] The last two columns refer to the semiparametric Whittle method of

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

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 2 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 the Czech Republic, Latvia and Rumania with some of the specifications. For the remaining cases, we cannot reject the null of I(1) behaviour or, if it is rejected, it is in favour of higher degrees of integration.

[Table 3 about here]

In addition, we present, in Table 3, 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. This is why the test is applied to the inflation rates and not the differentials. 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 inflation rates are unit root processes. 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 to that of the Eurozone's inflation rate. The hypothesis of equality of d with respect to that of the Eurozone cannot be rejected for the Czech Republic, Hungary, and Latvia, implying that Bulgaria, Lithuania, Poland and Romania's inflation rates d are not similar to the d of the euro area's rate of inflation. Hence, shocks affecting the inflation rates seem to have similar effects in the Eurozone, the Czech Republic, Hungary and Latvia. This implies that a common shock over these countries and the Eurozone would tend to disappear more or less at the same time, which of course it is good news for policy coordination.

These results reinforce 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 Lithuania's inflation differential appears to be stationary according to the unit root analysis (Table 1), the results in Table 3 indicate that inflationary shocks experienced by this country tend to disappear faster than in the euro area, which is consistent with our findings in Table 2. 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 1 for the case of Latvia. However, the results presented in Table 3 Panel (b) indicate similar orders of integration in her inflation rate to that of 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 Latvia and the Eurozone. The result is interesting since during our sample period Latvia had been preparing for euro adoption, joining on January, 1st, 2014, which may explain this convergence pattern. Poland on the other hand, has been in charge of her monetary policy during the period analysed, which has given the country the opportunity to accommodate asymmetric shocks and reduce her inflation rate.

[Table 4 about here]

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 artefact 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). Table 4 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.[10] 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 date across countries, i.e. around the time of the creation of the euro and also 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 ERM II (for the case of Lithuania). In all cases, it appears that the 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 slightly decreased the speed of mean reversion. The latter phenomenon can be explained 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, to assess whether the inflation rates of the CEECs potential euro candidates 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 (see appendix equation 17) with the critical value -1.65, for different groups of countries. In our case the t-statistic is -1.47, which is greater than the critical value, when Bulgaria and Poland are excluded; hence we cannot reject the null

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

hypothesis that this group of countries, with the exception of Bulgaria and Poland, form a convergence club.

5. Conclusions

Focusing upon a group of CEECs which at some point in the future are expected to adopt the single currency is of policy relevance. Specifically, analysing how similar the evolution of their inflation rates 'vs' the Eurozone is timely given that these countries do not have an opt-out clause. 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.

In order to investigate this issue, we explicitly test for the dispersion of inflation, using time series econometrics, both for the inflation differentials and inflation rates of each country. We account, in particular, for the order of integration taking into account a number of different data generation processes, namely non-linear and fractional integration. Whilst some of the 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. 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 fact the Bulgaria has had a currency board has not facilitated the adjustment process. The results point against the convergence hypothesis and the order of integration of Bulgaria's inflation rate is much lower 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,[11] which basically implies that losing for their monetary policy and exchange rate management

[11] However, Lithuania has kept a currency board for a number of years.

will not, in principle, pose major problems in the case of asymmetric macro shocks. Hungary and Romania are interesting case studies. Both countries inflation rates' appear to have converged to the inflation rate of the Eurozone, however, 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 no statistical evidence in favour of the convergence hypothesis. This is a positive sign though for their future within an enlarged euro area. 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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Appendix

Harvey et al. (2008) test

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:

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

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:

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

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:

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

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.[12]

Sollis (2009) unit root test

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

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

where

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

and

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

Hence, the null hypothesis of a unit root can be specified as $H_0: \gamma_1 = 0$. One problem is that, under the null hypothesis some of the parameters cannot be identified. Sollis (2009), by means of Taylor approximations, proposes testing for unit roots in this nonlinear framework using the following auxiliary equation:

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

In this context, testing for a unit root in model (11) implies testing the null of $H_0: \beta_1 = \beta_2 = 0$. 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 $H_0: \beta_2 = 0$, using any standard test of hypotheses.

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

Fractional integration

Fractionally integrated or I(d) models are as follows:

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

where u_t is a stationary process and L stands for the lag operator, i.e. $\mathrm{L}y_t = y_{t-1}$. 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 the 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 for different countries, which will provide some insights about the degree of homogeneity of persistence of shocks on the variables.

Phillips and Sul (2007) club convergence approach

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}:

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

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

$$h_{it} = \delta_{it} \Big/ \frac{1}{N}\sum_{i=1}^{N}\delta_{it} \tag{15}$$

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,

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

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:

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

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. The method can be used to identify clubs of convergence if the null of overall convergence is rejected for the whole panel.

TABLE 1: LINEARITY AND UNIT ROOT TESTS RESULTS

Country	W_λ	Sollis	Symmetry	ADF
Bulgaria	83.685**	0.906	–	–
Czech Republic	55.222**	12.692**	2.713**	–
Hungary	3.718	–	–	-3.825**
Latvia	11.149**	1.863	–	–
Lithuania	1.522	–	–	-2.555**
Poland	50.916**	13.013**	0.195	–
Romania	94.125**	3.460	-	–

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 Akaike Information Criterion. 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 2: 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	1.52** (1.34, 1.74)	Convergence not achieved	0.990	1.031	1.039
Czech Republic	1.16 (1.07, 1.27)	1.17 (0.97, 1.39)	0.500*	0.745	1.500**
Hungary	1.16** (1.07, 1.26)	1.20** (1.01, 1.46)	1.021	0.821	1.208**
Latvia	1.27** (1.20, 1.36)	1.42** (1.25, 1.63)	0.524*	1.235**	1.500**
Lithuania	1.16** (1.07, 1.24)	1.19** (1.06, 1.36)	1.086	1.271**	1.420**
Poland	1.24** (1.14, 1.35)	1.23** (1.08, 1.44)	0.954	0.968	1.376**
Romania	1.55** (1.40, 1.72)	0.87 (0.63, 1.32)	0.523*	0.593*	1.046

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, 4th 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 3: ROBINSON (1995B) FRACTIONAL INTEGRATION TESTS

PANEL (A): Estimation of d for inflation rates

Country	Estimated d	p-value
Bulgaria	1.033	0.00
Czech Republic	1.003	0.00
Hungary	0.850	0.00
Latvia	1.067	0.00
Lithuania	0.885	0.00
Poland	0.913	0.00
Romania	0.689	0.00
Eurozone	**1.033**	**0.00**

Note: F-tests for equality of d coefficients; F(7,880)=19.716, Prob > F=0.0000.

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

Pair	F	p-value
Bulgaria	56.348	0.000
Czech Republic	0.815	0.367
Hungary	2.807	0.095
Latvia	1.228	0.268
Lithuania	9.036	0.002
Poland	4.877	0.028
Romania	5.017	0.026

Note: Test for equality of d for Eurozone, Czech Republic and Latvia F(3,508)=0.849, Prob > F = 0.4676

TABLE 4: FRACTIONAL INTEGRATION AND BREAKS

	No. of breaks	Break dates	d_1	d_2	d_3
Bulgaria	1	2000m7	0.37 (0.05, 1.06)	1.22** (1.09, 1.38)	---
Czech Republic	2	1999m1 & 2008m1	1.27** (1.03, 1.24)	1.13 (1.00, 1.29)	0.93 (0.81, 1.07)
Hungary	1	2007m1	1.24** (1.13, 1.39)	1.00 (0.88, 1.16)	---
Latvia	0	–	1.27** (1.20, 1.36)	---	---
Lithuania	2	1999m10 & 2009m1	0.64 (0.51, 0.87)	1.07 (0.97, 1.21)	1.22** (1.09, 1.37)
Poland	1	2000m7	1.30** (1.09, 1.58)	1.24** (1.15, 1.34)	---
Romania	0	–	1.55** (1.40, 1.72)	---	---

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: INFLATION DIFFERENTIALS