Euro Area Inflation Differentials: Unit Roots and Non-Linear Adjustment

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Abstract

This paper examines the time series properties of inflation differentials in twelve EMU countries. Unlike previous studies, we focus on inflation deviations from the target, given by the Maastricht reference value and the ECB target. The results from linear unit root tests indicate that inflation misalignments are non-stationary. When we utilize a non-linear mean reverting adjustment mechanism for inflation differentials, we discover that although deviations of inflation from the target can exhibit a region of non-stationary behavior, overall they are stationary. Our non-linear findings of stationarity, provide evidence in favor of the EMU being an optimal currency area.

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1. Introduction

In June 1989 the European Council decided that the first stage towards European Monetary Union (EMU) should begin in July 1990[1]. The Treaty of Maastricht was agreed by the heads of state of the European Union (EU) in December 1991 setting out the framework for stages two and three of progress towards EMU. Considering price stabilization, the Maastricht Treaty requires that inflation is not greater than 1.5% from the average of the three lowest inflation rates in the EU, in order for a country to qualify for the third stage[2]. The Maastricht criteria were set in an effort to promote the convergence of the prospective eurozone economies in the lead-up to euro, and the effectiveness of common monetary policy. A well-known result of the Optimal Currency Area (OCA) literature is that a common central bank is better equipped than a national central bank when it comes to deal with economic shocks that are homogeneous across countries (see among others, Giavazzi and Giovannini, 1989). According to the Treaty, during the third stage, the primary objective of the European Central Bank (ECB) is price stability, which the ECB has interpreted as an annual Euro area inflation rate below, but close to, 2% in the medium run[3].

While post-euro, the eurozone countries' inflation rates are not explicitly bounded by the Treaty, and ECB itself admits that "monetary policy can only influence the price level of the Euro

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[1] In the first stage, the members of the European Monetary System (EMS) abolished all remaining capital controls. Also, there was an increase in the degree of co-operation among the EMS central banks, while exchange rate realignments remained possible. The second stage started on 1/1/1994. During that stage, the European Monetary Institute, the precursor of the European Central Bank, was created. In order to participate to the third stage, which started on 1/1/1999 (apart from Greece where it started on 1/1/2001) countries had to satisfy five convergence criteria.

[2] The convergence protocol states: "The criterion on price stability referred to in the first indent of Article 121(1) of this Treaty shall mean that a Member State has a price performance that is sustainable and an average rate of inflation, observed over a period of one year before the examination, that does not exceed by more than 1½ percentage points that of, at most, the three best performing Member States in terms of price stability. Inflation shall be measured by means of the consumer price index on a comparable basis, taking into account differences in national definitions".

[3] The Treaty states that the ECB should also be concerned with output and employment, albeit without prejudicing its main objective of price stability. The monetary policy framework adopted by the ECB to fulfill these tasks is based on two analytical perspectives or two *pillars*, namely economic analysis and monetary analysis. The ECB has repeatedly stated that achieving price stability is the most effective way to contribute to output and employment growth.

area as a whole and cannot affect inflation differentials across regions" (see *The Monetary Policy of the ECB*, 2004), national inflation rates should not, nevertheless, diverge considerably and persistently from the ECB target of 2%. Since the ECB sets the nominal interest rate according to the Euro-wide inflation rate, persistence in inflation differentials may imply that 'one size does *not* fit all'. Diverging national inflation rates and common monetary policy imply that every member has to face a different real interest rate: countries with strong economic growth and high inflation[4] benefit from lower rates providing further stimulation to their economy; the reverse happens to low growth economies that experience low inflation rates[5]. It has been also argued that persistent divergence of national inflation rates hampers the efficient communication of monetary policy and complicates the design of an optimal policy response (see Benigno, 2003).

Within this context, the time series properties of eurozone countries inflation deviations from the policy reference value implied by the Maastricht convergence criterion and the ECB target are vital. Discovering that inflation misalignments are characterised by unit root behaviour can suggest that the idiosyncratic shocks impacting upon individual countries' inflation rates have persistent effects, which raises issues on whether EMU really constitutes an OCA to be effectively managed by the ECB. In addition, it raises the question of whether the member countries truly converge during the pre-euro period. On the other hand, finding that inflation misalignments are only temporary, part of a rebalancing process between fast-growing and slow-growing regions, which would be characterised by a stationary process, would imply that the

[4] There is sufficient empirical evidence indicating that inflation and demand pressures, as measured by the output gap, are positively related in the euro area countries (see e.g. ECB, 2003).

[5] This statement should be treated with caution. As Busetti *et al.* (2006) argue, if inflation differentials are due to differences in administered prices, or due to different import prices and/or wage growth that don't affect profit margins, then the resulting differences in real interest rates may affect private consumption, but should not impact upon investment. In fact, the elasticity of investment expenditure with respect to the real interest rate will depend on the degree of market integration within the EMU. Another argument against the procyclicality of inflation differentials works through the real exchange rate channel. According to this view, given that the nominal exchange rate is fixed, inflation differentials are part of a countercyclical adjustment mechanism: the competitiveness of countries with high inflation declines, therefore reducing economic activity. Busetti *et al*. (2006) point out that the answer to whether inflation differentials are procyclical or countercyclical will largely depend on the magnitude and persistence of inflation misalignments.

ECB can effectively implement and communicate its policies without exacerbating the differences that exist between EMU countries. Therefore, the degree of persistence of inflation misalignments is of primary importance in order to establish whether the economic area exhibits imbalances which require structural interventions, or whether the asymmetries are just temporary phenomena which in the long run eliminate themselves.

Some previous empirical evidence indicates that in the run-up to the single currency the dispersion of inflation rates across the prospective eurozone members has decreased, reaching the lowest point during 1999, it increased in 2000 and remained fairly stable since 2001[6]. Honohan and Lane (2003) suggest that the differential impact of the euro depreciation during the first years of the single currency may have caused higher inflation differentials. Contrary to inflation differentials within the United States, and among regions of individual euro area countries, inflation differentials across euro area countries are more persistent, with most countries' inflation rates persistently below (e.g. Germany, Austria) or above (e.g. Greece, Ireland) the euro area average since 1999 (see e.g. ECB, 2003).

Various explanations have been suggested for the persistence in euro area inflation differentials. In the context of the Balassa (1964) and Samuelson (1964) effect, persistent national inflation differentials within a monetary union may be associated with the process of real convergence: higher productivity growth in the tradable sector of the low income countries results into higher real wages in both their tradable and non-tradable sector implying higher overall inflation. Recent empirical evidence however, indicates that the Balassa-Samuelson effect alone cannot fully account for the observed persistence in inflation differentials (see e.g. Rogers, 2002). Another set of explanations focuses on the interaction between nominal and real rigidities,

[6] See among others, Duarte (2003) and ECB (2003).

structural differences, price and wage setting, and common or idiosyncratic shocks[7]. Angeloni and Ehrmann (2004) suggest that the level of inflation persistence in each member country largely determines the persistence of inflation differentials in the euro area[8].

A number of papers have applied various unit root and cointegration tests to analyze the persistence of inflation differentials in the euro area. An early contribution to the literature is Siklos and Wohar (1997) who find evidence of a single stochastic trend (i.e. evidence of convergence) for the time period 1974-95. Kocenda and Papell (1997) also report evidence of inflation convergence during the pre-euro period using panel Augmented Dickey Fuller (ADF; Dickey and Fuller, 1979) unit root tests[9]. Busetti *et al.* (2006) apply univariate and multivariate unit root tests on bilateral inflation differentials and agree that the pre-euro (1980-1997) period is characterised by convergence (stationary differentials). They also provide evidence of diverging behaviour following the introduction of the Euro[10]. Rodriguez-Fuentes *et al.* (2004) use the ADF, ADF with GLS detrending (Elliot *et al.*, 1996), the Elliot, Rothemberg and Stock optimal point (Elliot *et al.*, 1996), Phillips-Perron (Phillips and Perron, 1988) and KPSS (Kwiatkowski *et al.*, 1992) tests to investigate whether inflation differentials exhibit a unit root over the period 1980-1998. In contrast to previous studies that suggest convergence over the pre-euro period, Rodriguez-Fuentes *et al.'s* (2004) results indicate that national inflation deviations from the euro area inflation are non-stationary in eight out of eleven EMU countries.

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[7] For example, differences in the economic structure can result in diverse propagation of the various shocks: an industry-focused economy with low availability of raw materials will face higher inflation as a consequence of higher oil prices. Contrary to that, a country whose economy is mainly based on services may have a small impact from the same price hike.

[8] A number of empirical papers document that euro area inflation is inertial and responds sluggishly to changes in monetary policy. See for example, Angeloni *et al.* (2005) who document that inflation persistence in the euro area did not decline after the introduction of the euro.

[9] They compute the inflation differential as the difference between an individual inflation and the average for all the countries.

[10] Their results suggest the existence of two clusters within the EMU: one of high inflation countries and another of low inflation countries. There is stationarity amongst the countries belonging to each cluster, but divergence between the two clusters.

This paper contributes to the existing literature in two important aspects. First, unlike previous studies which consider countries in pairs, or employ the deviation of national inflation from the euro area value, we make use of the inflation differential between each country and the policy reference value implied by the Maastricht convergence criterion and the ECB target. We consider this to be a more appropriate approach to model inflation differentials. This is because during the pre-euro period, national inflation rates were explicitly bounded by the Treaty, and in the post-euro period, the inflation rate of each EMU country should not diverge considerably from the ECB target of 2%.

Second, we present new empirical evidence, which explicitly allows for the possibility that inflation can be characterized by a non-linear mean-reverting process. This process may exhibit near unit root behaviour in a specific range, so inflation deviations from the policy reference value can appear non-stationary from the perspective of test procedures, which specify a linear non-stationary process as the null hypothesis. We propose an alternative hypothesis where the speed of adjustment increases, the greater the deviation of inflation from the policy reference value. Essentially, the more distant inflation is from the target, the greater the probability that remedial action will be taken to revert inflation back to the policy reference value. Gregoriou and Kontonikas (2006) have applied non-linear mean-reverting unit root tests to inflation deviations in a sample of seven inflation targeting countries. The motivation of their study stems from a new class of monetary policy models that relax the assumptions of the conventional linear-quadratic preferences framework in an effort to introduce non-linearities in the response of monetary policy to inflation[11].

[11]Orphanides and Wieland (2000) point out that if the central bank assigns at least some weight to output stabilization, the output stabilization objective will dominate when inflation is within the targeted zone, and inflation stabilization will dominate when inflation deviations from the target become large.

The rest of the paper is structured as follows. The next section describes the dataset. Sections 3 and 4, respectively, outline the linear and non-linear unit root testing framework and results. Section 5 concludes.

2. Data

Our dataset comprises of twelve EMU countries that adopted the euro and common monetary policy: Austria, Belgium, Finland, France, Germany, Greece, Italy, Ireland, Luxembourg, Netherlands, Portugal and Spain. The sample period is January 1996 - December 2005 effectively covering (part of) the second and the third stage of the process towards monetary union. Since both the inflation convergence criterion and the ECB reference value monitor the evolution of annual inflation, we measure inflation, π_t, as the twelfth difference of the natural log of the monthly Harmonised Consumer Price Index (HCPI): $\pi_t = 100 * (\ln HCPI_t - \ln HCPI_{t-12})$ [12]. Our sample includes 120 observations for each country. The data was collected from Datastream.

[INSERT TABLE 1 HERE]

In line with the policy objectives, during the period January 1996 to December 1998 the reference value for inflation, π_t^*, is calculated (year-per-year) as 1.5% plus the average inflation of the best three performing countries in terms of inflation control. For instance, we observe from Table 1, the three best performers during 1996 were Sweden (0.78%), Finland (1.06%) and Luxembourg (1.16%). Thus, the policy reference value for 1996 is

[12] Both the Maastricht Treaty inflation criterion and the ECB target are calculated by using annual HCPI inflation. The HCPI is a statistical indicator whose objective is to reflect the focus of the general public on the consumer goods prices, and to provide a common measurement of inflation which facilitates carrying out international comparisons. The HCPI series starts at January 1995 thereby providing the first observation of annual inflation at January 1996. In this paper, we also experimented using the conventional Consumer Price Index inflation that allows us to perform our empirical analysis from the start of the second stage towards the EMU (January 1994) thereby adding 24 observations to our sample for each country. The results (not reported and available upon request), remain the same.

$\frac{1}{3}\left(0.78+1.06+1.16\right)\%+1.5\%=2.5\%$. In a similar fashion we calculate the policy reference value for 1997 and 1998 as 2.7% and 2.19%, respectively. In the case of Greece, where stage two of the process towards EMU lasted until December 2000, we also calculate the policy reference value for 1999 and 2000 as 2.04% and 2.82%, respectively. From January 1999 until the end of the sample, we set π_t^* equal to the ECB target of 2%. We acknowledge that the ECB target of 2% concerns the Euro area as a whole[13], but nevertheless individual member countries inflation rates should converge around this reference value. In particular, national economic policies (fiscal, structural, wage-setting) in Euro zone should be employed to deal with persistent inflation differentials (Weber, 2004). Otherwise, changes in the Euro-wide nominal interest rate may be translated into diverse real interest rate changes thereby hampering the efficiency of ECB policies in stimulating Euro-area economic growth.

Table 1 reveals some interesting patterns for the inflation rates of the EMU member countries. Greece had the highest inflation, which was nevertheless declining in the effort to join the EMU, until 1998. Thereafter, during four of the remaining seven sample years, Ireland had the highest inflation rate, a sign of overheating in the Irish economy. Greece and the other three club-Med countries (Italy, Portugal and Spain) typically feature in the list of high-inflation countries. Over time, the list of best performers usually includes two of the core EMU countries, France and Germany, as well as Austria, Finland, UK and Sweden. Luxembourg, which was one of the best performers in 1996, turns out to be the worst performer in 2004 and 2005. We can also observe an increase in the inflation rate of both high and low inflation countries between 1999 and 2001. This is a reflection that, during the early years of the new monetary regime the Euro area was affected by a variety of price shocks such as the tripling of oil prices between early 1999

[13] The Euro area HCPI is computed by the Eurostat as the weighted average of the national HCPI's. The country weights are derived from national accounts data for 'household final monetary consumption expenditure' converted into purchasing power standards.

and mid-2000, the depreciation of the common currency over this period, and, in 2001, significant increases in food prices, due to a series of livestock epidemics.

The degree of misalignment between inflation and the policy reference value is given by:

$$e_t = \pi_t - \pi_t * \qquad (1)$$

where π_t^* is the policy reference value and π_t is the actual inflation, both as previously defined.

[INSERT FIGURE 1 HERE]

Figure 1 plots e_t for our sample countries; positive (negative) values indicate that inflation is higher (lower) from the target. In May 1998, when the European council decided about EMU membership, only Greece exhibited a significant positive inflation misalignment (2.68%). However, by June 2000, Greek inflation had been sharply reduced (e_t = -0.64%) which, in conjunction with satisfaction of the other criteria, allowed Greece to join the EMU in January 2001. During 1996-1998, inflation misalignments in Austria, Belgium, Finland, France, Germany and Luxembourg were always negative, while Ireland, Italy, Portugal and Spain exhibited both positive and negative misalignments. During 1999-2005, relatively large positive misalignments are observed in Greece, Ireland, Netherlands, Luxembourg, Portugal and Spain. Overall, inflation deviations appear to be characterised by cyclical behaviour. The horizon of the cycles is relatively long since they seem to last for at least two years with a sharp correction occurring at the peak of each cycle. This pattern is consistent with our proposed hypothesis of higher speed of adjustment of inflation towards the policy reference value, the greater the deviation from it.

3. Linear unit root tests

3.1 ADF unit root test

The standard linear ADF test uses the following regression model to test the stationarity of inflation deviations from the policy reference value:

$$\Delta e_t = \gamma_0 + \gamma e_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta e_{t-i} + \varepsilon_t \qquad (2)$$

where e_t is defined in (1), the $\gamma's$ are constants and ε_t is a random disturbance term: $\{\varepsilon_t\} \sim iid(0, \sigma_\varepsilon^2)$. The terms in Δe_{t-i} are included to remove any serial correlation in ε_t. Rejecting the null of unit root requires the estimates of γ to be negative and significantly different from zero.

The linear ADF results can be seen in Table 2. We witness over the full sample (1996-2005), that the null-unit root hypothesis is accepted for all countries with the exception of Greece and Italy, in the presence of a constant only in the ADF specification. When, in addition to the constant we incorporate a linear trend in the ADF specification there is evidence of unit roots in all sample countries apart from Italy. Overall, the linear ADF tests provide strong evidence of unit root behavior in the deviations of EMU countries inflation rates from the policy reference value[14].

For robustness, we compute unit root tests in two subsamples: 1996-1998 and 1999-2005. This allows us to investigate whether the introduction of the euro and common monetary policy on 1/1/1999 affects the time series structure of inflation deviations from the policy reference value. The results, presented in Table 2, provide further evidence of unit roots in inflation

[14] In addition, we considered the case of no intercept and trend. The results (available upon request) are very similar to the reported cases (of constant, constant and trend), in that, at the 5% level of significance the null of unit root is rejected only in two out of twelve countries (Greece and Italy).This finding, combined with the fact that the constant term is always insignificant (in regressions with intercept only) indicates that the non-rejection of the null-unit root is not due to the low power that may result from including a redundant intercept term in the unit root tests.

misalignments. Only in Italy (1996-1998 subperiod) and Belgium (1999-2005 subperiod), the null of unit root is rejected. Thus, the linear ADF evidence indicates that persistence in inflation misalignments from the policy reference value is a robust phenomenon extending over the two subsamples, pre and post the euro. This is a rather puzzling empirical finding that could reflect either that the Euro-area is not resembling OCA, or inadequate unit root testing procedures.

[INSERT TABLE 2 HERE]

3.2 Ng Perron unit root test

The Ng and Perron (2001) MZ_α test modifies the Phillips (1987) and Phillips and Perron (1988) Z_α test in a number of ways in order to increase the test's size and power. This testing procedure ensures that non-rejections of the null-unit root are not due to a low probability of rejecting a false null hypothesis, while rejections are not related to size distortions. The test statistic is defined as[15]:

$$MZ_a = \left[T^{-1}{e_T}^2 - s_{AR}^2\right]\left[2T^{-2}\sum\nolimits_{t=1}^{T}{e_{t-1}}^2\right]^{-1} \tag{3}$$

where t = 1…T , $s_{AR}^2 = \hat{\sigma}_k^2/[1-\hat{\gamma}(1)]^2$ is an autoregressive estimate of the spectral density at frequency zero of $\upsilon_t = \theta(L)\varepsilon_t = \sum\nolimits_{j=0}^{\infty}\theta_j\varepsilon_{t-j}$ with $\sum\nolimits_{j=0}^{\infty} j\left|\theta_j\right| < \infty$; $\hat{\gamma}(1) = \sum\nolimits_{i=1}^{k}\hat{\gamma}_i$ and $\hat{\sigma}_k^2 = (T-k)^{-1}\sum\nolimits_{t=k+1}^{T}\hat{\varepsilon}_t^2$ are calculated using the OLS estimates from Eq. (2). Following Elliot *et al*. (1996), Ng and Perron (2001) employ the local-to-unity GLS detrending procedure in order to

[15] The test statistic corresponds to the case where the variable into consideration (e_t) contains no deterministic term. If we allow for a constant, or constant and trend, then e_{t-1} and e_T in Eq. (3) should be replaced by their detrended counterparts.

benefit from the increased power offered by GLS detrending[16]. They also suggest that the autoregressive truncation lag, k, should be chosen using the Modified Akaike Information Criterion (MAIC) in an effort to avoid size distortions while maintaining power. The MAIC is calculated as follows:

$$MAIC(k) = \ln(\hat{\sigma}_k^2) + \frac{2[\tau_T(k) + k]}{T - k_{\max}} \tag{4}$$

where $\tau_T(k) = \left(\hat{\sigma}_k^2\right)^{-1} \hat{\gamma}^2 \sum_{t=k_{\max}+1}^{T} \left(e_{t-1}^d\right)^2$, k_{max} is the maximum value of k considered[17], e_t^d is the GLS detrended e_t and $\hat{\sigma}_k^2$ is defined as before using $k = k_{max}$.

The Ng Perron linear unit root test results are presented in Table 2. Using the full sample, the null-unit root is rejected only in Luxembourg (constant and trend). For all the other countries the results suggest that inflation deviations from the policy reference value are non-stationary. Using subsample regressions, we find that the null of unit root is never rejected. Hence, it appears that the non-rejection of the unit root null hypothesis is not the consequence of a low probability of rejecting a false null hypothesis. The Ng Perron results further support the findings of the ADF results suggesting that inflation differentials from the policy reference value follow a non-stationary process.

4. Non-linear modeling

Failure to reject non-stationarity may be the result of lack of power of linear unit root tests if the true adjustment process of inflation differentials is non-linear. Hence, in the next subsection

[16] For any series $\{e_t\}_{t=0}^{T}$ define $(e_0^{\bar{\alpha}}, e_t^{\bar{\alpha}}) = (e_0, (1-\bar{\alpha}L)e_t)$ for t = 1…T, and some chosen $\bar{\alpha} = 1 + \bar{c}/T$. The GLS detrended series is defined as: $e_t^d = e_t - \hat{\psi}' z_t$, where $\hat{\psi}$ minimizes $S(\bar{\alpha}, \psi) = (e^{\bar{\alpha}} - \psi' z^{\bar{\alpha}})'(e^{\bar{\alpha}} - \psi' z^{\bar{\alpha}})$, and z_t denotes the set of deterministic components of e_t. Elliot *et al*. (1996) suggest to set the value of $\bar{c}$ at -7 in the case of constant only, and -13.5 in the case of constant and linear trend.

[17] The upper bound is calculated as $k_{\max} = \text{int}(12/(T/100)^{1/4})$, where int(x) denotes an integer part of x. See Hayashi (2000, p.594) for a discussion of the selection of this upper bound.

we test for the presence of non-linearities. If non-linearities are detected we determine the appropriate non-linear specification. Finally, we perform the non-linear unit root tests.

4.1 Tests for linearity and STAR model selection

Consider two possible regimes comprising a pure 'small' and pure 'large' adjustment of inflation deviations from the policy reference value. Following Granger and Terasvirta (1993) and Terasvirta (1994) we write a Smooth Transition Autoregressive (STAR) model of order k, for e_t :

$$e_t = \theta_0 + \theta_1 ' x_t + \left(\delta_0 + \delta_1 ' x_t\right) F\left(e_{t-d}\right) + w_t \tag{5}$$

where $x_t = \left(e_{t-1}, e_{t-2}, \dots e_{t-k}\right)$, $\theta_1 = \left(\theta_1, \theta_2, \dots \theta_k\right)'$, $\delta_1 = \left(\delta_1, \delta_2, \dots \delta_k\right)'$, $w_t \sim iid\left(0, \sigma^2\right)$, $F(\bullet)$ is the continuous transition function, e_{t-d} is the switching variable, and d is the delay parameter. $F(\bullet)$ is a monotonically increasing function with $F(-) = 0$ and $F(\bullet) = 1$ which yields a non-linear asymmetric adjustment.

Consider the following Logistic STAR (LSTAR) function:

$$F\left(e_{t-d}\right) = \left\{1 + \exp\left[-a\left(e_{t-d} - c\right)\right]\right\}^{-1} \tag{6}$$

where a measures the smoothness of transition from one regime to another and c is some threshold value for e that indicates the halfway point between the two regimes.

The LSTAR model assumes that different regimes may have different dynamics and that adjustment takes place in every period but the smoothness of adjustment varies with the extent of the deviation from equilibrium. The transition function of LSTAR is monotonically increasing in

e_{t-d} and yields asymmetric adjustment towards equilibrium in the model. Moreover, $F(\bullet) \to 0$ as $e_{t-d} \to -\propto$ and $F(\bullet) \to 1$ as $e_{t-d} \to +\propto$ thus $F(\bullet)$ is bounded between 0 and 1 where $F(\bullet) = 0.5$ if $e_{t-d} = c$. The smaller is a, the smoother the transition. In the extreme, $a = 0$ means that $F(\bullet)$ becomes a constant and so (5) becomes a linear model. On the other hand, as $a \to \propto$ there is an even sharper transition at $e_{t-d} = c$ where $F(\bullet)$ jumps from 0 to 1.

Terasvirta and Anderson (1992) define the Exponential STAR (ESTAR) function as:

$$F(e_{t-d}) = 1 - \exp\left[-a(e_{t-d} - c)^2\right] \tag{7}$$

where, as previously, a measures the speed of transition from one regime to another and c is some threshold value for e which indicates the halfway point between the two regimes. The ESTAR function in (7) defines a transition function about c where $F(\bullet)$ is still bounded between 0 and 1.

The initial testing for the presence of non-linearities in e_t is based on three stages. First, a linear autoregressive model for e is specified in order to determine the lag length k. The lag length selection is based on the Schwarz information criteria and the Ljung-Box statistic for serial correlation. The residuals are saved from the chosen autoregressive model and denoted as v. Second, having determined k, the next stage is to test for the presence of non-linearities. This is done through the estimation of

$$v_t = \beta_0 + \beta_1 ' x_t + \beta_2 ' x_t e_{t-d} + \beta_3 ' x_t e^2_{t-d} + \beta_4 ' x_t e^3_{t-d} + w_t \tag{8}$$

where the linearity test is on the null hypothesis $H_0 : \beta_2' = \beta_3' = \beta_4' = 0.$ Equation (8) is estimated across a range of values for d where the smallest p-value attached to the linearity test determines d in the estimation of (5). The final stage of the non-linearity test is to determine which smooth transition model – LSTAR or ESTAR – is appropriate for the data. This is done by running the following sequence of nested tests.

$$H_{04} : \beta_4' = 0 \tag{9}$$

$$H_{03} : \beta_3' = 0 / \beta_4' = 0 \tag{10}$$

$$H_{02} : \beta_2' = 0 / \beta_4' = \beta_3' = 0 \tag{11}$$

Rejection of (9) implies selecting the LSTAR model. If we accept (9) and (10) we choose the ESTAR model. Accepting (9) and (10) and rejecting (11) lead to an LSTAR model. However, Granger and Terasvirta (1993) and Terasvirta (1994) show that application of this sequence of tests may lead to incorrect conclusions, because the higher order terms of the Taylor expansion used in deriving these tests are disregarded[18]. They therefore recommend that we should compute the p-values of all the F tests of (9)-(11) and make the choice of STAR model on the basis of the lowest p-value.

Table 3 displays the tests for non-linearity for e_t for the countries in our sample. The Ljung-Box statistic suggests white noise residuals for all autoregressive models. Following the selection of the lag length k for each autoregressive process, the delay parameter d is constrained to be $1 \leq d \leq 8.$ Using 0.05 as a threshold P-value, in all twelve countries apart from

[18] For more details see Terasvirta (1994) pages 211-212.

Greece and Luxembourg the test rejects linearity, classifying the series as non-linear. We can therefore proceed to build non-linear models for e_t in the ten remaining sample countries where the null of linearity is rejected. The tests for the choice between LSTAR and ESTAR models are shown in Table 4. Using the hypothesis tests outlined in equations (9)-(11), the results indicate that the ESTAR model is the most appropriate non-linear model in all cases.

[INSERT TABLES 3 and 4 HERE]

4.2 Non-linear unit root test

The ESTAR model assumes that the adjustment of inflation towards the policy reference value is characterized by a symmetric non-linear process[19]:

$$e_t = \beta e_{t-1} + \delta e_{t-1}\left(1 - \exp[-\alpha e_{t-1}{}^2]\right) + u_t \qquad (12)$$

where u_t is the error term and the other variables are as previously defined. Under the null-non stationarity, $\beta = 1$ and $a = 0$, inflation follows a random walk around π_t^*. In the case of stationarity ($a > 0$), inflation reverses to π_t^*. Computing a first-order Taylor series approximation to (12) under the null and allowing for serial correlation in u_t, we obtain the following auxiliary regression model (Kapetanios *et al*., 2003):

$$\Delta e_t = \gamma_0 + \gamma e_{t-1}{}^3 + \sum_{i=1}^{k} \gamma_i \Delta e_{t-i} + v_t \qquad (13)$$

where v_t is the error term and the other variables are defined as previously. The null hypothesis of equation (13) is that $\gamma = 0$. Equation (13) does not provide a direct method to test the

[19] See, among others, Granger and Terasvirta (1993) for other applications of the ESTAR model.

statistical significance of γ.[20] This is because the cubic term embedded in γ is a non-linear function of the underlying parameter estimate resulting in the distribution of γ being unknown. Therefore, we use a bootstrap technique to obtain an asymptotic *t* statistic to test the significance of γ .The model that represents the null is

$$\Delta e_t = \gamma_0 + \sum_{i=1}^{k} \gamma_i \Delta e_{t-i} + v_t \tag{13a}$$

The model (13a) is a fully specified parametric model, which means that each set of parameter values for γ_0 and γ_1 defines just one data generating process (dgp). The first step in constructing a bootstrap dgp is to estimate (13a) by OLS, yielding the restricted estimates $\bar{\gamma}_0$, $\bar{\gamma}_1$. Then the boostrap dgp is given by

$$\Delta e_t{*} = \gamma_0{*} + \sum_{i=1}^{k} \gamma_i \Delta e^*_{t-i} + v_t{*} \tag{14}$$

which is just the element of the model (13a) characterized by the parametric estimates under the null, with stars to indicate that the data are simulated. By computing 10000 bootstrapped resamples of $v_t{*}$ for each of our sample countries we obtain 95% confidence intervals to test the null hypothesis of $\gamma = 0$ in equation (13). The idea in 10000 replications is to determine the appropriate critical values for the t test under the null hypothesis of $\gamma = 0$. In our empirical estimates we report the p-values obtained through the simulation exercise for the estimated *t* values.

The non-linear unit root test results are presented in Table 5. The Jarque-Bera normality test indicates that the residuals are normally distributed in all cases. Thus, the non linearities in

[20] Critical values are provided by Kapetanios et al (2003) for an asymptotic distribution. These critical values are only valid for large samples.

inflation misalignments with respect to the policy reference value are not the outcome of any outliers in the data. The non-linear ADF tests show that inflation deviations follow a stationary process at all levels of significance. The decisive rejection of the null-unit root appears to be the result of the significant increase in the magnitude of the estimated ADF coefficient, γ. This finding holds across all ten countries and is unaffected by the inclusion of a linear trend in the regressions. Hence, the evidence of non-stationarity obtained from the linear unit root tests disappears when we allow for non-linear adjustment in inflation deviations.

[INSERT TABLE 5 HERE]

5. Conclusions

This paper investigates the time series properties of inflation adjustment to the target, given by the Maastricht reference value and the ECB target, within a framework of twelve EU countries that adopted the euro and common monetary policy. We find that the existence of non-linearities in inflation deviations from the policy reference value seriously affects the inference results from unit root tests. In particular, using linear unit root tests we discover that in ten out twelve countries, inflation deviations follow a non-stationary process. We show that this result is robust to subsample division, and to alternative specifications of linear unit root tests.

A possible explanation for these findings could be that the true rate of adjustment of inflation towards the target is increasing in the deviation from the target value, as opposed to being constant, which is the case in all linear unit root tests. Consequently, we test for non-linear behavior in inflation misalignments and reject linearity in ten out of twelve nations. For the ten remaining countries, we undertake econometric tests that lead us to the conclusion that the ESTAR model is the appropriate non-linear specification. Upon application of the ESTAR unit root test to the data, we discover that inflation adjustments relative to the target follow a stationary process for all the countries that exhibit non-linear behavior. The implications of our empirical findings are that inflation misalignments are not explosive and persistent, suggesting that they do not exacerbate real divergence. Therefore, regional inflation dispersion has not posed a threat to the 'one size fits all' policy of the ECB. Given the importance of inflation adjustment of EMU countries and the low power of the standard linear unit root tests, the empirical findings in this paper should not be ignored.

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Table 1: Average annual inflation rate (%) of three best and worst performing EU countries.

Year	Three Best Performers			Three Worst Performers		
1996	Sweden 0.78	Finland 1.06	Luxembourg 1.16	Greece 7.57	Italy 3.9	Spain 3.5
1997	Austria 1.15	Finland 1.21	Ireland 1.23	Greece 5.4	Italy 1.88	Portugal 1.87
1998	Germany 0.59	France 0.66	Austria 0.82	Greece 4.42	Portugal 2.18	Ireland 2.12
1999	Austria 0.51	Sweden 0.54	France 0.56	Ireland 2.42	Spain 2.21	Portugal 2.14
2000	UK 0.78	Germany 1.38	France 1.81	Ireland 5.13	Luxembourg 3.7	Spain 3.41
2001	UK 1.23	France 1.76	Germany 1.88	Netherlands 4.98	Portugal 4.31	Ireland 3.9
2002	UK 1.24	Germany 1.34	Belgium 1.54	Ireland 4.61	Greece 3.84	Netherlands 3.79
2003	Germany 1.03	Austria 1.28	Finland 1.29	Ireland 3.92	Greece 3.38	Portugal 3.21
2004	Finland 0.14	Denmark 0.88	Sweden 1.01	Luxembourg 3.17	Spain 3	Greece 2.98
2005	Finland 0.76	Sweden 0.81	Denmark 1.68	Luxembourg 3.68	Greece 3.42	Spain 3.32

Note: The inflation rate was calculated as the twelfth difference of the monthly HCPI.

Table 2: Linear Unit root test results

Country	Linear ADF t-test statistic						Ng Perron MZ_α test statistic					
	Constant			Constant and Trend			Constant			Constant and Trend		
	1996-2005	1996-1998	1999-2005	1996-2005	1996-1998	1999-2005	1996-2005	1996-1998	1999-2005	1996-2005	1996-1998	1999-2005
Austria	-1.53 [12]	-1.83 [1]	-2.23 [1]	-2.25 [1]	-3.26 [0]	-1.86 [1]	-6 [12]	-5.94 [1]	-1.65 [1]	-13.73 [12]	-13.01 [0]	-4.43 [1]
Belgium	-0.98 [5]	-1.87 [0]	-1.71 [5]	-1.55 [5]	-2.63 [0]	-1.63 [5]	-1.13[5]	-6.56 [0]	-3.38 [6]	-4.76 [5]	-8.98 [0]	-6.41 [6]
Finland	-1.96[12]	-2.35 [0]	-1.28 [2]	-1.91 [12]	-2.13 [0]	-2.31 [2]	-3.01 [12]	-4.27 [0]	-1.91 [2]	-6.95 [12]	-7.84 [0]	-2.64 [2]
France	-1.27 [12]	-1.09 [0]	-2.47 [4]	-2.43 [12]	-2.2 [0]	-1.08 [7]	-4.86 [12]	-1.87 [0]	-1.82 [11]	-8.71 [12]	-7.58 [0]	-3.04 [2]
Germany	-1.03 [12]	-1.25 [0]	-2.13 [4]	-2.29 [4]	-0.23 [9]	-2.74 [1]	-2.55 [12]	-5.54 [0]	-2.41 [1]	-11.7 [4]	-8.68 [0]	-9.54 [1]
Greece	-3.13 [3] *	-0.58 [0]	-1.72 [2]	-2.57 [3]	-1.49 [0]	-2.31 [2]	-0.78 [3]	-0.43 [0]	-5.25 [2]	-3.79 [3]	-4.08 [0]	-6.45 [2]
Italy	-4.2 [0]**	-4.3 [0]**	-1.97 [4]	-4.5 [0]***	-2.24 [0]	-1.54 [4]	-0.19 [12]	-2.6 [3]	-1.71 [3]	-2.02 [0]	-2.27 [1]	-2.48 [4]
Ireland	-1.38 [12]	-1.26 [0]	-2.05 [4]	-0.6 [12]	-1.41 [0]	-2.03 [2]	-2.19 [12]	-3.04 [0]	-4.52 [3]	-2.97 [12]	-3.24 [0]	-6.28 [3]
Luxembourg	-1.41 [12]	-1.55 [1]	-1.46 [2]	-2.53 [12]	-1.06 [1]	-2.04 [4]	-1.69 [12]	-5 [1]	-0.51 [7]	-144 [12]**	-7.41 [0]	-2.6 [4]
Netherlands	-1.65 [3]	-2.14 [0]	-1.57 [3]	-1.43 [0]	-2.29 [0]	-1.88 [3]	-4.34 [12]	-5.28 [0]	-4.28 [3]	-5.21 [3]	-8.99 [0]	-4.78 [3]
Portugal	-1.26 [12]	-1.12 [0]	-1.74 [0]	-1.12 [12]	-0.95 [0]	-1.79 [0]	-2.17 [12]	-3.02 [0]	-5.24 [0]	-3.61 [12]	-2.97 [0]	-5.56 [0]
Spain	-1.2 [12]	-1.33 [0]	-3.04 [3] *	-2.66 [0]	-1.18 [0]	-2.96 [3]	-4.41 [12]	-1.54 [1]	-1.18 [3]	-6.2 [12]	-2.67 [0]	-6.55 [3]

Note: The number in the bracket shows the number of lagged difference terms in the ADF and Ng-Perron linear unit root tests. It was chosen by the Modified Akaike Criterion. The reported *t*-statistics and MZ_α statistics test the null hypothesis that inflation contains a unit root using equations (2) and (3), respectively. **, * indicate rejection of the null-unit root hypothesis at the 1, 5% level of significance.

Table 3: Tests for Non-Linearities

Country	**k**	**d**	**p-value**	**Q(1)**
Austria	3	1	0.04	0.312
Belgium	3	1	0.03	0.354
Finland	2	1	0.04	0.369
France	2	1	0.03	0.387
Germany	2	1	0.03	0.314
Greece	4	1	0.13	0.311
Italy	3	1	0.04	0.312
Ireland	2	1	0.04	0.311
Luxembourg	2	1	0.11	0.344
Netherlands	3	1	0.04	0.376
Portugal	3	1	0.03	0.379
Spain	3	1	0.02	0.316

Note: Table 3 reports the linearity tests of e_t over the time period 1996-2005. The null of non-linearity is based on equation (8). The column headed 'p-value' corresponds to the test H_0 where the null is linearity. It should be noted that the Schwartz criterion is used to determine the lag length *k* of the autoregressive process. The residuals from the autoregressive processes were then saved. Having determined k, a range of delay parameters ($d \leq 1 \leq 8$) were employed. The value of d chosen is that which gives rise to the lowest p-value of the linearity test using the data for the residuals of the autoregressive process. The linearity test is a variable-deletion F test on the restriction applied to equation (8). The column headed Q(1) refers to the p-value associated with the Ljung-Box Q statistic for serial correlation among the residuals.

Table 4: Specification of the Non-Linear Model

Country	H_{04}	H_{03}	H_{02}	**Type of Model**
Austria	0.132	0.02#	0.139	ESTAR
Belgium	0.139	0.03#	0.146	ESTAR
Finland	0.144	0.04#	0.149	ESTAR
France	0.146	0.05#	0.151	ESTAR
Germany	0.149	0.07#	0.155	ESTAR
Italy	0.151	0.06#	0.163	ESTAR
Ireland	0.138	0.05#	0.148	ESTAR
Netherlands	0.135	0.04#	0.139	ESTAR
Portugal	0.134	0.03#	0.140	ESTAR
Spain	0.129	0.02#	0.135	ESTAR

Note: Table 4 reports the variable deletion tests portrayed in equations (9), (10) and (11) over the time period 1996-2005. # denotes the lowest p-value associated with the variable-deletion tests and therefore the determination of the relevant STAR model. The values of k and d are reported in Table 3.

Table 5: Non-Linear Unit root test results

Country	Non-Linear ADF t-test statistic			
	Constant	Constant and Trend	NORM (2) Constant	NORM (2) Constant and Trend
Austria	-4.76** [0] (0.005)	-4.99**[0] (0.005)	0.144	0.149
Belgium	-4.34** [12] (0.004)	-4.39** [12] (0.004)	0.147	0.153
Finland	-4.22** [12] (0.005)	-4.25** [12] (0.005)	0.141	0.142
France	-4.33** [12] (0.005)	-4.34** [12] (0.005)	0.143	0.146
Germany	-4.62** [12] (0.005)	-4.65** [1] (0.005)	0.146	0.148
Italy	-4.65** [12] (0.005)	-4.69** [12] (0.005)	0.148	0.150
Ireland	-4.24** [13] (0.004)	-4.27** [13] (0.004)	0.149	0.151
Netherlands	-4.27** [13] (0.004)	-4.28** [13] (0.004)	0.151	0.154
Portugal	-4.30** [12] (0.004)	-4.32** [12] (0.004)	0.152	0.155
Spain	-4.24** [12] (0.004)	-4.26** [12] (0.005)	0.147	0.150

Note: The non-linear unit root tests were undertaken over the time period 1996-2005. The number in the bracket shows the number of lagged difference terms in the regressions. It was chosen by the Modified Akaike Criterion. The reported *t*-statistics test the null hypothesis that inflation contains a unit root using equation (13). **, * indicate rejection of the null-unit root hypothesis at 1, 5% level of significance. NORM(2) Constant is the P-value of the Normality test with a constant and NORM(2) Constant and Trend is the P-value of the Normality test with a constant and a trend. Figures in the round brackets represent the p-value of the t statistic obtained through bootstrap simulation.

Figure 1: Inflation deviations from the policy reference value, 1996-2005.