The background of the entire page is a grayscale photograph of a modern university building. The building features a prominent glass facade and a courtyard area in the foreground with trees and benches. The text is overlaid on this image.

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The Nature of Persistence in Euro Area Inflation:
A Reconsideration

by

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The Nature of Persistence in Euro Area Inflation: A Reconsideration*

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Abstract

Recent empirical studies find little evidence of a change in euro area inflation persistence over the post-1970 period. Their methodology is, however, primarily based on standard unit root and structural break tests that are not designed to detect a change in persistence when the process shifts from stationarity to non-stationarity or vice-versa. This paper employs five classes of tests for a change in persistence that allow for such shifts as well as consistent break date estimation methods to argue that euro area inflation shifted from a unit root process to a stationary one around the time the Maastricht Treaty came into effect with an explicit mandate for price stability as the primary objective of monetary policy. The evidence presented is consistent with the view that the degree of inflation persistence varies with the transparency and credibility of the monetary regime.

Keywords: persistence, price stability, unit root, monetary policy

JEL Classification: C22, E3, E5

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1 Introduction

In the last decade and a half or so, the issue of the nature of inflation persistence and its relation to monetary policymaking has been a subject of intense debate among economists. Given the recent adoption of inflation targeting as the primary objective of long-run monetary policy in many countries, a question that has been receiving an increasing amount of attention is whether inflation persistence is a deep inherent characteristic of the economy and thus invariant to policy shifts. Evidence in favor of high and unchanged persistence across different policy regimes is construed as suggesting that inflation is structural in the sense of Lucas (1976) and should therefore be a feature that a reasonable model for the economy should be able to replicate. This has motivated the development of general equilibrium macroeconomic models which have explicitly incorporated inflation persistence into the structure of the economy. These models could then be used for the computation of optimal policies as well as evaluating the effectiveness of alternative monetary policy regimes. The idea is that by uncovering the deep structural parameters that characterize the economy as well as explicitly modeling expectations, one can hope to capture the dependence of agents' behavior on the functions describing policy.

Several approaches have been adopted in developing the microeconomic foundations for structural inflation persistence. Some authors assume that the persistent behavior of inflation results from the structure of nominal contracts (Fuhrer and Moore, 1995, Fuhrer, 2000, Calvo et al., 2001 and Christiano et al., 2005). An alternative approach assumes that private agents face information-processing constraints (Roberts, 1998, Mankiw and Reis, 2002, and Woodford, 2003). Yet another mechanism has been to simply generate inflation persistence through the exogenous structural shocks affecting the economy (Rotemberg and Woodford, 1997 and Ireland, 2004).

If, however, inflation persistence is not an inherent characteristic of the economy but instead depends on shifts in monetary policy, these backward looking models are no longer structural and therefore their use in policy evaluation is likely to provide misleading implications. This is indeed the essence of the Lucas Critique according to which the parameters of macroeconomic models depend implicitly on agents' expectations of the policy process and are unlikely to remain stable as policymakers change their behavior. Given that a strict mandate for price stability was adopted for the euro area following the signing of the Maastricht Treaty in 1992, it seems useful to assess the empirical importance of this policy shift with respect to the dynamics of euro area inflation. Such an analysis is likely to provide a

guide as to whether a backward looking component in the inflation process should be introduced as a structural feature in rational expectations models for the euro area or a purely forward looking model provides a better approximation to the observed behavior of inflation.

In a recent article, O'Reilly and Whelan (2005) find little evidence of a change in euro area inflation persistence over the post-1970 period. Full sample estimates of the persistence parameter are generally close to unity and they fail to reject the hypothesis that this parameter has been stable over time. They interpret their results as favoring the simple backward-looking rule-of-thumb model of expectations over models that feature forward looking rational expectations. Their methodology is based on standard unit root and structural break tests on the persistence parameter in an autoregressive specification for the inflation process. These procedures are, however, not designed to detect a change in persistence when a sub-sample of the data has a unit root, i.e., when the process shifts from stationarity to non-stationarity or vice-versa. In particular, unit root tests have poor power in detecting processes which exhibit stationary behavior in a certain part of the sample and are non-stationary otherwise. The reason is that the unit root component of such processes dominate the stationary component so that the tests are not consistent. Further, the usual structural break tests (e.g., Andrews, 1993) are based on the presumption that the variables are stationary in all regimes and therefore preclude the possibility of a unit root in a particular regime.

In this paper, we employ five classes of tests for a change in persistence that allow for such non-stationary alternatives to argue that euro area inflation shifted from a unit root process to a stationary one at some point in the sample. Statistical methods to select the break date identify the change in the second quarter of 1993 around the time the Maastricht Treaty came into effect, thereby suggesting a role for inflation targeting. Bias-adjusted estimates of the persistence parameter, half life estimates and confidence intervals for the largest autoregressive root all suggest a marked decline in persistence after the break. We further illustrate that the hypothesis of stationarity with a mean shift but a stable persistence parameter is not compatible with the data. This contrasts with the multi-country evidence in Cecchetti and Debelle (2006) and Levin and Piger (2003), who argue that inflation persistence is stable once one allows for a structural break in the mean of the series. The evidence presented is therefore consistent with the view that the nature of inflation persistence varies with the transparency and credibility of the monetary regime.

Our findings can be viewed as complementary to recent evidence provided by Benati (2008). He shows, based on comparing estimates of the persistence parameter before and

after the introduction of the European Monetary Union (EMU), that there has been a significant decline in euro inflation persistence so that the latter is not a characteristic that should be built into the structure of macroeconomic models for the euro area. We confirm his results by employing persistence change tests that account for the possibility of a unit root in an unknown sub-sample of the data as well as econometric methods for endogenous selection of the break date instead of exogenously imposing such a date.

Our results are also consistent with the theoretical model developed in Erceg and Levin (2003). These authors formulate a dynamic general equilibrium model with staggered nominal contracts in which private agents have limited information about the central bank's objectives. In particular, households and firms use optimal filtering to disentangle persistent shifts in the inflation target from transitory disturbances to the monetary policy rule. Under these assumptions, the speed at which private agents recognize a new inflation target depends on the transparency and credibility of the central bank. The model exhibits moderate persistence when monetary policy is transparent and credible, and much higher persistence when agents must use signal extraction to make inferences about the central bank's inflation target.

The paper is organized as follows. Section 2 provides a brief overview of the existing literature on inflation persistence in the euro area. In Section 3, we discuss tests for a change in persistence that form the basis of our empirical analysis. Section 4 presents the empirical results. Section 5 contains a discussion of the results and some concluding remarks are given in Section 6.

2 Euro Area Inflation Persistence

There is a vast and growing literature studying the nature and characteristics of inflation dynamics in the euro area. The European Central Bank (ECB) and the National Central banks comprising the Eurosystem embarked on a comprehensive research effort in the form of the Inflation Persistence Network (IPN) which investigated both the characteristics of inflation persistence and the pattern of determinants of price setting in the euro area and its member countries.¹ The network addressed the patterns, causes and policy implications of inflation persistence based on data from individual consumer and producer prices, surveys on firms' price-setting practices, aggregated sectoral, national and area-wide price indices. In what follows, we will not attempt an exhaustive review of the literature, but instead focus on

¹More information about the general purpose, organization and publication of the IPN can be found on their website http://www.ecb.int/home/html/researcher_ipn.en.html

the main findings of the network as well as related work that bear relevance to the current study. Moreover, we shall primarily confine ourselves to the macro evidence on inflation persistence given the main theme of the paper and postpone a discussion of the relationship between the micro and macro evidence to Section 6. An overview of the main findings of the IPN is presented in two companion papers by Altissimo et al. (2006) and Alvarez et al. (2005), with the former discussing the macro evidence on the degree of inflation persistence while the latter reviews the micro evidence on price setting practices.

Analysis of aggregate data for the euro area typically yields very high estimates of inflation persistence for sample periods spanning multiple decades. For instance, Batini (2002) presents evidence based on analyzing the autocorrelation properties of inflation as well as the lag in the response of inflation to monetary policy shocks to argue that the persistence of euro area inflation seems to have varied only marginally over 1970-2002, despite substantive shifts in the monetary policy regime after the collapse of the Bretton Woods' exchange rate system. O'Reilly and Whelan (2005) apply standard unit root and structural break tests on the persistence parameter to conclude in favor of little change in euro area inflation persistence over the period 1970-2002. They cannot reject the hypothesis that this parameter has been stable over time and find estimates of the persistence parameter that are very close to unity. They construe this result as providing support for purely backward looking structural macroeconomic models or hybrid models with only a weak forward looking element (see also Rudebusch, 2005). These studies thus provide empirical support for the view that inflation persistence is a structural parameter and hence invariant to changes in the policy regime.

According to an alternative view, the estimated degree of inflation persistence falls substantially once we allow for time variation in the mean level of inflation, either by explicitly allowing for discrete breaks in the regression intercept or by focusing on shorter sample periods. The intuition is that ignoring occasional shifts in mean leads to spuriously high estimates of the persistence parameter (see Perron, 1990). Such breaks in the mean inflation rate have been found to coincide with observed shifts in monetary policy and are associated with breaks in the mean of nominal as opposed to real variables (Corvoiser and Mojon, 2005). Levin and Piger (2003) apply classical and Bayesian econometric procedures to study inflation dynamics for twelve industrial countries over the period 1984-2003, using four different price indices for each country. For many of these countries, they find a break in mean in the late 1980s or early 1990s, allowing for which reduces the extent of estimated persistence significantly, thereby leading them to infer that high inflation persistence is not an inherent characteristic of industrial economies. Similar conclusions are also reached in

analyses conducted by Gadzinski and Orlandi (2004) for 79 inflation series covering the EU countries, the euro area and the US, and Cecchetti and Debelle (2006) based on aggregate as well as disaggregate inflation data for 17 countries over various time periods.

More recently, Benati (2008) documents that inflation persistence in the euro area, among other regions, is not a deep structural feature of the economy that should be specifically incorporated in general equilibrium macroeconomic models. In particular, he shows that such persistence is indeed quite low in the regime following the introduction of the EMU. His results therefore suggest that evaluation of the pros and cons of alternative monetary policy regimes based on models featuring intrinsic inflation persistence is likely to deliver misleading conclusions. The results of the current paper can be treated as complementary to those in Benati (2008), being based on a different econometric methodology which explicitly allows us to test for a change in persistence in the potential presence of a unit root in a sub-sample of the data, as well as endogenously determine the date of the break as opposed to specifying it a priori.

3 Tests for a Change in Persistence

In this section, we will briefly discuss the four classes of tests for a change in persistence that we employ in our empirical analysis. Some of these tests have been developed for the null hypothesis of a unit root throughout the sample while others take a stable stationary process as the null. Further, the tests also differ according to the particular alternative hypothesis that they are designed to detect. Specifically, tests that are used to identify a potential shift from unit root to stationary behavior can be different from those designed to detect a change from stationarity to a unit root. Both types of tests are, however, useful in identifying the possible direction of the shift. In addition, we also present results of tests that do not presume a particular alternative hypothesis but are rather aimed at simply determining if the persistence parameter has been stable over the sample.

A process y_t that is stationary $[I(0)]$ for a fraction $\lambda_0 \in (0, 1)$ of the sample and subsequently shifts to a unit root process $[I(1)]$ for the rest of the sample can be represented as

$$y_t = \begin{cases} \beta_0 + z_{t,1} & \text{if } t = 1, \dots, [\lambda_0 T] \\ \beta_1 + z_{t,0} & \text{if } t = [\lambda_0 T] + 1, \dots, T \end{cases} \quad (1)$$

where $z_{t,0} = z_{t-1,0} + u_t$ and $z_{t,1}, u_t$ are stationary processes.

Similarly, a process that shifts from $I(1)$ to $I(0)$ at the point λ_0 is given by

$$y_t = \begin{cases} \beta_0 + z_{t,0} & \text{if } t = 1, \dots, [\lambda_0 T] \\ \beta_1 + z_{[\lambda_0 T],0} + z_{t,1} & \text{if } t = [\lambda_0 T] + 1, \dots, T \end{cases} \quad (2)$$

Note that the data generating processes (1) and (2) allow for a shift in level to occur simultaneously with a shift in persistence. We now briefly describe the test procedures that are designed to test the null hypothesis of stable persistence against the alternative hypotheses given by (1) and (2).

3.1 Ratio-based Tests

Kim (2000) and Busetti and Taylor (2004) (henceforth BT) consider testing the null hypothesis of stationarity throughout the sample versus $I(0)$ - $I(1)$ behavior based on the ratio statistic computed for some break fraction λ :

$$K_M(\lambda) = \frac{[(1-\lambda)T]^{-2} \sum_{t=[\lambda T]+1}^T \left(\sum_{i=[\lambda T]+1}^t \hat{\varepsilon}_{1,i} \right)^2}{[\lambda T]^{-2} \sum_{t=1}^{[\lambda T]} \left(\sum_{i=1}^t \hat{\varepsilon}_{0,i} \right)^2}$$

where $\hat{\varepsilon}_{0,t} = y_t - (1/[\lambda T]) \sum_{i=1}^{[\lambda T]} y_i$, for $t = 1, \dots, [\lambda T]$ and $\hat{\varepsilon}_{1,t} = y_t - \{1/[(1-\lambda)T]\} \sum_{i=[\lambda T]+1}^T y_i$ for $t = [\lambda T] + 1, \dots, T$. Since the breakpoint is unknown, the statistic $K_M(\lambda)$ is first computed for each value of $\lambda \in \Lambda$, where Λ is a given sub-interval of $[0, 1]$. In our empirical analysis, we set $\Lambda = [0.2T, 0.8T]$. The final test statistic is then obtained by taking an appropriate function of the resulting sequence of statistics $\{K_M(\lambda), \lambda \in \Lambda\}$. Three such functions are considered. First, after Andrews (1993), the maximum over the sequence of statistics, *viz.*,

$$H_1(K_M(\cdot)) = \max_{\lambda \in \Lambda} K_M(\lambda) \quad (3)$$

Second, Hansen's (1992) mean score statistic

$$H_2(K_M(\cdot)) = \int_{\lambda \in \Lambda} K_M(\lambda) d\lambda \quad (4)$$

Finally, after Andrews and Ploberger (1994), the mean-exponential statistic

$$H_3(K_M(\cdot)) = \log \left\{ \int_{\lambda \in \Lambda} \exp \left(\frac{1}{2} K_M(\lambda) \right) d\lambda \right\} \quad (5)$$

In each case, the null hypothesis is rejected for large values of the $H_j(K_M(\cdot))$ statistics, $j = 1, 2, 3$.

BT show that when the process is given by (1), tests that reject for large values of statistics based on the reciprocal of $K_M(\lambda)$ can provide consistent inference. BT suggest that the inconsistency of the ratio-based tests against the “wrong” alternative can be used constructively to help identify the direction of change. For instance, if the tests against the $I(1)$ - $I(0)$ alternative reject while those against the $I(0)$ - $I(1)$ alternative do not, this could be interpreted as evidence in favor of an $I(1)$ - $I(0)$ process. When the direction of change is unknown, they also propose using the statistic

$$\max H_j(K) = \max \{H_j(K_M(\cdot)), H_j((K_M(\cdot))^{-1})\}, \quad j = 1, 2, 3 \quad (6)$$

in each case rejecting for large values of the statistics.

Once these tests provide evidence against the null, it is then desirable to estimate the breakpoint. Based on the residual processes $\hat{\varepsilon}_1$ and $\hat{\varepsilon}_0$, BT and Kim et al. (2002) independently propose the following estimator for the breakpoint:

$$\hat{\lambda}_M = \operatorname{argmax}_{\lambda \in \Lambda} S_M(\lambda)$$

where

$$S_M(\lambda) = \left([(1 - \lambda)T]^{-2} \sum_{t=[\lambda T]+1}^T \hat{\varepsilon}_{1,t}^2 \right) \left([\lambda T]^{-2} \sum_{t=1}^{[\lambda T]} \hat{\varepsilon}_{0,t}^2 \right)^{-1}$$

The estimator is consistent for the true breakpoint regardless of the direction of shift, i.e., it is valid whether the true data generating process involves an $I(1)$ - $I(0)$ shift or an $I(0)$ - $I(1)$ shift.

3.2 Modified Ratio-based Tests

A limitation of the ratio-based tests is that they are unable to discern between a change in persistence and a constant $I(1)$ process. In particular, they tend to reject the null hypothesis of stability when the process has a unit root throughout the sample. To remedy this problem, Harvey et al. (2006) propose modified versions of the ratio-based tests that have the same asymptotic critical values regardless of whether the process is $I(0)$ or $I(1)$ throughout. Consequently, the null hypothesis for the modified tests is that of constant persistence (either

constant $I(1)$ or constant $I(0)$). The tests are defined as (for $j = 1, 2, 3$)

$$H_j^m(K_M(\cdot)) = \exp(-b \min[J_{\min}, J_{\min}^R]) H_j(K_M(\cdot)) \quad (7)$$

$$H_j^m((K_M(\cdot))^{-1}) = \exp(-b \min[J_{\min}, J_{\min}^R]) H_j((K_M(\cdot))^{-1}) \quad (8)$$

$$\max H_j^m(K) = \max \{H_j^m(K_M(\cdot)), H_j^m((K_M(\cdot))^{-1})\} \quad (9)$$

In (7) and (8), $J_{\min} = \min_{\lambda \in \Lambda} J_{1, [\lambda T]}$ and $J_{\min}^R = \min_{\lambda \in \Lambda} J_{[\lambda T], T}$, where J_{\min} is T^{-1} times the Wald statistic for testing the joint hypothesis $\gamma_{k+1} = \dots \gamma_9 = 0$ in the regression

$$y_t = c_1 + \sum_{i=k+1}^9 \gamma_i t^i + \text{error}, \quad t = 1, \dots, [\lambda T]$$

and $J_{[\lambda T], T}$ is T^{-1} times the Wald statistic for testing the joint hypothesis $\gamma_{k+1} = \dots \gamma_9 = 0$ in the regression

$$y_t = c_2 + \sum_{i=k+1}^9 \gamma_i t^i + \text{error}, \quad t = [\lambda T] + 1, \dots, T$$

The tests in (7)-(9) are shown to have adequate empirical size under the null while retaining decent power in finite samples against both $I(1)$ - $I(0)$ and $I(0)$ - $I(1)$ alternatives.

3.3 Locally Best Invariant (LBI) Tests

BT also propose testing the null of a stable stationary process using a locally best invariant (LBI) test for the $I(0)$ - $I(1)$ alternative:

$$L_1(\lambda) = \hat{\sigma}_L^{-2} (T - [T\lambda])^{-2} \sum_{t=[\lambda T]+1}^T \left(\sum_{j=t}^T \hat{\varepsilon}_j \right)^2$$

for some $\lambda \in \Lambda$, where $\hat{\varepsilon}_t = y_t - T^{-1} \sum_{t=1}^T y_t$ ($t = 1, \dots, T$) are the full sample OLS residuals and the estimate of the long-run variance is given by

$$\hat{\sigma}_L^2 = T^{-1} \sum_{t=1}^T \hat{\varepsilon}_t^2 + 2T^{-1} \sum_{s=1}^m [1 - s/(m+1)] \sum_{t=s+1}^T \hat{\varepsilon}_t \hat{\varepsilon}_{t-s}$$

where m is the lag-truncation parameter. We choose m using Andrews' (1993) data dependent method based on an AR(1) approximation for the residuals. The test for the $I(1)$ -

$I(0)$ alternative is given by

$$L_0(\lambda) = \hat{\sigma}_L^{-2}([T\lambda])^{-2} \sum_{t=1}^{[T\lambda]} \left(\sum_{k=t}^T \hat{\varepsilon}_k \right)^2$$

Here we again consider the functions $H_j(L_1(\cdot))$ and $H_j(L_0(\cdot))$, $j = 1, 2, 3$ defined in (3)-(5) applied to the sequences $\{L_1(\lambda), \lambda \in \Lambda\}$ and $\{L_0(\lambda), \lambda \in \Lambda\}$ respectively. Unlike the ratio-based tests, a LBI test directed against a particular alternative has substantial power when the other alternative is in fact the true data generating process. So we might expect a rejection by both tests if any one of the alternatives is true. For an unknown direction of change, BT propose a test statistic analogous to (6):

$$\max H_j(L) = \max \{H_j(L_1(\cdot)), H_j(L_0(\cdot))\}, \quad j = 1, 2, 3$$

Again, we reject the null for large values of the test statistics.

3.4 Sub-sample Stationarity Tests

Another class of tests suggested by BT is based on the application of the stationarity test proposed by Nyblom and Mäkeläinen (1983) to different sub-samples of the data. The test statistic for the $I(0)$ - $I(1)$ alternative is

$$NM_1(\lambda) = (T - [T\lambda])^{-2} [\hat{\sigma}_1(\lambda)]^{-2} \sum_{t=[\lambda T]+1}^T \left(\sum_{j=[\lambda T]+1}^t \hat{\varepsilon}_{1,j} \right)^2$$

while that for the $I(1)$ - $I(0)$ alternative is

$$NM_0(\lambda) = ([T\lambda])^{-2} [\hat{\sigma}_0(\lambda)]^{-2} \sum_{t=1}^{[\lambda T]} \left(\sum_{j=1}^t \hat{\varepsilon}_{0,j} \right)^2 \quad (10)$$

The residuals $\hat{\varepsilon}_1$ and $\hat{\varepsilon}_0$ are as defined for the ratio-based tests. The estimates $\hat{\sigma}_1^2(\lambda)$ and $\hat{\sigma}_0^2(\lambda)$ are given by

$$\begin{aligned}\hat{\sigma}_1^2(\lambda) &= (T - [\lambda T])^{-1} \sum_{t=[\lambda T]+1}^T \hat{\varepsilon}_{1,t}^2 + 2(T - [\lambda T])^{-1} \sum_{s=1}^m [1 - s/(m+1)] \sum_{t=s+[\lambda T]+1}^T \hat{\varepsilon}_{1,t} \hat{\varepsilon}_{1,t-s} \\ \hat{\sigma}_0^2(\lambda) &= ([\lambda T])^{-1} \sum_{t=1}^{[\lambda T]} \hat{\varepsilon}_{0,t}^2 + 2([\lambda T])^{-1} \sum_{s=1}^m [1 - s/(m+1)] \sum_{t=s+1}^{[\lambda T]} \hat{\varepsilon}_{0,t} \hat{\varepsilon}_{0,t-s}\end{aligned}$$

The functions $H_j(NM_1(\cdot))$ and $H_j(NM_0(\cdot))$, $j = 1, 2, 3$ are defined as before. Evidence against the null is provided by large values of the test statistics.² As with the ratio-based tests, BT show that the NM tests generally have low power against the “wrong” alternative.

3.5 Cusum of Squares-based Tests

Leybourne et al. (2006) propose tests of the null hypothesis of a stable unit root against the alternative of a change in persistence from trend stationarity to difference stationarity, or vice-versa. They can be treated as complementary to tests that take trend stationarity as the null hypothesis. The tests are based on standardized cumulative sums of squared sub-sample residuals. One of these is a ratio test defined as

$$R = \frac{\inf_{\lambda \in \Lambda} K^f(\lambda)}{\inf_{\lambda \in \Lambda} K^r(\lambda)} \equiv \frac{N}{D}$$

where

$$\begin{aligned}K^f(\lambda) &= \frac{([\lambda T])^{-2} \sum_{t=1}^{[\lambda T]} \hat{\varepsilon}_{0t}^2}{\hat{\omega}_f^2(\lambda)} \\ K^r(\lambda) &= \frac{(T - [\lambda T])^{-2} \sum_{t=1}^{(T-[\lambda T])} \hat{\varepsilon}_{1t}^2}{\hat{\omega}_r^2(\lambda)} \\ \hat{\omega}_f^2(\lambda) &= ([\lambda T])^{-1} \sum_{t=1}^{[\lambda T]} \Delta \hat{\varepsilon}_{0t}^2 + 2([\lambda T])^{-1} \sum_{s=1}^m [1 - s/(m+1)] \sum_{t=1}^{[\lambda T]} \Delta \hat{\varepsilon}_{0,t} \Delta \hat{\varepsilon}_{0,t-s} \\ \hat{\omega}_r^2(\lambda) &= (T - [\lambda T])^{-1} \sum_{t=1}^{T-[\lambda T]} \Delta \hat{\varepsilon}_{1t}^2 + 2(T - [\lambda T])^{-1} \sum_{s=1}^m [1 - s/(m+1)] \sum_{t=1}^{T-[\lambda T]} \Delta \hat{\varepsilon}_{1,t} \Delta \hat{\varepsilon}_{1,t-s}\end{aligned}$$

²Busetti and Taylor (2004) note that the maximum of the two NM tests turned out to have very low power and were therefore not considered in their analysis.

Leybourne et al. (2006) show that a consistent test of the unit root null against the $I(0)$ - $I(1)$ [$I(1)$ - $I(0)$] alternative can be obtained from the left-tail (right-tail) distribution of R . Further, when the ratio test correctly rejects the null of no persistence change, the tail in which the rejection occurs can be used to identify the direction of change since the test almost never rejects in the right [left] tail when there is a change from $I(0)$ to $I(1)$ [$I(1)$ to $I(0)$]. A consistent test against the $I(0)$ - $I(1)$ [$I(1)$ - $I(0)$] change can also be obtained by using the test which rejects for small values of N [D]. Consequently, to test the null hypothesis against either alternative, one could consider the statistic $M = \min\{D, N\}$. A test which rejects for small values of M is consistent against either alternative.

As a byproduct of their analysis, Leybourne et al. (2006) also propose consistent estimators for the unknown breakpoint. In the $I(0)$ - $I(1)$ case, the estimate is

$$\hat{\lambda}_{01} = \arg \inf_{\lambda \in \Lambda} ([T\lambda])^{-2} \sum_{t=1}^{[T\lambda]} \hat{\varepsilon}_{0t}^2$$

while in the $I(1)$ - $I(0)$ case, it is

$$\hat{\lambda}_{10} = \arg \inf_{\lambda \in \Lambda} (T - [T\lambda])^{-2} \sum_{t=1}^{(T-[T\lambda])} \hat{\varepsilon}_{1t}^2$$

4 Empirical Results

The data for our empirical analysis is obtained from the ECB's Area Wide Model (AWM) quarterly database described in Fagan, Henry and Mestre (2001). The sample period is 1970Q1-2005Q4. The measure of inflation (y_t) used is the annualized quarterly log-difference of the GDP deflator.³ This is calculated as $y_t = 400(\ln P_t - \ln P_{t-1})$, where P_t is the GDP deflator in quarter t . Figure 1 provides a plot of the data. The figure suggests the possibility of non-stationary behavior in the '70s and '80s while inflation appears to be much more stable from the early '90s. The plot also indicates a substantially lower level of inflation in this latter period. In this section and the next, we will argue that inflation in the euro area is better characterized as a process which has undergone a shift in level as well as a shift in

³An alternative measure is based on the Harmonized Index of Consumer Prices (HICP). O'Reilly and Whelan (2005) report results for parameter stability tests only for the GDP deflator measure noting that the non-seasonally adjusted HICP data may exhibit instabilities over time due to changing seasonal patterns. Since our empirical analysis is primarily based on tests for parameter stability, we focus on the GDP deflator measure in this paper.

persistence as opposed to a shift in level alone.

It is useful to note at this point that the empirical analysis in the paper is exclusively based on tests for a one time structural change. While the possibility of multiple structural changes cannot be ruled out, the assumption of a single break appears reasonable from an informal inspection of Figure 1. Moreover, our empirical results provide strong evidence against the null of stable persistence for euro area inflation. Given that the single break tests generally have low power against multiple changes, this further suggests that the assumption of a single change is likely to be adequate. It would nevertheless be useful to estimate the number of structural changes endogenously from the data. However, the econometric literature on persistence change has been primarily developed for the case of a single break and no procedures are currently available (to the best of our knowledge) to consistently estimate the number of breaks in persistence allowing for unit root non-stationarities in some regimes.⁴

The empirical results of our analysis are organized in four subsections. Specifically, in subsection 4.1, we present the results of persistence change tests described in Section 3. Subsection 4.2 contains the estimates of the parameters which include the break date and the persistence parameter (the sum of the autoregressive coefficients) which is computed from least squares regressions as well as employing the bias correction procedure advocated in Andrews and Chen (1994). Alternative measures of persistence such as the half life of shocks and confidence intervals for the largest autoregressive root are presented in subsection 4.3. Finally, subsection 4.4 demonstrates that the hypothesis of a stationarity with a mean shift but a stable persistence parameter is not compatible with euro area inflation data.

4.1 Persistence Change Tests

Table 1 presents the results from the persistence change tests applied to euro area inflation. Given the wide range of tests employed, we also present in Table 1 the corresponding null and alternative hypotheses for each of the tests. In addition to the tests discussed in Section 3, we also present results for the commonly used Sup- F test (Andrews, 1993) as well as the Mean- F and Exp- F tests (Andrews and Ploberger, 1994) for stability of the persistence parameter. We will argue that the empirical results point to an $I(1)$ - $I(0)$ alternative as

⁴A potential strategy to investigate the possibility of multiple changes is to apply the single break tests to subsamples identified by the first break. Monte Carlo simulations, however, indicated that the tests can have poor finite sample properties for samples of size less than 100. Thus, one is likely to obtain a misleading estimate of the number of structural changes from applying such a procedure.

opposed to a $I(0)$ - $I(1)$ alternative. First, the ratio-based tests that are directed against the $I(0)$ - $I(1)$ alternative provide only very weak evidence against the null. In contrast, the tests based on $K_M(\cdot)^{-1}$ designed for the $I(1)$ - $I(0)$ alternative all comprehensively reject the null of stable persistence. As discussed in BT and Section 3, the ratio-based tests against the incorrect alternative have low power thereby suggesting a shift from unit root behavior to stationarity. The maximum of the ratio-based tests are also all significant at conventional levels. Second, none of the modified tests designed to detect the $I(0)$ - $I(1)$ alternative reject the null of stability. On the other hand, the modified tests against the $I(1)$ - $I(0)$ alternative as well as the maximum of the modified tests decisively reject the null in favor of a change in persistence. Third, all LBI-based tests are significant, at least at the 5% level. This is also to be expected since, as noted before, these tests have power against both types of alternatives. Fourth, The NM tests against both types of alternatives are generally significant and so do not provide any direct evidence regarding the direction of the shift. Fifth, the R test rejects in the right tail which, following the reasoning in Section 3, is again supportive of an $I(1)$ - $I(0)$ change. Further, the D and M tests reject the null while the N test does not, again indicative of the $I(1)$ - $I(0)$ nature of the process. Finally, it is worth noting that none of $\text{Sup-}F$, $\text{Mean-}F$ or $\text{Exp-}F$ provide any evidence against the null of stability.

Given that the above test results are suggestive of an $I(1)$ - $I(0)$ type behavior for inflation, we now turn to estimating the break date and the persistence parameters over the regimes identified by the estimated break date.

4.2 Parameter Estimates

The breakpoint estimates $\hat{\lambda}_M$ and $\hat{\lambda}_{10}$ are presented in panel (A) of Table 2. The break date identified by both methods is the second quarter of 1993 irrespective of the sample period used. As discussed in section 5, such a date may be expected based on economic events that may have contributed to a reduction in the level of inflation persistence. In order to uncover the extent of persistence in the two identified regimes, we estimate an $AR(p_i)$ model over each regime i ($i = 1, 2$) where the lag length p_i is selected using Ng and Perron's (1995) general-to-specific sequential procedure based on a 5% two-sided t -test on the last included lag (assuming a maximum of 10 lags). The model for regime i is

$$y_t = \alpha_i y_{t-1} + \sum_{j=1}^{p_i-1} \phi_{i,j} \Delta y_{t-j} + u_t \quad (11)$$

In (11), α_i ($i = 1, 2$) denote the sum of the autoregressive coefficients. The corresponding least squares estimates, denoted $\hat{\alpha}_{1,ols}$ and $\hat{\alpha}_{2,ols}$, are seen to differ markedly over the two regimes. In the first regime, the persistence parameter is estimated at 0.93 while the estimate $\hat{\alpha}_{2,ols}$ is 0.43, far smaller. This is clearly indicative of a substantial reduction in estimated persistence after the break. It is well known that the least squares estimates may potentially suffer from a substantial downward bias when the process is persistent. In order to mitigate the effects of such a bias, we also present bias-corrected estimates as well as confidence intervals based on the procedure proposed in Andrews and Chen (1994).⁵ The bias-adjusted estimate in each regime (denoted $\hat{\alpha}_{i,ac}$; $i = 1, 2$ in Table 2) is hardly any different from the corresponding least squares estimate. Both the standard and Andrews and Chen confidence intervals include unity in the first regime, consistent with the results from persistence change tests. For the second regime, the standard interval is somewhat wide although the upper end is only 0.77, much below unity. The Andrews and Chen confidence interval is much shorter with an upper bound similar to that for the standard interval. The main message from these results is, however, clear: there has been an important decline in inflation persistence in the period following the break.

4.3 Alternative Measures of Persistence

The usefulness of the sum of autoregressive coefficients α_i as a measure of persistence arises from the fact that, for $\alpha_i \in (-1, 1)$, it is directly related to the cumulative impulse response following a shock, given by $1/(1 - \alpha_i)$. A larger α_i therefore intuitively corresponds to higher persistence of inflation. Phillips (1991) discusses a second interpretation of the parameter α_i in terms of the spectrum of y_t . The spectrum at zero frequency is a measure of the low-frequency autocovariance of the series. For the model (11), it is given by $\text{Var}(u_t)/(1 - \alpha_i)^2$. Hence, according to this measure too, the persistence of y_t depends on the magnitude of the parameter α_i . The main problem with this measure, as discussed by Pivetta and Reis (2007), is that it is large for a process with an impulse response function where inflation rises quickly and subsequently falls steeply back to zero, compared to a process with a relatively slow initial increase and a slowly decaying impulse response, despite the second being intuitively more persistent.

An alternative measure of persistence is the largest autoregressive root (LAR), which we denote by ρ_i . In the distant future, the impulse response of inflation to a shock is dominated

⁵These estimates were computed using 1000 Monte Carlo replications.

by the largest root so that the magnitude of ρ_i is an important determinant of the length of time for which the effects of shocks will persist. Methods to obtain median unbiased estimates and confidence intervals for ρ_i are discussed in Stock (1991). Based on the local-to-unity model $\rho_i = 1 + c_i/T_i$, where T_i denotes the number of observations in regime i , he provides tables based on which we can obtain a median unbiased estimate and a confidence interval for c_i , from where a confidence interval for ρ_i follow directly. The median unbiased estimates, denoted $\hat{\rho}_{1,med}$ and $\hat{\rho}_{2,med}$, together with 90% confidence intervals (in brackets beside the point estimates) for ρ_1 and ρ_2 , are presented in panel (B) of Table 2. The point estimate in the first regime exceeds unity while for the second regime, the LAR estimate is only 0.69. The confidence interval for the first regime includes unity while that for the second regime does not, although the upper end is somewhat close to one. These results are again consistent with the notion that inflation persistence has declined considerably after 1993.

Our third measure of persistence is the half life, defined as the number of periods in which inflation remains above 0.5 following a unit shock. Rossi (2005) proposes measures of half life in general autoregressive models. For an $AR(p_i)$ model estimated over regime i observations, the median unbiased estimate of the half life is

$$\hat{h}_{i,med} = \max \left\{ \frac{\ln(1/2\hat{b}_i(1))}{\ln(\hat{\rho}_{i,med})}, 0 \right\}$$

We set $\hat{h}_{i,med} = \infty$ if $\hat{\rho}_{i,med} \geq 1$. The estimate $\hat{b}_i(1)$ is given by

$$\hat{b}_i(1) = 1 - \sum_{j=1}^{p_i-1} \hat{\phi}_{i,j}$$

where $\{\hat{\phi}_{i,j}\}$ are the least squares estimates from (11). Based on the confidence interval for ρ_i , we can then directly get, by monotonicity, the corresponding 90% confidence interval for the half life. The half life estimates and confidence intervals are presented in panel (B) of Table 2. The half life estimate is infinity in the first regime given that the median unbiased estimate $\hat{\rho}_{1,med}$ exceeds unity. For the second regime, the half life estimate is about 2 quarters suggesting that the effects of shocks dissipate quite rapidly in this regime. The confidence interval is however quite wide with the upper bound being about 10 quarters.

It is useful to note that the LAR and the half life as measures of persistence are also not immune to criticism. For instance, the problem with LAR is that it ignores the effects

of the other roots. While the LAR may be a reasonable approximation to persistence, considering more roots will provide better approximations. Moreover, the half life is likely to underestimate the true persistence of the process if the impulse response function is oscillating. Our objective in considering alternative persistence measures is to strengthen our conclusion regarding a change in persistence following the break in the inflation process.

4.4 Stationarity with a Mean Shift

As discussed in Section 2, recent work by Gadzinski and Orlandi (2004), Levin and Piger (2003), Cecchetti and Debelle (2006), Corvoisier and Mojon (2005) suggest that inflation in the euro area is well characterized by a process which undergoes a shift in its mean but whose persistence parameter remains stable. They argue that ignoring a mean break leads to overestimating the extent of inflation persistence and that once one controls for the break, measured persistence is much lower. Indeed, the annualized inflation rate is 7.51% over 1970Q1-1993Q2 and 1.98% over 1993Q3-2005Q4. Moreover, as shown in Belaire-Franch (2005), persistence change tests which take stationarity as the null have power against processes which display a pure mean shift so that a rejection by these tests could occur even if the true process for inflation involves no change in persistence. In what follows, we illustrate that, for the euro area, the hypothesis of a pure mean shift in inflation without an accompanying shift in persistence is generally not supported by the data.

Our analysis is based on unit root as well as stationarity tests that allow for a break in mean. First, we conduct unit root tests proposed by Perron and Vogelsang (1992) that allow for a break in mean under both the null and alternative hypotheses. They consider two models: the additive outlier (AO) model where the change is assumed to take effect instantaneously and the innovative outlier model (IO), where the change affects the level of the series only gradually. We present results for both models. The construction of the test statistics is first discussed for a given break date T_b . Methods to select T_b will be discussed subsequently. First, consider the AO model. For a fixed break date T_b , we obtain the residuals from running the OLS regression

$$y_t = \mu + \delta DU_t + \tilde{y}_t, \quad t = 1, \dots, T \quad (12)$$

where $DU_t = I(t > T_b)$. We then construct the t -statistic, denoted $t_\alpha(\text{AO}, T_b, k)$, for testing

$\alpha = 1$ in the regression

$$\tilde{y}_t = \sum_{a=0}^k \omega_a D(TB)_{t-a} + \alpha \tilde{y}_{t-1} + \sum_{j=1}^k c_j \Delta \tilde{y}_{t-j} + e_t, \quad t = k+2, \dots, T$$

with $D(TB)_t = I(t = T_b + 1)$. For the IO model, we estimate the following regression by OLS:

$$y_t = \mu + \delta DU_t + \theta D(TB)_t + \alpha y_{t-1} + \sum_{j=1}^k c_j \Delta y_{t-j} + e_t, \quad t = k+2, \dots, T$$

We then compute the t -statistic for testing $\alpha = 1$, denoted $t_\alpha(\text{IO}, T_b, k)$.

Perron and Vogelsang (1992) propose two methods to select T_b for both AO and IO models. In the first method, the break date is selected by minimizing the t -statistic over all permissible break dates. That is, the test statistic in the AO case is

$$PV_1(\text{AO}) = \inf_{T_b \in (k+2, T)} t_\alpha(\text{AO}, T_b, k)$$

while that for the IO case is

$$PV_1(\text{IO}) = \inf_{T_b \in (k+2, T)} t_\alpha(\text{IO}, T_b, k)$$

The second procedure selects T_b by maximizing the t -statistic for testing $\delta = 0$ in each regression. The resulting statistics are denoted by $PV_2(i)$, $i = \text{AO}, \text{IO}$. The lag length k is chosen using a sequential procedure which entails testing the significance of the last included lag, starting from a prespecified maximum order k_{\max} , and stopping when this lag coefficient is significant. We use a 5% two-sided test for evaluating significance with $k_{\max} = 5$.

As a complement to the above unit root tests, we also present results from stationarity tests in the presence of a mean shift, proposed in Kurozumi (2002). Unlike the unit root tests that only allow but do not impose the existence of a break, the stationarity tests presume that a break in mean exists and so in order to apply these tests, we first need to verify the presence of a break. To do so, we apply the mean shift tests proposed by Vogelsang (1998) which are valid whether or not the errors have a unit root. For a given break date T_b , consider estimating regression (12) by OLS and constructing the standard Wald statistic for testing $\delta = 0$. Let $PS_T(T_b)$ denote this Wald statistic divided by the sample size T . Next,

we estimate by OLS the regression

$$y_t = \mu + \delta DU_t + \sum_{i=1}^9 \gamma_i t^i + u_t$$

and compute the statistic, denoted by $J_T(T_b)$, defined as T^{-1} times the standard Wald statistic for testing the joint hypothesis that $\gamma_1 = \gamma_2 = \dots = \gamma_9 = 0$. Then, since the shift date is unknown, Vogelsang (1998) considers the following three functionals:

$$\begin{aligned} \text{mean } PS_T &= \left\{ T^{-1} \sum_{T_b \in \Lambda} PS_T(T_b) \right\} \exp(-bJ_T^*) \\ \text{exp } PS_T &= \log \left\{ T^{-1} \sum_{T_b \in \Lambda} \exp\left(\frac{1}{2} PS_T(T_b)\right) \right\} \exp(-bJ_T^*) \\ \text{sup } PS_T &= \left\{ \sup_{T_b \in \Lambda} \text{exp } PS_T(T_b) \right\} \exp(-bJ_T^*) \end{aligned}$$

where $J_T^* = \inf_{T_b \in \Lambda} J_T(T_b)$ and b is a constant chosen such that the critical values in the stationary case are close to those in the unit root case.

Once the presence of a mean break is confirmed by these tests, an estimate of the break date, denoted \hat{T}_b , is obtained by minimizing the sum of squared residuals of y_t on a constant and DU_t . The residuals based on the estimated breakpoint are denoted \tilde{y}_t^* . Given the estimated break date and the associated vector of residuals, we can proceed to test for stationarity in the presence of a mean break.

Kurozumi (2002) proposes tests for the null hypothesis of (trend) stationarity with a structural change against a unit root. The first test is an LM test defined by

$$S_T = \frac{1}{\tilde{\sigma}^2 T^2} \sum_{j=1}^T \left(\sum_{t=1}^j \tilde{y}_t^* \right)^2$$

where

$$\tilde{\sigma}^2 = T^{-1} \sum_{t=1}^T (y_t^*)^2 + 2T^{-1} \sum_{s=1}^m [(1 - s/(m+1))] \sum_{t=s+1}^T \tilde{y}_t^* \tilde{y}_{t-s}^*$$

A second test, whose asymptotic distribution under the null hypothesis does not depend on the breakpoint, based on previous work by Park and Sung (1994) [and hence the superscript

ps], is given by

$$S_T^{ps} = \frac{1}{\tilde{\sigma}^2 T^2} \sum_{j=1}^{T-1} \left(\sum_{t=1}^j \tilde{y}_t^{ps} \right)^2$$

with

$$\tilde{y}_t^{ps} = \begin{cases} (T/\hat{T}_b)y_t & t = 1, \dots, \hat{T}_b \\ [T/(T - \hat{T}_b)]y_t & t = \hat{T}_b + 1, \dots, T \end{cases}$$

The parameter m is chosen according to the data dependent method suggested in Kurozumi (2005):

$$\hat{m}_l = \min \left(1.1447 \left\{ \frac{4\hat{a}^2 T}{(1 + \hat{a})^2 (1 - \hat{a})^2} \right\}^{1/3}, 1.1447 \left\{ \frac{4l^2 T}{(1 + l)^2 (1 - l)^2} \right\}^{1/3} \right)$$

where \hat{a} is the OLS estimate obtained by estimating an AR(1) model for the residuals. Size and power simulations in Kurozumi (2002) indicate that $l = 0.7$ is a reasonable choice in finite samples.

The results of the unit root, mean shift and stationarity tests are presented in Table 3. First, the unit root tests do not provide any evidence against the null for both the AO and IO models. Next, Vogelsang's mean shift tests generally reject the null of a constant mean. Given the presence of a mean shift, the stationarity tests S_T and S_T^{ps} were computed. The null of stationarity is rejected by both tests. This analysis therefore suggests that a process that is stationary except for a shift in mean somewhere in the sample does not provide an adequate representation of the inflation process for the euro area.

5 Discussion

Both methods that were employed to select the date of the break in the inflation process identified the change in the second quarter of 1993. While inflation was found to be strongly persistent in the period prior to the break, it was markedly less so in the period following the break. This was confirmed by different measures of persistence, including the sum of autoregressive coefficients, confidence intervals for the largest autoregressive root as well as estimates of the half life. One plausible explanation for the identified break date is that it occurred during the time the groundwork for the EMU was being established, with price stability mandated as the primary objective of monetary policy. Plans for the EMU were

formalized in provisions within the Maastricht Treaty signed in February 1992, which founded the European Union. The Treaty was subsequently ratified by all of the member states and finally came into effect in November 1993. It set up the conditions, or “convergence criteria” which each member state in the European Union must meet before it could join the EMU. These criteria included the stipulation that a country had to achieve a rate of inflation within 1.5% of the rates in the three participating countries with the lowest rates. The Treaty thus created an institutional commitment to price stability by mandating it be the long-run objective of monetary policy. In other words, while not explicitly adopted, the fundamentals of inflation targeting were codified in the Maastricht Treaty (see Bernanke et al., 1999, for a discussion pertaining to a proposal for inflation targeting in the EMU). A commitment to price stability entails vigorous efforts to communicate with the public about the plans and objectives of the monetary authorities, thereby strengthening the credibility of the monetary regime for attaining these objectives. To the extent that increased credibility anchors inflation expectations, it reduces the real economic costs of a disinflation and hence contributes to an improved trade-off between inflation and unemployment (a lower sacrifice ratio).

As argued in Taylor (1998), the credibility of a monetary policy regime can be modeled in terms of its influence on the degree to which expectations are forward looking. If the inflation target is perfectly credible, it would cause all price setters to adopt purely forward looking inflation expectations, anchored on the inflation target, as in most New Keynesian models. A credible target is therefore likely to induce a marked decline in inflation persistence from the previous regime where there was a strong backward looking element to expectations.

Recent work by Erceg and Levin (2003) suggests that the persistence of inflation is not an inherent characteristic of the economy but rather varies with the credibility and transparency of the monetary regime. They show that inflation persistence can emanate from the public’s limited information about the central bank’s policy objectives. These authors formulate a dynamic general equilibrium model with optimizing agents and staggered nominal contracts, in which private agents use optimal filtering to make inferences about the central bank’s inflation target. The speed at which private agents recognize a new inflation target depends on the transparency and credibility of the central bank. The signal-to-noise ratio plays a key role in determining the persistence of inflation forecast errors which in turn influences the persistence of actual inflation. Orphanides and Williams (2003) simulate a similar model which illustrates that the absence of a long-run inflation objective for the monetary authority leads to substantially higher inflation persistence relative to an

environment where the inflation objective is clearly understood by price-setters.

Our results thus suggest that inflation persistence should not be treated as a structural feature of the economy in dynamic macroeconomic models for the euro area. For instance, the hybrid New Keynesian Phillips Curve introduces a backward looking component in the otherwise purely forward looking model by assuming that a fraction of the firms set their prices according to a rule of thumb (Gali and Gertler, 1999; Steinsson, 2003). The hybrid model is popular in empirical work as it is better able to match the supposedly high observed persistence of inflation (see Gali and Gertler, 1999 and Gali, Gertler and Lopez-Salido, 2001 for the U.S. and euro area evidence, respectively). Based on our empirical analysis, we argue that such models are unlikely to be structural in the sense of Lucas (1976) and are therefore not suitable for evaluating the desirability of alternative monetary policy regimes. It is also questionable to what extent these models can be used to formulate optimal monetary policies. If optimal policy based on a particular estimate of inflation persistence differs significantly from actual policy over the estimation period, the implementation of such a policy would change the structure of the economy thereby rendering the policy sub-optimal.

6 Concluding Remarks

This paper revisits the issue of the nature and degree of inflation persistence in the euro area. The results in this paper do not support the view that inflation persistence has been high and stable over the post-1970 period. We argue that O'Reilly and Whelan's (2005) finding of a stable persistence parameter can be attributed to the use of procedures which do not specifically allow for the presence of a unit root in a sub-sample of the data. Based on persistence change tests which allow for such unit root behavior, we find evidence in favor of a marked drop in persistence in the second quarter of 1993. The decline in persistence is confirmed by alternative persistence measures. Our findings therefore suggest that inflation persistence is not an intrinsic characteristic of the economy that should be incorporated into the structure of general equilibrium macroeconomic models. The reason is that such backward looking models are not structural in the sense of Lucas (1976) and so are of not much use for quantitative policy evaluation or the computation of optimal policies. Our results are instead consistent with forward looking rational expectations models which imply that the nature of inflation persistence varies with the transparency and credibility of the monetary regime.

It is important to note that the findings of the paper are based on data for the euro

area as a whole. Benigno and Lopez-Salido (2006) provide evidence suggesting the presence of heterogeneity in inflation dynamics across euro area countries. For instance, German inflation is found to have a dominant forward looking component while inflation dynamics in France, Italy, Spain and the Netherlands are characterized by significant inertial behavior.⁶ These inflation differentials across regions are shown to matter for the design of monetary policy. It would be useful to employ the methods of this paper to conduct a cross country inflation analysis to understand how the degree of inertia and structural conditions differ across countries. Another aggregation issue pertains to the fact that our analysis is based on an aggregate measure of inflation as opposed to its individual components. Altissimo et al. (2007) conduct a sectoral analysis of inflation persistence for the euro area over the period 1985-2003 and find substantial heterogeneity across sectors, with non-processed food and energy generally exhibiting less persistence, while services and industrial goods are more persistent. There also seems to be a distinct difference between the persistence of aggregate inflation and the average persistence of the disaggregated time series. One potential explanation is that idiosyncratic shocks to the sub-indices will tend to offset each other if a sufficiently large number of series is aggregated so that the aggregate time series will appear smoother as it can be expected to be dominated by the common shocks only. Again, the methods of this paper can, in principle, be used to identify whether a change in persistence across subcomponents have a common source. Further exploration of the relationship between the dynamic properties of aggregate euro area inflation and its country and sectoral subcomponents is left as an important avenue for future research.

⁶Altissimo et al. (2006) provide a useful summary of studies that estimate the extent of inflation persistence in euro area countries. There is a wide range of estimates across countries and studies and the studies disagree considerably on the country rankings.

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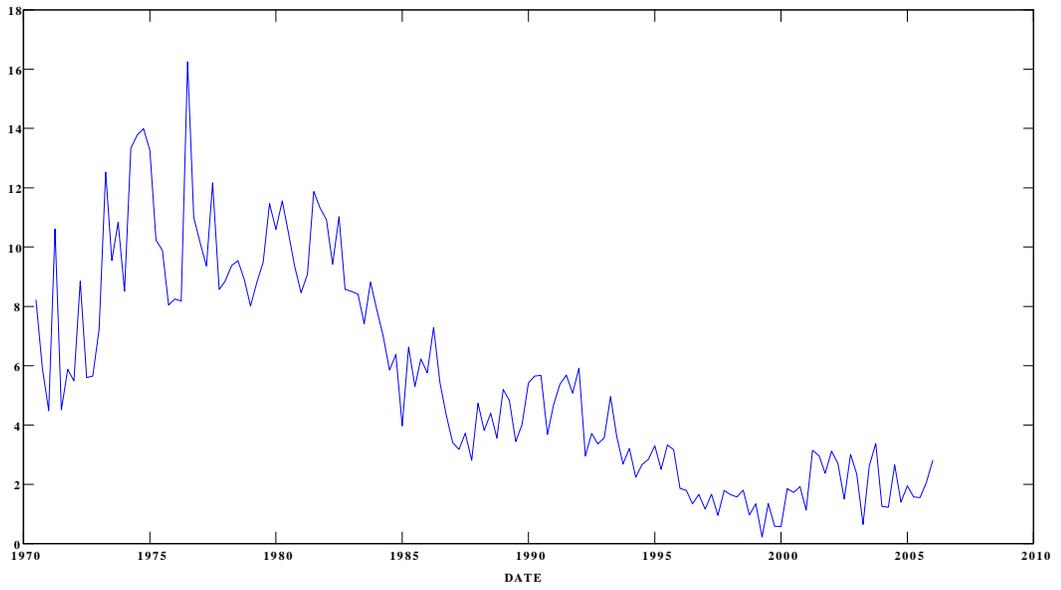


Figure 1: Euro Area Inflation (1970Q1-2005Q4)

Table 1: Tests for a Change in Persistence

Source	Test Statistic	H_0	H_1	Sample Value
Andrews ('93)	Sup- F	$I(0)$	$I(0)-I(0)$	4.25
Andrews & Ploberger ('94)	Mean- F	$I(0)$	$I(0)-I(0)$	1.38
	Exp- F	$I(0)$	$I(0)-I(0)$	0.96
Kim ('00)	$H_1(K_M(\cdot))$	$I(0)$	$I(0)-I(1)$	12.64
	$H_2(K_M(\cdot))$	$I(0)$	$I(0)-I(1)$	3.31
	$H_3(K_M(\cdot))$	$I(0)$	$I(0)-I(1)$	3.87*
Buseti & Taylor ('04)	$H_1((K_M(\cdot))^{-1})$	$I(0)$	$I(1)-I(0)$	288.38***
	$H_2((K_M(\cdot))^{-1})$	$I(0)$	$I(1)-I(0)$	55.67***
	$H_3((K_M(\cdot))^{-1})$	$I(0)$	$I(1)-I(0)$	139.73***
	$\max H_1(K)$	$I(0)$	$I(1)-I(0)/I(0)-I(1)$	288.38***
	$\max H_2(K)$	$I(0)$	$I(1)-I(0)/I(0)-I(1)$	55.67***
	$\max H_3(K)$	$I(0)$	$I(1)-I(0)/I(0)-I(1)$	139.73***
	$H_1(L_1(\cdot))$	$I(0)$	$I(0)-I(1)$	2.27**
	$H_2(L_1(\cdot))$	$I(0)$	$I(0)-I(1)$	1.97***
	$H_3(L_1(\cdot))$	$I(0)$	$I(0)-I(1)$	1.00***
	$H_1(L_0(\cdot))$	$I(0)$	$I(1)-I(0)$	2.43**
	$H_2(L_0(\cdot))$	$I(0)$	$I(1)-I(0)$	1.95***
	$H_3(L_0(\cdot))$	$I(0)$	$I(1)-I(0)$	1.00***
	$\max H_1(L)$	$I(0)$	$I(1)-I(0)/I(0)-I(1)$	2.43**
	$\max H_2(L)$	$I(0)$	$I(1)-I(0)/I(0)-I(1)$	1.97***
	$\max H_3(L)$	$I(0)$	$I(1)-I(0)/I(0)-I(1)$	1.00***
	$H_1(NM_1(\cdot))$	$I(0)$	$I(0)-I(1)$	0.99**
	$H_2(NM_1(\cdot))$	$I(0)$	$I(0)-I(1)$	0.58***
	$H_3(NM_1(\cdot))$	$I(0)$	$I(0)-I(1)$	0.30***
	$H_1(NM_0(\cdot))$	$I(0)$	$I(1)-I(0)$	0.90*
	$H_2(NM_0(\cdot))$	$I(0)$	$I(1)-I(0)$	0.46**
	$H_3(NM_0(\cdot))$	$I(0)$	$I(1)-I(0)$	0.24**
	Harvey et. al ('06)	$H_1^m(K_M(\cdot))$	$I(1)/I(0)$	$I(0)-I(1)$
$H_2^m(K_M(\cdot))$		$I(1)/I(0)$	$I(0)-I(1)$	(2.28,1.91,1.08)
$H_3^m(K_M(\cdot))$		$I(1)/I(0)$	$I(0)-I(1)$	(2.19,1.72,0.84)
$H_1^m((K_M(\cdot))^{-1})$		$I(1)/I(0)$	$I(1)-I(0)$	(207.24*,175.55**,102.19***)
$H_2^m((K_M(\cdot))^{-1})$		$I(1)/I(0)$	$I(1)-I(0)$	(42.79*,37.27**,24.44***)
$H_3^m((K_M(\cdot))^{-1})$		$I(1)/I(0)$	$I(1)-I(0)$	(92.37*,77.42**,44.27***)
$\max H_1^m(K)$		$I(1)/I(0)$	$I(1)-I(0)/I(0)-I(1)$	(175.75*,144.62**,75.91***)
$\max H_2^m(K)$		$I(1)/I(0)$	$I(1)-I(0)/I(0)-I(1)$	(37.31*,31.74**,18.95***)
$\max H_3^m(K)$		$I(1)/I(0)$	$I(1)-I(0)/I(0)-I(1)$	(77.45*,63.15**,32.75***)
Leybourne et. al ('06)	R	$I(1)/I(0)$	$I(1)-I(0)/I(0)-I(1)$	7.32**
	N	$I(1)$	$I(0)-I(1)$	0.10
	D	$I(1)$	$I(1)-I(0)$	0.01***
	M	$I(1)$	$I(1)-I(0)/I(0)-I(1)$	0.01***

Table 2: Parameter Estimates and Measures of Persistence

Estimator/Measure	Sample Value
(A) Parameter Estimates	
$\hat{\lambda}_M$	1993Q2
$\hat{\lambda}_{10}$	1993Q2
$\hat{\alpha}_{1,ols}$	0.93; [0.83,1.03]
$\hat{\alpha}_{2,ols}$	0.43; [0.09,0.77]
$\hat{\alpha}_{1,ac}$	0.93; [0.88,1.00]
$\hat{\alpha}_{2,ac}$	0.44; [0.37,0.80]
$k_{bic}(1)$	3
$k_{bic}(2)$	8
(B) Measures of Persistence	
$\hat{\rho}_{1,med}$	1.01
$\hat{\rho}_{2,med}$	0.69
90% <i>CI</i> for ρ_1	(0.94,1.04)
90% <i>CI</i> for ρ_2	(0.48,0.94)
$\hat{b}_1(1)$	1.91
$\hat{b}_2(1)$	0.92
$\hat{h}_{1,med}$	∞
$\hat{h}_{2,med}$	2.01; (1.17,10.30)

Table 3: Tests for Stationarity with a Mean Shift

Test	Sample Value
$PV(AO1)$	-3.44
$PV(AO2)$	-2.74
$PV(IO1)$	-3.48
$PV(IO2)$	-1.96
sup PS_T	(13.66*,10.72**,4.68)
mean PS_T	(3.901*,3.484**,2.436)
exp PS_T	(4.282*,3.231**,1.320)
S_T	0.35***
S_T^{ps}	1.03**

Note: In Tables 1 and 3, '*', '**', and '***' denote significance at the 10%, 5%, and 1% levels respectively. For the sup, mean and exp statistics, we present results in the form (a, b, c) , which represent the statistics computed at the 10%, 5% and 1% levels respectively.