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

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Abstract

Recent empirical studies investigating the time series behavior of euro area inflation suggest that inflation persistence has been generally high and stable over the post-1970 period. Their methodology is primarily based on standard unit root and structural break tests that are not designed to detect changes in persistence when the process shifts from stationarity to non-stationarity or vice-versa. This paper employs a variety of tests that allows for such shifts as well as consistent break date estimation methods to argue that euro area inflation shifted from a (near) 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. Bias-corrected estimates of regime-specific persistence parameters, half-life estimates and confidence intervals for the largest autoregressive root all suggest a marked decline in persistence after the estimated break date. We also argue that the hypothesis of stationarity with mean shifts but a stable persistence parameter does not appear to provide an adequate description of the data. The evidence presented is consistent with the view that the degree of inflation persistence varies with the transparency and credibility of the underlying monetary regime.

Keywords: persistence, price stability, structural break, unit root

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 an inherent characteristic of the economy that should be explicitly incorporated while formulating macroeconomic models. Evidence in favor of high and unchanged persistence from reduced form specifications across different policy regimes has generally been construed as suggesting that inflation persistence is a feature that any reasonable model for the economy should be able to replicate. Several approaches have been adopted in developing the microeconomic foundations for 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).

Given the sequence of monetary policy shifts that has occurred since the early 1970s in the euro area, one would a priori expect the backward looking reduced form representations for inflation to be characterized by substantial parameter instability. This is indeed the essence of the Lucas Critique according to which the parameters of macroeconometric models depend implicitly on agents' expectations of the policy process and are unlikely to remain stable as policymakers change their behavior, if agents are forward looking. An empirical finding of high and stable persistence in such a context can potentially be interpreted either in terms of the presence of a strong backward looking component in the dynamics of inflation induced through, say indexation or rule-of-thumb behavior on the part of the price setters, or in terms of the historical policy shifts being of relatively modest magnitude.[1] While the latter explanation seems implausible for the euro area given the large empirical literature that points at evidence to the contrary, the former suggests that purely forward looking models based on rational expectations such as the New Keynesian Phillips curve are incompatible

[1] For instance, Rudebusch (2005) uses a New Keynesian style macroeconomic model to demonstrate that if the underlying structure in such models places relatively low weights on forward looking expectational variables, then the inflation persistence parameter in reduced form specifications will be close to unity.

with euro area inflation dynamics. In such models, the extent of dependence of inflation on its past values is likely to diminish as the credibility of a central bank's commitment to low inflation increases (Taylor, 1998). This possibility appears especially relevant for the euro area given the European Central Bank's adoption of an explicit mandate for maintaining price stability as the overriding objective of monetary policy.

The reduced form specifications are also routinely employed to generate out-of-sample forecasts of the inflation rate. The stability of these representations is essential in this context since ignoring the presence of structural breaks is likely to result in very inaccurate and biased forecasts. Furthermore, a careful analysis of the stability of the inflation persistence parameter is warranted in that this parameter plays a crucial role in determining the impulse response and propagation mechanisms of inflationary shocks.

Recent empirical studies investigating the time series behavior of euro area inflation suggest that inflation persistence has been generally high and stable over the post-1970 period. Their econometric methodology is primarily 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 portion 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 (see Kim, 2003). 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 reassess the empirical relevance of the Lucas Critique for euro area inflation employing a variety of tests for shifts in persistence that allow for unit root non-stationarities of the form described above to argue that the reduced form process for euro area inflation shifted from a (near) 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 with its strict mandate for price stability came into effect, thereby suggesting a role for inflation targeting. The Treaty laid the foundation for the establishment of Europe's Economic and Monetary Union (EMU) including the conditions required to be met by each member state in the European Union for joining the EMU. 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 mean shifts but a stable persistence parameter does not appear to provide an adequate description of the inflation data. This contrasts with a recent strand of the literature which argues that inflation persistence is stable once one allows for structural breaks in the mean of the series. Our results therefore question the notion that inflation persistence is an inherent feature of the euro area that should be taken into account when building dynamic models that seek to explain the behavior of inflation. The evidence presented is more in line with the view that the nature of inflation persistence varies with the transparency and credibility of the monetary regime. Our results therefore suggest that forward looking rational expectations models or hybrid models that have a substantive forward looking component are more likely to be successful at explaining inflation dynamics in the euro area as opposed to models that are characterized by purely backward looking behavior on the part of agents based on, say, an adaptive expectations mechanism.

The rest of the paper is organized as follows. Section 2 provides a brief overview of the existing empirical literature on inflation persistence in the euro area. In Section 3, we discuss tests for shifts in persistence that form the basis of our empirical analysis. Tests for single as well as multiple shifts are considered. Section 4 presents details on the construction of point and interval estimates for the break dates. Section 5 presents the empirical results. Section 6 contains a discussion of the results and some concluding remarks are given in Section 7.

2 Euro Area Inflation Persistence: A Brief Review

There is a vast and growing literature studying the nature and characteristics of inflation dynamics in the euro area. The European Central Bank 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.[2] 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 the main findings of the network as well as related work that bear relevance to the current study.

[2] More 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

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 7. 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 to conclude in favor of little change in euro area inflation persistence over the post-1970 period. Full sample estimates of the persistence parameter are close to unity and they fail to reject the hypothesis that this parameter has been stable over time. They interpret 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 appear to 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 in Europe. 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 allows us to explicitly test for shifts 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. Such an analysis is important since structural changes in price formation could well have occurred before the euro actually came into effect, and yet the latter may be responsible for it through expectations mechanisms or preparatory policies that paved the way to the euro (Angeloni et al., 2006).

3 Tests for Shifts in Persistence

In this section, we will briefly discuss the different classes of tests for shifts in persistence that we employ in our empirical analysis. The problem of testing for persistence breaks entails a choice among three different null hypotheses: the null of a stable unit root process, the null of a stable stationary process or a null that allows the process to be either stationary or have a unit root. Given that full sample estimates of the persistence parameter in inflation are typically very close to unity (see, for example, O'Reilly and Whelan, 2005 and section 5), we are specifically interested in testing the unit root null hypothesis against shifting persistence alternatives. Consequently, we include only those tests that either take the unit root model as the null hypothesis or are robust to the persistence structure under 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

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

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

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

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

Note that the data generating processes (1) and (2) allow for a shift in level to occur simultaneously with a shift in persistence. Also, for (2), the term $z_{[\lambda_0 T],1}$ is included in the $I(0)$ regime to avoid spurious jumps to zero at the break date (see Busetti and Taylor, 2004). Section 3.1 and 3.2 discuss tests for a single change in persistence, i.e., those directed against the alternatives given by (1) and (2), while section 3.3 considers tests allowing for multiple persistence shifts, a feature that is potentially relevant from an empirical standpoint, given the span of the data.

3.1 Modified Ratio-based Tests

Harvey et al. (2006) propose a set of persistence change tests that allow the data to be either $I(0)$ or $I(1)$ under the null hypothesis, i.e., the null is one of constant persistence. In particular, the same asymptotic critical values are valid regardless of whether the process is $I(0)$ or $I(1)$. The motivation for introducing such a class of tests is that a rejection can be reliably interpreted as a change in persistence. The tests are based on modified versions of ratio-based tests proposed earlier in Kim (2000) and Busetti and Taylor (2004). Consider the following ratio statistic computed for some break fraction λ:

$$K_M(\lambda) = \frac{[(1-\lambda)T]^{-2} \sum_{t=[\lambda T]+1}^{T} \left(\sum_{i=[\lambda T]+1}^{t} \hat{\varepsilon}_{1,i} \right)^2}{[\lambda T]^{-2} \sum_{t=1}^{[\lambda T]} \left(\sum_{i=1}^{t} \hat{\varepsilon}_{0,i} \right)^2} \tag{3}$$

In (3), the residual processes $\hat{\varepsilon}_0$ and $\hat{\varepsilon}_1$ are defined as $\hat{\varepsilon}_{0,t} = y_t - (1/[\lambda T]) \sum_{t=1}^{[\lambda T]} y_t$, for $t = 1, ..., [\lambda T]$ and $\hat{\varepsilon}_{1,t} = y_t - \{1/[(1-\lambda)T]\} \sum_{t=[\lambda T]+1}^{T} y_t$ for $t = [\lambda T] + 1, ..., T$. 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 test statistics are based on 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.*,

$$H_1(K_M(.)) = \max_{\lambda \in \Lambda} K_M(\lambda) \tag{4}$$

Second, Hansen's (1992) mean score statistic

$$H_2(K_M(.)) = \int_{\lambda \in \Lambda} K_M(\lambda) d\lambda \tag{5}$$

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

$$H_3(K_M(.)) = \log \left\{ \int_{\lambda \in \Lambda} \exp\left(\frac{1}{2} K_M(\lambda)\right) d\lambda \right\} \tag{6}$$

The final test statistics are defined as (for $j = 1, 2, 3$)

$$\begin{aligned} H_j^m(K_M(.)) &= \exp(-bJ_{\min}) H_j(K_M(.)) && (7) \\ H_j^m((K_M(.))^{-1}) &= \exp(-bJ_{\min}^R) H_j((K_M(.))^{-1}) && (8) \\ \max H_j^m(K) &= \exp(-b \min[J_{\min}, J_{\min}^R]) \max\left\{H_j^m(K_M(.)), H_j^m((K_M(.))^{-1})\right\} && (9) \end{aligned}$$

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 $\gamma_{k+1} = ...\gamma_9 = 0$ in the regression

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

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

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

The constant b for each of the tests in (7)-(9) is chosen so that asymptotic critical values remain the same regardless of whether the errors are $I(1)$ or $I(0)$. For each test statistic, a large value provides evidence against the null hypothesis. The tests 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. Harvey et al. (2006) show that the inconsistency of their 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.

3.2 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. 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

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

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.

3.3 Wald Tests for Multiple Persistence Shifts

A feature of the tests discussed in the previous subsections is that they are only designed to detect a one time change in persistence. Allowing for a single structural change may appear restrictive in our empirical context, given that the span of data under consideration leaves open the possibility of multiple structural changes. While the econometrics literature contains a voluminous body of work on testing for a single persistence shift, the issue of multiple shifts has received attention only relatively recently. Leybourne et al. (2007) and Kejriwal et al. (2011) propose testing procedures for the null hypothesis that the process is $I(1)$ throughout the sample against the alternative that there are possibly multiple persistence shifts, where a persistence shift between any two regimes is defined as a movement from an $I(1)$ to an $I(0)$ regime or vice-versa. In an extensive set of simulation experiments, Kejriwal et al. (2011) demonstrate that the procedure proposed by Leybourne et al. (2007) can suffer from serious over-rejection problems (size distortions) for a wide range of processes that govern the serial correlation structure of the data. On the other hand, the Wald test statistics recommended by Kejriwal et al. (2011) are shown to possess very good finite sample properties, maintaining empirical size close to the nominal size as well as rejecting the null hypothesis in a large proportion of samples when the persistence parameter is in fact unstable. Consequently, we employ a subset of the class of tests proposed by Kejriwal et al. (2011) that appear most relevant for our empirical analysis. In particular, this subset includes tests whose construction does not require the specification of the number of breaks under the alternative hypothesis, except for an upper bound, as well as a test that does not require information regarding the direction of shift, viz., whether the initial regime is $I(1)$ or $I(0)$. These tests can also be used in conjunction with each other to distinguish between processes with multiple level shifts that are otherwise stationary and multiple persistence change processes (see Section 5.4).

The general model that allows for m persistence breaks is given by

$$y_t = c_i + \alpha_i y_{t-1} + \sum_{j=1}^{p_i-1} \phi_{i,j} \Delta y_{t-j} + \epsilon_t \tag{10}$$

for $t \in [T_{i-1}+1, T_i]$, $i = 1, ..., m+1$, where we use the convention $T_0 = 0$ and $T_{m+1} = T$. The vector of break fractions is denoted $\lambda = (\lambda_1, ..., \lambda_m)$ with $\lambda_i = T_i/T$ for $i = 1, ...m$. The null hypothesis is H_0: $c_i = 0$, $\alpha_i = 1$ for all i. We consider two specifications for the alternative hypothesis depending on the nature of persistence in the initial regime. These are given by

- Model 1a: $c_i = 0$, $\alpha_i = 1$ in odd regimes and $|\alpha_i| < 1$ in even regimes.
- Model 1b: $c_i = 0$, $\alpha_i = 1$ in even regimes and $|\alpha_i| < 1$ in odd regimes.

Given an upper bound A on the number of breaks, the first two tests are based on the maximum of the individual Wald tests for the null of no break versus m breaks ($m = 1, ..., A$):

$$UDmax_{1a}(A) = \max_{1 \leq m \leq A} \sup_{\lambda \in \Lambda_h^m} F_{1a}(\lambda, m), \tag{11}$$

$$UDmax_{1b}(A) = \max_{1 \leq m \leq A} \sup_{\lambda \in \Lambda_h^m} F_{1b}(\lambda, m). \tag{12}$$

In (11) and (12), $F_{1a}(\lambda, m)$ and $F_{1b}(\lambda, m)$ denote the Wald statistics for models 1a and 1b respectively, assuming m breaks under the alternative.[3] For an arbitrary small positive number h, we define $\Lambda_h^m = \{\lambda : |\lambda_{i+1} - \lambda_i| \geq h, \lambda_1 \geq h, \lambda_m \leq 1 - h\}$.

The third test statistic is based on the presumption that the nature of persistence in the first regime is unknown, viz., we do not have any a priori knowledge regarding whether the first regime contains a unit root or not. The statistic is

$$Wmax_1 = \max_{1 \leq m \leq A} [\sup F_{1a}(\lambda, m), \sup F_{1b}(\lambda, m)]$$

Kejriwal et al. (2011) show that the tests (11) and (12) can also be used to distinguish between persistence change processes and $I(0)$ processes with multiple level shifts. In particular, if the process is characterized by level shifts but is otherwise stationary, both these tests will tend to reject the null. On the other hand, if the process involves at least one change in persistence, then only one of the two tests will tend to reject depending on the initial regime. For instance, if the initial regime is $I(1)$ but involves subsequent shifts in persistence, then $UDmax_{1a}(A)$ would be likely to detect such a process while $UDmax_{1b}(A)$ may not be expected to provide evidence against the unit root null. In our empirical analysis, we set $A = 5$.

[3] In the construction of the Wald statistics, the coefficients $\phi_{i,j}$ are held fixed across regimes in order to direct the power of the tests against changes in persistence as opposed to changes in short-run dynamics.

4 Break Date Estimation: Point Estimates and Confidence Intervals

Following evidence against the null hypothesis of no persistence change, it is desirable to estimate the break dates and form confidence intervals for them. Given that our empirical analysis does not find any support for multiple breaks, we will accordingly focus on methods for estimating a single change point. Based on the residual processes $\hat{\varepsilon}_1$ and $\hat{\varepsilon}_0$ as defined in (3), Busetti and Taylor (2004) and Kim et al. (2002) independently propose the following estimator for the breakpoint:

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

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.

Leybourne et al. (2006) propose alternative consistent breakpoint estimators that differ depending on the direction of shift. 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

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

In addition to point estimates, it is important to provide a measure of precision associated with these estimates. To the best of our knowledge, the only study that considers construction of confidence intervals for persistence change processes is Chong (2001). However, the modeling framework he considers is overly restrictive being based on the assumption of an AR(1) model and a zero intercept in each regime. Hence, we construct confidence intervals for the breakpoint based on a sieve bootstrap procedure (see Bühlmann, 1997). The procedure is implemented as follows. Given the estimated break date from one of the procedures described above, we estimate an AR(p) model for each regime (where p is selected by BIC) and obtain the vector of residuals $\hat{\epsilon}$. To generate the bootstrap samples, we resample $\hat{\epsilon}$ to obtain the bootstrap error vector ϵ^*. The bootstrap samples are then generated according to

$$y_t^* = \begin{cases} \hat{c}_1 + \hat{\alpha}_1 y_{t-1}^* + \sum_{j=1}^{p_1-1} \hat{\phi}_{1,j} \Delta y_{t-j}^* + \epsilon_t^* & \text{if} \quad t <= \left[T\hat{\lambda}\right] \\ \hat{c}_2 + \hat{\alpha}_2 y_{t-1}^* + \sum_{j=1}^{p_2-1} \hat{\phi}_{2,j} \Delta y_{t-j}^* + \epsilon_t^* & \text{if} \quad \text{otherwise} \end{cases}$$

where $\hat{\lambda}$ is a general breakpoint estimator that takes the values $\hat{\lambda}_M$, $\hat{\lambda}_{01}$ or $\hat{\lambda}_{10}$ depending on which estimation procedure is employed. For the initial values, we set $y_t^* = \bar{y}_1$ for $t \leq 0$, where $\bar{y}_1$ denotes the mean of the data in the first regime. The first 50 observations are discarded in order to mitigate startup effects.

From each bootstrap sample, we compute the estimate λ^* using one of the methods above. Then the bootstrap distribution of $T_1^* - \hat{T}_1$ is used to approximate the unknown finite sample distribution of $\hat{T}_1 - T_1^0$, where $T_1^* = [T\lambda^*]$, $\hat{T}_1 = [T\hat{\lambda}]$ and $T_1^0 = [T\lambda^0]$, where λ^0 denotes the true break fraction. Note that we approximate the latter distribution since, as shown in Kim et al. (2002) and Leybourne et al. (2006), $\hat{T}_1 - T_1^0$ has a nondegenerate asymptotic distribution. Finally, the $100(1-\alpha)\%$ confidence intervals are obtained as $[\hat{T}_1 - cv_{1-\alpha/2}^*, \hat{T}_1 - cv_{\alpha/2}^*]$, where $cv_{\alpha/2}^*$ and $cv_{1-\alpha/2}^*$ denote the $\alpha/2$ and $(1-\alpha/2)$ quantiles of the bootstrap distribution of $T_1^* - \hat{T}_1$, respectively. In our empirical analysis, we report three such intervals for each of the two breakpoint estimation methods. The first interval uses the usual OLS estimates to generate the bootstrap samples and is denoted $BootCI_{ols}$, the second is based on the bias-corrected estimates suggested by Andrews and Chen (1994) and denoted $BootCI_{ac}$, while the third is based on estimates that impose the unit root restriction in the relevant regime and is denoted $BootCI_r$. The second interval accounts for the potential bias associated with least squares estimates in autoregressive models whereas the motivation for considering the last interval derives from the fact that substantial finite sample efficiency gains may be obtainable by imposing the relevant parameter restrictions, as shown in a related context by Perron and Qu (2006).

5 Empirical Results

The data for our empirical analysis are obtained from the European Central Bank's Area Wide Model (AWM) quarterly database described in Fagan, Henry and Mestre (2001). The sample period is 1970Q1-2009Q4. The measure of inflation (y_t) used is the annualized quarterly log-difference of the GDP deflator.[4] 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 persistence as opposed to pure shifts in level while retaining the same degree of persistence throughout the sample.

The empirical results of our analysis are organized in four subsections. In subsection 4.1, we present the results of persistence change tests described in Section 3. Subsection 4.2 reports the point and interval estimates for the breakpoint as well as the full sample and regime-specific persistence parameters (the sum of the autoregressive coefficients) computed both from ordinary least squares regressions and 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 presents arguments to show that the hypothesis of stationarity with mean shifts but a stable persistence parameter does not appear to provide an adequate description of the euro area inflation data.

5.1 Persistence Change Tests

Table 1 reports the results of the persistence change tests described in section 3 applied to euro area inflation. Given the variety of tests employed, we also present in Table 1 the corresponding null and alternative hypotheses for each of the tests.[5]We will argue that the empirical results suggest an $I(1)$-$I(0)$ process as a suitable approximation to the data generating process for the inflation data. First, the modified ratio-based tests that are directed against the $I(0)$-$I(1)$ alternative do not provide any statistical evidence against the null of stable persistence. In contrast, the tests based on $K_M(.)^{-1}$ designed to detect

[4] An 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 empiricial analysis is primarily based on tests for parameter stability, we focus on the GDP deflator measure in this paper.

[5] As in O'Reilly and Whelan (2005), we also computed 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. None of these tests were found to be significant at even the 10% level. The results are available upon request.

the $I(1)$-$I(0)$ alternative are all able to reject the null at even the 1% level. As discussed in Harvey et al. (2006) and Section 3, the modified tests directed against the "incorrect alternative" have low power thereby suggesting that the data are possibly consistent with a shift from (near) unit root behavior to stationarity rather than vice-versa. The maximum of the modified ratio-based tests are also all significant at conventional levels. Second, the R test rejects in the right tail which, following the reasoning in Section 3, is again supportive of an $I(1)$-$I(0)$ type 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, none of the multiple shift tests are significant at even the 10% level.

Given that the above test results appear to be suggestive of an $I(1)$-$I(0)$ type behavior for inflation, we now present point and interval estimates for the break date as well as the persistence parameters over the regimes identified by the estimated break date.

5.2 Parameter Estimates

The breakpoint estimates $\hat{\lambda}_M$ and $\hat{\lambda}_{10}$ and the bootstrap confidence intervals are reported in Table 2. The break date identified by both methods is the second quarter of 1993. As discussed in section 6, such a date may be expected based on economic events that may have contributed to a reduction in the level of inflation persistence. The confidence intervals based on the estimate $\hat{\lambda}_M$ appear relatively long and therefore are not particularly informative while those based on $\hat{\lambda}_{10}$ are much shorter and cover the period from the late 80's to the early 90's. Moreover, the three different methods for generating the bootstrap samples yield very similar confidence intervals.

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 BIC (assuming a maximum of 10 lags). The model for regime i is

$$y_t = c_i + \alpha_i y_{t-1} + \sum_{j=1}^{p_i - 1} \phi_{i,j} \Delta y_{t-j} + \epsilon_t \tag{13}$$

In (13), α_i $(i = 1, 2)$ denote the sum of the autoregressive coefficients which is a typical measure of persistence in autoregressive models. Given that the normal approximation to the sampling distribution of the persistence parameter estimates can be quite poor for values of α_i close to unity, we also report confidence intervals based on Hansen's grid bootstrap procedure which has been shown to work well globally in the parameter space.[6] Further, it is well known that the ordinary 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). The results are reported in Table 3. The full sample OLS estimate of the persistence parameter is very close to unity with a relatively tight confidence interval, consistent with the analysis in O'Reilly and Whelan (2005). But the least squares estimates are seen to differ markedly across the two regimes. In the first regime, the persistence parameter is estimated at 0.93 while the post-break estimate is 0.56, far smaller. This is clearly indicative of a substantial reduction in persistence after the break.[7] The bias-adjusted estimate in each regime (denoted $\hat{\alpha}_{ac}$ in Table 3) is larger than its OLS counterpart for both regimes but the difference between the estimates in the two regimes is still substantive. 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 intervals are somewhat wider although they are consistent with a range of low persistence parameter values in contrast to the intervals for the first regime. The main message from these results is, however, clear: there appears to have been an important decline in inflation persistence in the period following the break.

5.3 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 (13), it is given by $\mathrm{Var}(u_t)/(1-\alpha_i)^2$. 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

[6] We present results for the so called grid-α bootstrap procedure based on 399 bootstrap replications. As argued in Rossi (2005), an alternative grid-t procedure also suggested in Hansen (1999) can lead to poor coverage rates for half life confidence intervals.

[7] These estimates were computed using 1000 Monte Carlo replications.

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 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 LAR estimates, denoted $\hat{\rho}_i$, the median unbiased estimates, denoted $\hat{\rho}_{1,med}$ and $\hat{\rho}_{2,med}$, and the 95% confidence intervals (in brackets beside the point estimates $\hat{\rho}_i$) for ρ_1 and ρ_2, are presented in Table 3. The point estimate in the first regime and for the full sample are both very close to unity while for the second regime, the LAR estimate is only 0.67. The median unbiased estimates are only slightly higher in each regime. The confidence interval for the first regime as well as for the full sample are tightly concentrated around unity while for the second regime it is wider and includes unity, although it is also compatible with much lower values for the LAR. Overall, 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 half life estimate 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 (13). Based on the confidence interval for ρ_i, we can then directly get, by monotonicity, the corresponding 95% confidence interval for the half life. As in Rossi (2005), we report confidence intervals based on the methods proposed by Stock (1991), Hansen (1999) and Elliott and Stock (2001). The median half life estimates and the three confidence intervals denoted $CI_S(h)$, $CI_H(h)$ and $CI_{ES}(h)$ respectively, are presented in the last four rows of Table 3. The median half life estimate is infinity for the full sample and for the first regime given that the corresponding median unbiased estimates exceed unity. For the second regime, the median estimate is slightly less than 2 quarters suggesting that the effects of shocks possibly dissipate quite rapidly in this regime. The half life confidence intervals based on all three methods are, however, rather wide and do not rule out the possibility of an infinite half life.

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 the true 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.

5.4 Stationarity with Mean Shifts

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 occasional shifts in its mean but whose persistence parameter remains stable. They argue that ignoring mean breaks leads to overestimating the extent of inflation persistence and that once one controls for such breaks, measured persistence is much lower. Indeed, the annualized inflation rate is 7.51% over 1970Q1-1993Q2 and 1.93% over 1993Q3-2009Q4. Moreover, as argued in Belaire-Franch (2005) and Busetti and Taylor (2004), persistence change tests 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. However, for such a process, tests directed against *both* the $I(1)$-$I(0)$ alternative as well the $I(0)$-$I(1)$ alternative are likely to reject the null of stable persistence. In contrast, for processes involving a change in persistence, only one of the two tests may be expected to provide evidence against the null, as discussed in Section 3. Given the pattern of rejections reported in Table 1, it is clear that the data are likely to be consistent with the latter type of processes, viz., one that undergoes a $I(1)$-$I(0)$ change in persistence accompanied by a shift in mean. For processes with multiple level shifts that are otherwise stationary, all three tests considered in Section

3.3 would be expected to reject, as shown in Kejriwal et al. (2011). Indeed, none of these tests turned out to be statistically significant at conventional levels thereby providing no evidence in favor of a stationary process with multiple level shifts. In what follows, we employ an alternative methodology that is intended to reinforce the conclusion 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. Further, the results obtained in this section are argued to be consistent with a data generating process that is non-stationary for a non-negligible fraction of the sample but stationary otherwise.[8]

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

$$y_t = \mu + \delta DU_t + \widetilde{y}_t, \quad t = 1, ..., T \tag{14}$$

where $DU_t = I(t > T_b)$. We then construct the t-statistic, denoted $t_\alpha(\mathrm{AO}, T_b, k)$, for testing $\alpha = 1$ in the regression

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

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

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

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

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

[8]This methodology in this section is restricted to a single break date given that most of the procedures employed in this section have not been extended to account for the possibility of multiple shifts.

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 $PV_2(i),\ i =$ AO, 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 $k_{\max} = 10$.

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 (14) 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 $\gamma_1 = \gamma_2 = ... = \gamma_9 = 0$. Then, since the shift date is unknown, Vogelsang (1998) considers the following three functionals:

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

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 $\widetilde{y}_t^*$. 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.

Busetti and Harvey (2001) and 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

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

where

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

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

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

with

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

The parameter m is chosen according to the data dependent method suggested in Kurozumi (2002):

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

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 4. First, the unit root tests do not provide any evidence against the unit root null for both the AO and IO models. Note that a non-rejection by these tests could be expected for data generating processes that are characterized by a unit root for a fraction of the sample, given that the behavior of the unit root component dominates such tests (see Kim, 2003 for a discussion on the low power of unit root tests in this context). 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. Again, these rejections are consistent with the presence of a unit root in a subsample of the data given that the unit root dominates the behavior of such tests. To summarize, our 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. The pattern of rejections/non-rejections obtained from the various testing procedures seems to be more compatible with persistence change data generating processes of the form considered in this paper.

6 Discussion

The empirical results presented in the foregoing section provide clear evidence against a stable reduced form representation for euro area inflation. Consequently, the usefulness of such a specification for forecasting or policy analyses is likely to be very limited. Moreover, the impulse response and propagation mechanisms derived from such a reduced form are likely to be misleading. The results also question the adequacy of the simple backward looking rule-of-thumb model of expectations in which agents extrapolate from past inflation rates to formulate the expectation used in current period wage and price setting. Rather, in the light of the Lucas Critique, a forward looking component based on rational expectations is likely to be important in explaining the dynamics of inflation in the euro area if the identified structural change can be linked to a major change in the monetary policy regime.

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.

Consistent with our empirical results, recent theoretical 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.

It is important to note that the results of our univariate reduced form analysis do not distinguish between the different sources of inflation persistence and should rather be interpreted as providing an overall statistical measure of persistence as in Pivetta and Reis (2007). Altissimo et al. (2006) discuss three sources of persistence in inflation: (a) persistence that is inherited from fluctuations in the determinants of inflation such as marginal costs or the output gap ("extrinsic persistence"), (b) persistence that emanates from the dependence of inflation on its own past ("intrinsic persistence") and (c) persistence arising due to the formation of inflation expectations ("expectations-based persistence"). These sources interact with each other and their relative importance will depend on the monetary regime and the policy reaction function. The creation of EMU may in principle have affected all three sources of persistence. First, the establishment of EMU has led to increased competition, not only because the single numeraire across the area facilitates systematic price comparisons across countries, but also because EMU may have been a triggering factor for product market reforms aimed at strengthening competition. This increased competition is likely to have led to a stronger incentive to set prices in an optimal, forward looking manner and reduce the extent of explicit or implicit indexation, thereby reducing intrinsic inflation persistence. Second, as discussed earlier, the European Central Bank's clear mandate for maintaining price stability has presumably resulted in a stronger achorage of inflation expectations and a lower dependence on past inflation. This might have reduced the persistence due to expectations. Third, the EMU may have affected the monetary reaction function and the transmission mechanism of monetary policy which could potentially have had an impact on extrinsic inflation persistence (see Altissimo et al., 2006 and Angeloni et al., 2006 for a more detailed discussion of the different sources of inflation persistence).

7 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 the finding of a stable persistence parameter in the extant literature 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 a wide range of 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 inherent characteristic of the economy that should necessarily be incorporated into the structure of general equilibrium macroeconomic models. 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.

The findings of this 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.[9] 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

[9] Altissimo 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.

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.

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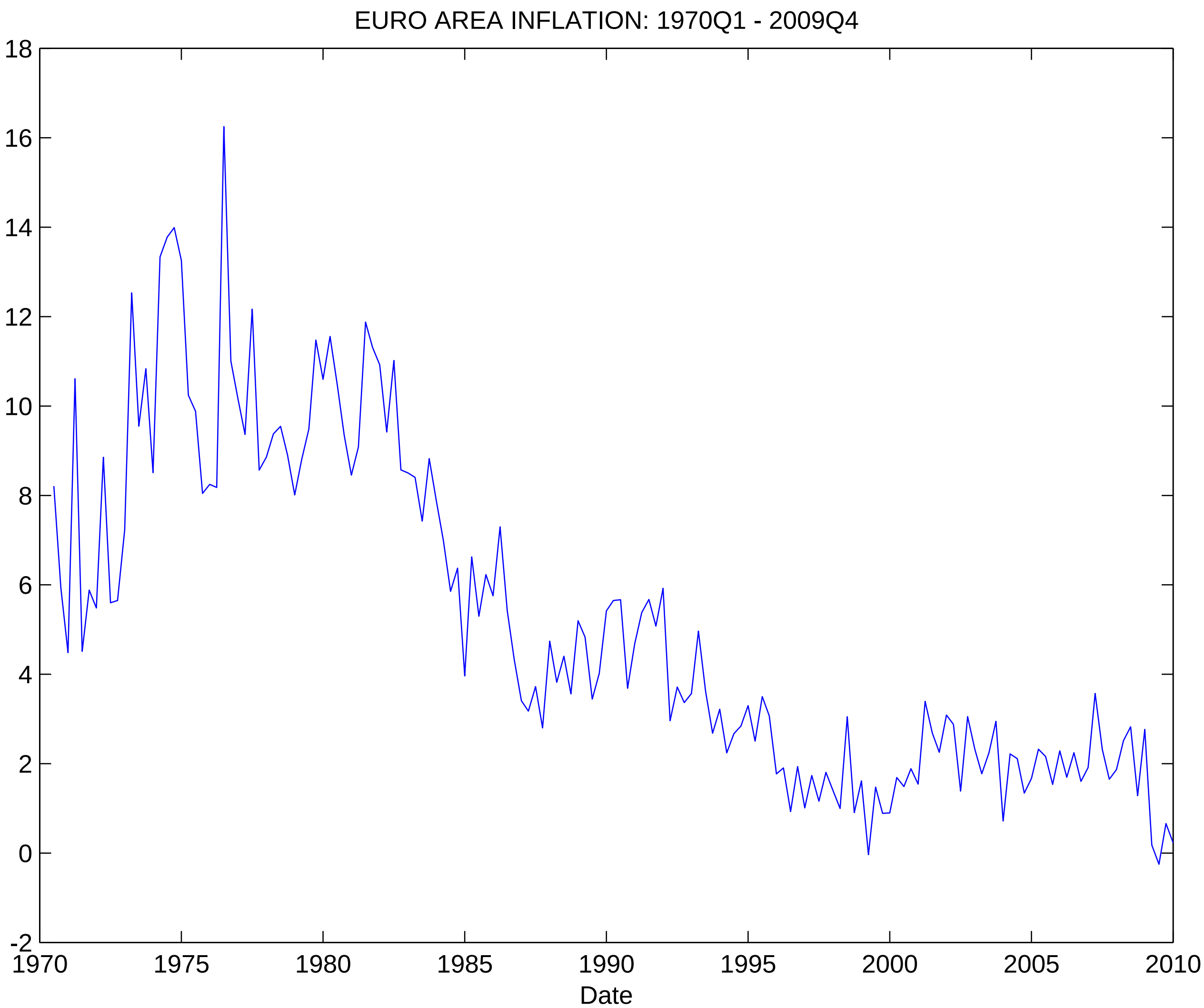
EURO AREA INFLATION: 1970Q1 - 2009Q4
18
16
14
12
10
8
6
4
2
0
-2
1970
1975
1980
1985
1990
1995
2000
2005
2010
Date

Table 1: Tests for Shifts in Persistence

Source	Test Statistic	H_1	Sample Value
(A) $\mathbf{H}_0$: Stable $\mathbf{I(1)/I(0)}$			
Harvey et al. ('06)	$H_1^m(K_M(.))$	$I(0)$-$I(1)$	(8.85,7.24,3.86)
	$H_2^m(K_M(.))$	$I(0)$-$I(1)$	(2.05,1.75,1.04)
	$H_3^m(K_M(.))$	$I(0)$-$I(1)$	(2.39,1.93,1.01)
	$H_1^m((K_M(.))^{-1})$	$I(1)$-$I(0)$	(280.41*,233.74**,129.12***)
	$H_2^m((K_M(.))^{-1})$	$I(1)$-$I(0)$	(80.98*,69.61**,43.81***)
	$H_3^m((K_M(.))^{-1})$	$I(1)$-$I(0)$	(125.03*,103.02**,55.81***)
	$\max H_1^m(K)$	$I(1)$-$I(0)/I(0)$-$I(1)$	(234.03*,188.98**,93.19***)
	$\max H_2^m(K)$	$I(1)$-$I(0)/I(0)$-$I(1)$	(69.66*,58.35**,33.14***)
	$\max H_3^m(K)$	$I(1)$-$I(0)/I(0)$-$I(1)$	(103.06*,82.38**,40.09***)
Leybourne et al. ('06)	R	$I(1)$-$I(0)/I(0)$-$I(1)$	8.51***
(B) $\mathbf{H}_0$: Stable $\mathbf{I(1)}$			
Leybourne et al. ('06)	N	$I(0)$-$I(1)$	0.10
	D	$I(1)$-$I(0)$	0.01**
	M	$I(1)$-$I(0)/I(0)$-$I(1)$	0.01**
Kejriwal et al. ('11)	$UDmax_{1a}$	$I(1)$ Initial	4.23
	$UDmax_{1b}$	$I(0)$ Initial	5.30
	$Wmax_1$	$I(1)$ Initial/$I(0)$ Initial	5.30

Table 2: Break Date Estimates and Bootstrap Confidence Intervals

Estimator	Estimate	$BootCI_{ols}$	$BootCI_{ac}$	$BootCI_r$
$\hat{\lambda}_M$	1993Q2	[1985Q3,1993Q3]	[1985Q3,1994Q3]	[1985Q4,1994Q3]
$\hat{\lambda}_{10}$	1993Q2	[1989Q1,1994Q2]	[1988Q2,1994Q4]	[1989Q2,1994Q3]

Table 3: Parameter Estimates and Measures of Persistence

Estimator/Measure	Full Sample	Regime 1	Regime 2
$\hat{\alpha}_{ols}$	0.97; [0.91,1.03]	0.93; [0.80,1.03]	0.56; [0.28,0.83]
$\hat{\alpha}_{ac}$	0.99; [0.95,1.00]	0.99; [0.87,1.00]	0.63; [0.32,0.91]
$BootCI$	[0.94,1.06]	[0.86,1.10]	[0.34,0.95]
p_{bic}	4	4	2
$\hat{\rho}$	0.98; [0.97,1.03]	0.96; [0.93,1.05]	0.67; [0.57,1.01]
$\hat{\rho}_{med}$	1.00	1.00	0.77
$\hat{h}_{med}$	∞	∞	1.75
$CI_S(h)$	[0.16,∞)	[0.87,∞)	[1.02,∞)
$CI_H(h)$	[0.15,∞)	[0.77,∞)	[0.88,∞)
$CI_{ES}(h)$	[22.49,∞)	[5.66,∞)	[4.24,∞)

Table 4: Tests for Stationarity with a Mean Shift

Test	Sample Value
$PV(AO1)$	-3.62
$PV(AO2)$	-1.60
$PV(IO1)$	-3.94
$PV(IO2)$	-0.87
$\sup PS_T$	(17.69*,13.82**,5.92)
mean PS_T	(4.57*,4.07**,2.83)
exp PS_T	(5.64*,4.23**,1.70)
S_T	0.38***
S_T^{ps}	1.06**

Note: In Tables 1 and 4, '*', '**', 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.