Inflation Persistence in OECD and non-OECD economies

1. Introduction

Inflation is an important macroeconomic variable that affects the options of economic agents as well as their future expectations and it is often regarded as a result of domestic policies combined with the effects of globalization and therefore, a sign of how governments have been well succeeded in their political options. This issue is reflected in the mandate of many monetary authorities to maintain price stability and, therefore, no wonder it plays a critical role in policies' design as its effects spread out in the economy as a whole either in terms of economic efficiency and equity, two of the most important concerns of any government's policy.

This explains the attention political authorities and economic agents in general have given to the evolution of inflation and the fact that its control has been stated as a priority for governments all over the world. These issues have became increasing relevant as the international monetary context has experienced important changes such as the adoption of inflation targeting regimes by some countries, the arrival of monetary union in Europe, and a general deflationist process in industrial economies.

At the center of this concern is the evolution of inflation persistence, considered as a key property of inflation, being usually defined as the duration of the effects of shocks hitting inflation. (For other definitions of persistence, see, *inter alia*, Andrews and Chen, 1994; Pivetta and Reis, 2007; Batini and Nelson, 2002; Batini, 2002; and Willis, 2003). The analysis of inflation persistence has been widely studied in the literature by a large number of theoretical and empirical papers, but no consensus has been achieved yet. While the former group has contributed with advances in the statistical treatment of time series data and improved the tools of analysis (see, *iter alia*, Cat *et al*, 1999; Andrews and Chen, 1994; Pivetta and Reis, 2007; Kim, 2000; Kim *et al*., 2002; and Busetti and Taylor, 2004). The last group has presented mixed evidence about changes in inflation persistence. Recent empirical evidence not only suggests that inflation has varied over time but also that inflation is not an intrinsically persistent process (see inter alia Barsky, 1987; Evans and Wachtel, 1993; Brainard and Perry, 2000; Taylor, 2000; Kim et al., 2001; Cogley and Sargent, 2001; Ravenna, 2000;

Benati, 2003; Levin and Piger, 2004; Harvey *et al*, 2006). This contrasts with the findings of O'Reilly and Whelan (2004), Gadea and Mayoral (2006), Marques (2004) and Piveta and Reis (2006), who report evidence of unchanging inflation persistence. In this line, some studies also report a discrepancy between short and long samples-based measures of inflation persistence (Levin and Piger, 2004; Benati, 2003, 2004; Benati and Wood, 2004, Altíssimo, 2003, Gadzinski and Orlandi, 2004; Goodhart and Hofmann, 2003; Cecchetti and Debelle, 2004, Marques, 2004 and Dias and Marques, 2004).

Therefore, one possible explanation for the inconclusive nature of the results reported in the literature is it has ignored the occurrence of structural changes that have caused breaks in the mean of inflations.

Paradoxically, maybe because of the inconclusive nature of the results on the analysis of inflation persistence, the issue of the possible existence of structural breaks in the mean of inflation has received little attention. This issue is addressed directly by Corvoisier and Mojon (2005) in measuring the effects M3 growth and real unit labor costs on inflation, and Marques (2004) which that the U.S. inflation shows persistence parameter instability when using conventional structural break tests.

This paper is a contribution to the analysis of inflation persistence, focused on how sluggishly inflation returns to its long-run equilibrium level after an exogenous shock. In particular, our objective is to test whether the impact of shocks to inflation will be transitory with inflation presenting a mean-reverting behaviour, or persistent with long lasting effects. While this is a central key to assess the short term impact of monetary policy decisions and to determine the short-run trade-off between inflation and real activity, it plays no minor role in our capability in forecasting inflation and selecting structural models. Successful out-of-sample forecasting of inflation requires in first instance to decide the appropriate degree of persistence and, eventually, of non-stationarity of the data generating process. This affects not only the point forecasts but, perhaps more importantly, impulse responses and predictive density forecasts.

We depart from the existing literature in two important ways. First, and methodological speaking, we depart from the recently proposed tests of persistence changed proposed by Harvey *et al*. (2006). We adopt the evidence suggested in Corvoisier and Mojon (2005) and Marques (2004) and propose modified tests to account for the existence of a structural break in the mean of inflation. Second, we empirically analyse the inflation persistence in a total of 30 countries with economic and structural differences and which are classified in three categories whether they are European Union (EU) members, OECD members and non-OECD members. By using country specific data it is possible to assess the extent to which the shocks affecting it are themselves persistent, or diverse structural factors in the economy, such as the presence of fiscal dominance, the formation of inflation expectations, the exchange rate regime in place, the degree of price indexation and the monetary regime followed by the central bank. These two directions may prove to be a useful complement to identify different patterns of inflation and achieve some consensus in the literature.

The paper is organized as follows. Section 2 presents the data and some preliminary empirical results. Section 3 presents the econometric tests for changes in persistence. Section 4 reports the main results on inflation persistence and final conclusions.

2. Data and Preliminary Empirical Results

2.1 Data: sources and description

The data in this study consist of monthly observations of the inflation rate, measured as the monthly percentage change in consumer prices index (CPI) from 1980:01 to 2009:12 for a set of OECD countries non-OECD economies in a total of 30 countries.

As OECD countries, we consider Austria (AT), Belgium (BE), Canada (CA), Denmark (DN), Finland (FI), France (FR), Germany (GE), Greece (GR), Hungary (HG), Iceland (IC), Ireland (IR), Italy (IT), Japan (JP), Korea (KO), Luxemburg (LU), Mexico (ME), the Netherlands (NH), Norway (NW), Portugal (PO), Spain (SP), Sweden (SE), Switzerland (SW), Turkey (TU), the United Kingdom (UK) and the United States (US). For the non-OECD countries group, we consider Chile (CL), India (IN), Indonesia (ID), Israel (IS) and South Africa (SA). The source is the OECD database, which is available on-line, and the choice of the countries was determined by data availability.

The aggregated data for both the OECD and non-OECD countries are presented in Figures 1 and 2.

Figure 1: Inflation in the OECD countries

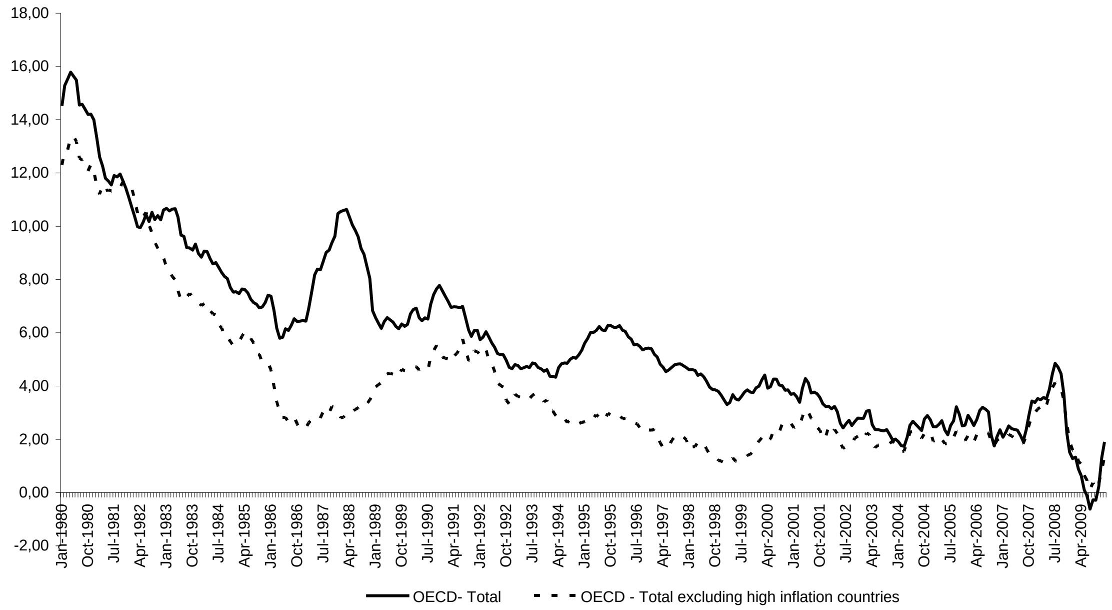
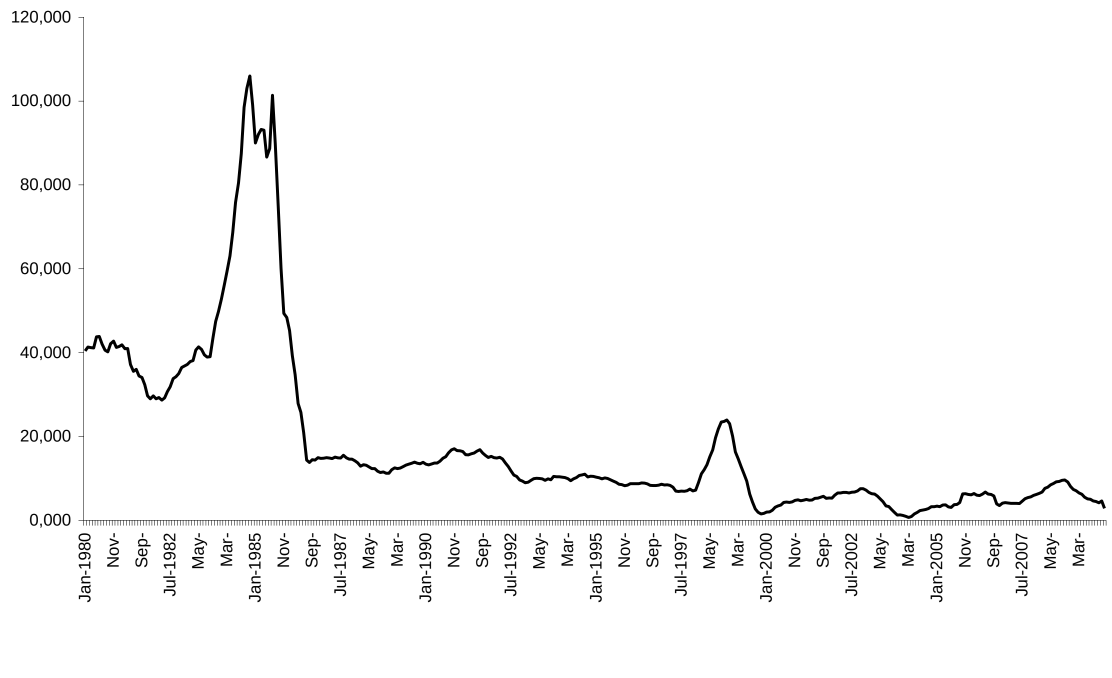


Figure 2: Inflation in a sample of non-OECD countries



A visual inspection shows the well known trends of inflation. Starting from relatively high levels in the beginning of the 1980s, inflation initiate a reduction movement but rose dramatically in the mid of the decade. However, it is evident the results of the battle towards inflation reduction embarked by both OECD and non-OECD countries embarked in the second half of the 80s. Considering the inflation levels in the beginning of the year for comparison purposes, in the OECD member States, inflation went from an average of 14.51% in 1980, to 6.71% 1990, 3.99% in 2000 and 1.27% in 2009. This decreasing trend was also observed in traditionally high-inflation countries, as illustrated in Figure 1. The non-OECD countries also started a downsizing trajectory in monthly inflation. After having reached a maximum value of 40.39% in 1980, the value was reduced to 13.38% in 1990, and 1.94% in 2000. However, in the last decade inflation has been impacted by upward forces in both groups.

The evolution in the last decades highlights the high heterogeneity in the levels of the series, but less than in the past, between OECD and non-OECD countries. The monthly average inflation rate in the sample period was 5.78% and 17.96% in the OECD- the non-OECD economies respectively. Differences are also observed among the OECD countries. The average monthly inflation rate was 6.81%, but this value is reduced to 3.98% when the high inflation countries are excluded. While these statistics illustrate the disparity of the inflation behaviour between groups of the countries, a country by country detailed analysis explicitly uncovers any individual differences.

Table 1 reports some statistic data by country, which highlights some relevant issues. First, the discrepancy of inflation levels is now clearer within and outside the OECD group. Second, most series present high volatility. Third, the variance and the level of inflation seem to be positively related, and is clear the so-called “great moderation” reflected by the very low mean and standard deviation in several countries during the last ten years of the sample period.

As a first look at inflation persistence we analyze the evolution over time of the sample first-order serial correlation. Almost all individual estimates are marginally lower that unity over the sample period although this value seems to have declined in most countries over the last decades. However, the magnitude of this decrease was only marginally different from zero.

Table 1: Descriptive statistics in OECD economies

	Sample period			1980s			1990s			2000s		
	Mean	**St. Dev.**	**Corr**	**Mean**	**St.Dev.**	**Corr**	**Mean**	**St.Dev.**	**Mean**	**St.Dev.**	**SD**	**Corr**
Austria	3,49	2,17	0,96	3,84	2,00	0,98	2,41	1,19	0,97	1,99	0,87	0,92
Belgium	3,67	2,89	0,99	4,90	2,84	0,99	2,15	0,88	0,93	2,12	1,27	0,93
Canada	3,88	3,29	0,99	6,51	3,18	0,98	2,20	1,68	0,92	2,12	1,01	0,87
Denmark	5,20	3,86	0,98	6,92	3,25	0,97	2,11	0,49	0,80	2,11	0,75	0,92
Finland	5,45	4,45	0,99	7,19	2,98	0,98	2,19	1,80	0,95	1,70	1,36	0,95
France	4,88	3,99	0,99	7,39	4,32	0,99	1,88	0,93	0,96	1,73	0,77	0,91
Germany	2,80	1,80	0,98	2,90	2,16	0,99	2,56	1,56	0,97	1,59	0,73	0,90
Greece	9,09	8,06	0,99	19,49	4,12	0,96	11,13	6,08	0,98	3,16	0,89	0,89
Hungary	12,59	8,75	0,99	9,01	4,41	0,94	22,26	7,25	0,98	6,14	2,33	0,96
Iceland	19,21	20,80	0,99	39,73	21,48	0,98	4,29	4,64	0,91	6,25	3,91	0,98
Ireland	5,95	5,81	0,98	9,35	6,73	0,98	2,31	0,81	0,89	3,16	2,89	0,85
Italy	6,37	5,63	1,00	11,22	5,74	0,98	4,16	1,72	0,98	2,27	0,72	0,95
Japan	3,37	4,29	0,98	2,53	2,27	0,96	1,21	1,31	0,93	-0,26	0,84	0,93
Korea	12,50	16,86	0,93	8,43	8,90	0,98	5,74	2,44	0,95	3,12	0,98	0,87
Luxembourg	3,44	2,75	0,98	4,80	3,42	0,99	2,17	1,07	0,92	2,34	0,93	0,91
Mexico	27,24	32,92	0,99	69,88	41,77	0,98	20,51	10,47	0,98	5,22	1,75	0,93
Netherlands	3,62	2,68	0,98	2,87	2,63	0,99	2,44	0,55	0,91	2,13	1,00	0,96
Norway	4,84	3,36	0,98	8,35	2,91	0,98	2,45	0,87	0,92	2,09	1,31	0,88
Portugal	9,18	8,74	0,98	17,70	7,14	0,97	6,02	3,78	0,98	2,60	1,40	0,96
Spain	7,65	5,56	0,99	10,26	3,83	0,97	4,22	1,70	0,98	2,96	1,31	0,95
Sweden	4,91	3,65	0,98	7,94	3,06	0,97	3,30	3,68	0,97	1,50	1,22	0,94
Switzerland	2,80	2,36	0,98	3,27	1,81	0,96	2,34	2,09	0,98	0,96	0,83	0,91
Turkey	35,06	29,78	0,99	49,49	21,48	0,96	77,42	16,43	0,95	23,09	21,00	0,97
U. Kingdom	5,49	5,11	0,99	7,11	4,55	0,97	3,31	2,23	0,98	1,85	0,90	0,93
U.S.	3,97	2,88	0,99	5,56	3,53	0,97	3,00	1,12	0,96	2,57	1,44	0,92
Chile	55,00	123,04	0,99	21,48	8,30	0,96	11,79	7,25	0,98	3,49	2,40	0,96
India	7,45	5,70	0,98	9,22	2,56	0,93	9,57	3,53	0,90	5,45	2,38	0,90
Indonesia	12,60	11,78	0,97	9,66	4,10	0,92	14,87	18,24	0,98	8,48	4,13	0.92
Israel	45,40	82,59	0,99	132,17	124,90	0,99	11,26	4,56	0,96	2,01	2,32	0,96
S. Africa	8,23	5,14	0,99	14,61	2,28	0,91	9,91	3,62	0,95	5,45	3,71	0,98

2.2 Preliminary empirical results

In the light of the above discussion, it is relevant to be able to identify whether inflation follows a stationary process allowing for the possibility of changes in the degree of persistence, or a non-stationary process, especially when changes in the structure of the economy or in the monetary policy framework have taken place. This distinction is meaningfully for the purposes of our analysis since it helps understanding the effects of shocks to inflation: while the effects of such shock will be transitory for a stationary series, they will be permanent in the case of non-stationary series. In other words, while an I(0) time series will display mean-reverting behavior, an I(1) variable will be persistent, that is, shock to it will have long lasting effects.

In face of the previous analysis which underlines the possible existence of non-stationarities characterized by variable means and variances as well as significant autocorrelation coefficients, we carry out the testing procedure in two steps. First, we perform unit roots tests to establish the apparent order of integration of the series and, second, we test for the existence of breaks in the mean which may be due to changes in monetary policy, as they have been expected to occur in many countries.

The results of the unit root tests are reported in Table 2. We applied the Augmented Dickey Fuller (ADF) tests (Dickey and Fuller ???), the ADF-ERS test (Elliott, Rothenberg and Stock, 1996), a modification of the augmented Dickey-Fuller test that has substantially improved power when an unknown mean or trend is present, the Phillips-Perron (PP) test (Phillips and Perron ????), as well as the MZ_α, MZ_t, MSB and the MPT tests due to Ng and Perron (2001), designed to overcome both size distortion and low power problems when data is characterized by large absolute values autoregressive and moving average roots. In performing the tests, a constant was included as well as a linear trend when significant. As can be seen, there is an absolute concordance among the different tests in the sense that it is not possible to reject the null hypothesis of a unit root for inflation in almost all countries, confirming other results in the literature according to which the CPI is a second order integrated (I(2)) variable. For Chile, India and Indonesia, results seem to reject the null hypothesis of a unit root. These results are in accordance with the analysis of the previous section from which some evidence of degree of nonstationarity for most countries was drawn.

Table 2: Unit root tests (to be changed with result of the period 1980:01-2009:12)

Countries	ADF	ADF-GLS	PP	MZ_{α}	MZ_t	MSB	MPT
Austria	1.36	1.17	-3.09	-3.44	-1.20	0.35	24.57
Belgium	-1.88	-1.86	-2.03	-9.21	-2.10	-0.23	10.07
Canada	-5.17	-1.13	-3.13	-2.97	-1.14	0.38	28.6
Denmark	-2.31	-1.84	-3.55**	-5.92	-1.71	0.29	15.39
Finland	-2.42	-2.37	-3.00	-10.69	-2.28	0.21	8.66
France	-3.03	-1.67	-3.06	-6.07	-1.71	0.28	14.99
Germany	-2.28	-1.79*	-2.80*	-6.84	-1.85	0.27	3.59
Greece	-1.59	-1.58	-2.15	-5.38	-1.59	0.29	16.78
Hungary	-1.80	-1.28	-1.81	-3.59	-1.29	0.36	24.64
Iceland	-1.98	-1.84	-2.77	-7.25	-1.88	0.26	12.62
Ireland	-2.20	-1.97	-2.22	-8.66	-2.07	0.24	10.58
Italy	-1.55	-1.06	-2.11	-2.47	-1.04	0.42	33.91
Japan	-2.94	-1.75	-3.64	-6.29	-1.72	0.27	14.48
Korea	-2.93	-1.22	-7.31***	-0.47	-0.36	0.77	116.05
Luxembourg	-2.04	-1.57	-2.90	-4.87	-1.52	0.31	18.48
Mexico	-2.60	-2.25	-2.64	-12.12	-2.44	0.20	7.61
Netherlands	-2.69	-1.18	-3.13	-2.99	-1.18	0.39	29.31
Norway	-2.37	1.50	-3.46**	-4.62	-1.47	0.32	19.42
Portugal	-1.70	-1.35	-2.66	-4.14	-1.36	0.33	21.23
Spain	-1.87	-1.87	-2.29	-7.44	-1.91	0.26	12.29
Sweden	-2.10	-2.07	-2.63	-8.83	-2.08	0.24	10.39
Switzerland	-2.57	-1.76	-3.37*	-6.32	-1.73	0.27	14.42
Turkey	-1.28	-1.50	-2.11	-5.46	-1.52	0.28	16.35
United Kingdom	-2.26	-2.09	-2.46	-9.55	-2.18	0.23	9.56
United States	-2.16	-1.73	-2.82	-5.94	-1.69	0.29	15.31
Chile	-5.91***	-4.20***	-20.21***	-36.89***	-4.29***	0.12***	0.67***
India	-4.26***	-3.17***	-4.47***	-25.95***	-3.53***	0.14***	1.17***
Indonesia	-4.39***	-3.16**	-4.57***	-22.80**	-3.37**	0.15***	4.00**
Israel	-2.99	-2.47	-2.87	-14.45	-2.68	0.18	6.33
South Africa	-1.62	-1.51	-2.16	-4.77	-1.49	0.31	18.80

Note: * significance at 10%; ** significance at 5%; *** significance at 1%.

However, the tests applied so far work under the assumption of no breaks in the stochastic process under consideration. The presence of breaks is not a problem if the null of a unit root is rejected, as in the case of Chile, India and Indonesia, although there could still be breaks in these countries. The problematic case is for the majority of the countries where the test cannot reject a unit root, as this rejection can be driven by the presence of breaks. That is, the non rejection of the null may be due to the existence of breaks in the mean of the series, which may reflect changes in the monetary regime, as stated in the literature (see, *inter alia,* Cogley and Sargent, 2001; Clarida *et al*., 2000; Orphanides, 2003; and Corvoisier and Mojon, 2005). In this line, those breaks are very likely to occur in countries pursuing a target inflation policy with a very strong Central Bank.

In what follows, we perform the Lumsdain and Papell (1997) test for unit roots considering two breaks[1]. In all countries, the null hypothesis of a unit root is rejected in favor of the existence of two structural changes. The results are reported in Table 3. These results highly suggest the effects of the intervention of monetary authorities in an attempt to achieve an inflation targeting ……..We observe a first wave of breaks in the 1980s, affecting 10 countries[2]. The second wave took place over the 1990s and affected 27 countries. The third wave took place in the first half of the 2000s and changed the mean in 14 countries[3].

In face of these results, the inflation process follows non-stationary patterns with the presence of structural breaks that have led to changes in the mean, possibly due to the effects of the intervention of monetary authorities in an attempt to reduce the inflation level and to keep it in low levels.

[1] By using the Bai and Perron (2003) test we found no more of two structural breaks in the countries. Therefore a higher number of breaks for the purpose of unit roots testing was not considered.
[2] Two countries that admit two breaks in the 1980s are counted only once.
[3] These dates are highly consistent with dates obtained with the terst of Bai and Perron (2003).

Table 3: Dates of breaks in the mean (1980:01-2009:12)

Countries	Dates of breaks	
Austria	1989:05	1992:05
Belgium	1988:01	1998:08
Canada	1990:11	2006:03
Denmark	1998:05	2001:09
Finland	1992:09	1997:03
France	1995:03	1997:09
Germany	1992:03	1995:11
Greece	1985:01	2003:09
Hungary	1993:09	1998:03
Iceland	1987:07	2006:01
Ireland	1991:05	1994:05
Italy	1983:03	1983:07
Japan	1987:03	1995:01
Korea	1995:02	2000:01
Luxembourg	1995:05	2005:07
Mexico	1993:03	2002:11
Netherlands	1996:03	2001:03
Norway	1990:07	1993:11
Portugal	1993:11	1997:03
Spain	1992:01	1994:07
Sweden	1991:11	1998:11
Switzerland	1998:07	2002:0301
Turkey	1998:01	2006:03
United Kingdom	1998:11	2001:11
United States	2002:11	2004:03
Chile	1989:10	1998:09
India	1991:05	2005:09
Indonesia	1987:07	1997:07
Israel	1989:01	1991:04
South Africa	1984:09	1989:09

3. Methodology framework for the analysis of inflation persistence

3.1 The Persistence Change Model

For the purpose of presenting the persistence change tests, we follow Harvey *et al.* (2006) and Busetti and Taylor (2004) and consider the following data generation process,

$$y_t = x'_t \beta + v_t \tag{1}$$

$$v_t = \rho_t v_{t-1} + \varepsilon_t, \qquad t = 1,...,T \text{ and } v_0 = 0 \tag{2}$$

where y_t is the inflation rate, x_t is a set of deterministic variables, including a constant or a constant and a linear trend and a dummy to account for the change in the mean of the inflation rate. The dummy is equal to zero over the period before the structural break and equal to unity thereafter. The vector x_t is assumed to satisfy the mild regularity conditions of Phillips and Xiao (1998) and the innovation sequence $\{\varepsilon_t\}$ is assumed to be a mean zero process satisfying the familiar α-mixing conditions of Phillips and Perron (1988, p.336) with strictly positive and bounded long-run variance, $\omega^2 \equiv \lim_{T\to\infty} E\left(\sum_{t=1}^{T}\varepsilon_t\right)^2$; (see Harvey *et al.*, 2006, pp. 444).

For estimation purposes, we use the following specification $y_t = \beta_0 + \beta_1 D_t + \beta_2 t + v_t$, (in which; $D_t = 0, \;\; if \;\; t \le \lambda t$; and $D_t = 1, \;\; if \;\; t > \lambda t$; λ represents the timing of the break).

Within (1)-(2), four hypotheses can be considered:

i) H_{1}: y_t is I(1) (*i.e.* nonstationary) throughout the sample period;

ii) H_{01}: y_t is I(0) changing to I(1) (in other words, stationary changing to nonstationary) at time $\left[\tau^* T\right]$. The change point proportion is assumed to be an unknown point in $\Lambda = [\tau_l, \tau_u]$, an interval in (0,1) which is symmetric around 0.5;

iii) H_{10}: y_t is I(1) changing to I(0) (i.e. nonstationary changing to stationary) at time $\left[\tau^* T\right]$;

iv) H_{0}: y_t is I(0) (stationary) throughout the sample period.

With the dummy case, the structural change may occur before or after the change in persistence.

3.2 Ratio-based tests

In order to test the hypotheses in i) - iv), Kim (2000), Kim *et al.* (2002) and Busetti and Taylor (2004) develop tests for the constant $I(0)$ DGP (H_0) against the $I(0)-I(1)$ change DGP (H_{01}) which are based on the ratio statistic,

$$K_{[\tau T]} = \frac{\left(T-[\tau T]\right)^{-2} \sum_{t=[\tau T]+1}^{T} \left(\sum_{i=[\tau T]+1}^{t} \tilde{\upsilon}_{i,\tau} \right)^2}{[\tau T]^{-2} \sum_{t=1}^{[\tau T]} \left(\sum_{i=1}^{t} \hat{\upsilon}_{i,\tau} \right)^2} \tag{3}$$

where $\hat{\upsilon}_{t,\tau}$ is the residual from the OLS regression of y_t on x_t for observations up to $[\tau T]$ and $\tilde{\upsilon}_{t,\tau}$ is the OLS residual from the regression of y_t on x_t for $t=[\tau\ T],...,T$.

Since the true change point, τ^*, is assumed unknown Kim (2000), Kim *et al.* (2002) and Busetti and Taylor (2004) consider three statistics based on the sequence of statistics $\{K(\tau), \tau \in \Lambda\}$, where $\Lambda = [\tau_l, \tau_u]$ is a compact subset of [0,1], *i.e.*,

$$K_1 = T_*^{-1} \sum_{s=[\tau_l T]}^{[\tau_u T]} K\left(s/T\right) \tag{4}$$

$$K_2 = \ln\left\{ T_*^{-1} \sum_{s=[\tau_l T]}^{[\tau_u T]} \exp\left(\frac{1}{2} K\left(s/T\right)\right)\right\} \tag{5}$$

$$K_3 = \max_{s \in \{[\tau_l T],...,[\tau_u T]\}} K\left(s/T\right) \tag{6}$$

where $T_* = [\tau_u T] - [\tau_l T] + 1$ and τ_l and τ_u correspond to the (arbitrary) lower and upper values of τ^* (in the empirical section we set $\tau_l = 0.2$ and $\tau_u = 0.8$, as is frequently adopted in the literature). Limit results and critical values for the statistics in (4) - (6) can be found in Harvey *et al.* (2006).

Note that the procedure in (4) corresponds to the mean score approach of Hansen (1991), (5) is the mean exponential approach of Andrews and Ploberger (1994) and finally (6) is the maximum Chow approach of Davies (1977); see also Andrews (1993).

In order to test H_0 against the I(1) - I(0) change DGP (H_{10}), Busetti and Taylor (2004) propose further tests based on the sequence of reciprocals of K_t, $t = [\tau_l T], ..., [\tau_u T]$. They define K_1^R, K_2^R and K_3^R as the respective analogues of K_1, K_2 and K_3, with K_j, $j = 1, 2, 3$ replaced by K_j^{-1} throughout. Furthermore, to test against an unknown direction of change (that is either a change from I(0) to I(1) or vice versa), they also propose $K_i^M = \max[K_i, K_i^R]$, $i = 1, 2, 3$. Thus, tests which reject for large values of K_1, K_2 and K_3 can be used to detect H_{01}, tests which reject for large values of K_1^R, K_2^R and K_3^R can be used to detect H_{10} and tests which reject for large values of K_1^M, K_2^M and K_3^M can be used to detect either H_{01} or H_{10}.

3.3 Modified ratio-based test

As noted by Harvey *et al.* (2006), all statistics previously presented possess pivotal limit distributions under both H_0 and H_1. Thus, they employ the approach of Vogelsang (1998) to produce tests based on modified versions of these statistics which for a given test and significance level, the critical values are the same under the null and alternative hypothesis and are at the same time equal to corresponding unmodified test.

The modification is largely the same for all tests. In other words, following Vogelsang (1998) and Harvey *et al.* (2006) a modified variant of K_i can be considered as,

$$K_{im} = \exp(-b_i J_{1T}) K_i, \qquad i = 1, 2, 3 \tag{7}$$

Where b_i (with $i = 1, 2, 3$) are finite constants and J_{1T} is a T^{-1} times the Wald statistic for testing the joint hypothesis $\gamma_{k+1} = ... = \gamma_9 = 0$ in the regression,

$$y_t = x'_t \beta + \sum_{i=k+1}^{9} \gamma_i t^i + error, \qquad t = 1, ..., T .$$

Note that J_{1T} is the unit root test statistic proposed by Park and Choi (1988) and Park (1990) which is used to explicitly test for zero frequency nonstationarity of $\{y_t\}$. This statistic will serve as an activation mechanism of the correction factor $\exp(-b_i J_{1T})$. This results from the fact that if the series is stationary, asymptotically, $J_{1T} \to 0$ and hence $\exp(-b_i J_{1T}) \to 1$, whereas if the series is nonstationary J_{1T} will not converge to zero and therefore inducing $\exp(-b_i J_{1T})$ to provide the necessary scaling to adjust the critical value.

Harvey *et al*. (2006) also suggest a variant of this modification procedure, which is perhaps more natural to consider when testing against H_{01}, by replacing the correction factor J_{1T} with $J_{\min} = \min_{\tau \in \Lambda} J_{1,[\tau T]}$, where $J_{1[\tau T]}$ is T^{-1} times the Wald statistic for testing the joint hypothesis $\gamma_{k+1} = ... = \gamma_9 = 0$ in the regression

$$y_t = x'_t \beta + \sum_{i=k+1}^{9} \gamma_i t^i + error, \qquad t = 1, ..., [\tau T].$$

Note that for the reciprocal statistics the $J^R_{\min}$ correction is given from $J^R_{\min} = \min_{\tau \in \Lambda} J_{[\tau T], T}$, where $J_{[\tau T]}$ is T^{-1} times the Wald statistic for testing the joint hypothesis $\gamma_{k+1} = ... = \gamma_9 = 0$ in the regression

$$y_t = x'_t \beta + \sum_{i=k+1}^{9} \gamma_i t^i + error, \qquad t = [\tau T], ..., T$$

Furthermore, as regards the test against an unknown direction of change, the two modifications suggested by Harvey *et al*. (2006) to K_i^M are defined as,

$$K^M_{i,J} = \exp(-b_i^M J_{1,T}) K_i^M \tag{8}$$

and

$$K_{i,\min}^{M} = \exp\left(-b_i^{M^{\tau}} \min\left[J_{\min}, J_{\min}^{R}\right]\right)K_i^{M}, \qquad i = 1, 2, 3. \tag{9}$$

Regarding the necessary *b* values to implement the tests presented in this section we refer to Harvey *et al.* (2006, p. 453) who provide a table with the asymptotic *b* values for modified tests of stationarity or a unit root against a change in persistence.

4. Major Results

The main results are reported in Tables 4, 5 and 6. These results illustrate changes in inflation persistence form I(1) to I(0) in 12 countries, changes from I(0) to I(1) in 4 countries and, finally, no changes in 14 countries.

In general, we may concludee that the high persistence is a common phenomenon to OECD and non-OECD member-states. However, it is high in the majority of the OECD member-states, as well as in a number of European Union members.

The high persistence in a number of countries favors the common view that high inflation persistence is a structural part of the economies and contrasts the alternative view that monetary policy changes can affect inflation persistence.

Table 4: Changes in inflation persistence from I(1) to I(0)

	K_{1m}	K_{2m}	K_{3m}	K_{1m}^{R}	K_{2m}^{R}	K_{3m}^{R}
Canada	2.42	2.16	10.47	4.81**	18.47***	48.01***
Finland	0.18	0.09	1.49	9.45***	9.32***	24.77***
Germany	0.62	0.33	2.04	2.15	1.38	7.50***
Ireland	0.52	0.27	1.26	2.47*	1.72*	8.57**
Italy	0.36	0.19	1.41	3.76*	2.53	9.75
Netherlands	0.20	0.10	0.58	6.22***	5.38***	16.85***
Norway	0.10	0.05	0.54	13.67***	11.70***	31.84***
Sweden	0.41	0.21	1.30	3.53**	3.05**	13.08***
U. Kingdom	0.26	0.14	1.28	6.88**	8.15**	24.77**
U.S.	0.84	0.60	4.48	3.44**	2.20**	7.84*
Chile	0.21	0.11	1.72	9.03***	7.43***	19.74***
India	0.47	0.26	2.73	3.86*	3.25	12.36

Table 5: Changes in inflation persistence from I(0) to I(1)

	K_{1m}	K_{2m}	K_{3m}	K_{1m}^{R}	K_{2m}^{R}	K_{3m}^{R}
France	2.74*	1.51*	4.95	0.44	0.23	1.62
Japan	2.18	1.41	7.07*	0.87	0.55	3.78
Mexico	2.31	1.51*	8.21*	0.55	0.28	1.51
Israel	8.05***	8.20**	24.74**	1.42	5.62**	20.92*

Table 6: No evidence of changes in inflation persistence

	K_{1m}	K_{2m}	K_{3m}	K_{1m}^{R}	K_{2m}^{R}	K_{3m}^{R}
Austria	1.20	0.63	2.44	1.03	0.58	4.20
Belgium	1.54	0.86	4.13	0.89	0.49	3.20
Denmark	0.87	0.50	3.54	1.91	1.25	6.27
Greece	0.54	0.28	1.77	2.76	1.84	9.30
Hungary	1.19	0.70	5.07	1.26	0.72	3.55
Iceland	0.61	0.32	1.87	2.05	1.13	3.85
Korea	1.21	0.63	3.55	0.88	0.45	1.43
Luxembourg	1.28	0.86	5.17	1.56	1.01	5.88
Portugal	1.20	1.07	7.24	2.05	1.27	5.66
Spain	0.48	0.25	2.26	2.88	2.19	11.87
Switzerland	1.24	0.67	3.20	1.10	0.62	3.56
Turkey	1.37	1.22	7.65	1.44	0.79	2.91
Indonesia	1.65	1.19	6.38	1.14	0.69	3.96
S. Africa	0.93	0.50	2.21	1.59	0.98	4.92

References

Altíssimo, F. (2003), “Testting for breaks: as application to euro area inflation” - delete

Andrews, D. W. K. (1993), “Tests for parameter instability and structural change with unknown change point,” *Econometrica* 61, 821-856.

Andrews, D. W. K. and H. Chen (1994), “Approximately median-unbiased estimation of autoregressive models,” *Journal of Business and Economic Statistics* 12, 187-204.

Andrews, D. W. K. and W. Ploberger (1994), “Optimal tests when a nuisance parameter is present only under the alternative,” Econometrica 62, 1383-1414.

Bai, J. and P. Perron (2003), “Computation and analysis of multiple structural change models,” *Journal of Applied Econometrics* 18, 1-22.

Barsky, R. (1987), “The Fisher hypothesis and the forecastibility and persistence of inflation,” *Journal of Monetary Economics* 19, 3-24.

Batini, N. and E. Nelson (2002), “The lag from monetary policy actions to inflation: Friedman revisited,” Bank of England, Discussion Paper 6.

Batini, N. (2002), “Euro area inflation persistence,” ECB Working Paper 201.

Benati and Wood, 2004

Benati, L. (2003), “Structural breaks in inflation dynamics,” *Computing in Economics and Finance* 169.

Benati, L. (2004), “International evidence on inflation persistence,”

Brainard, W. and G. Perry (2000), “Making policy in a changing world,” In Perry, G., Tobin, J. (Eds.), Economic Events, Ideas and Policies: The 1960s and After, Brookings Institution Press, Washington.

Busetti, F. and A. Taylor (2004), “Tests of stationarity against a change in persistence,” *Journal of Econometrics* 123, 33-66.

Cati, R. C., M. G. P. Garcia and P. Perron (1999), “Unit roots in the presence of abrupt governmental intervention with an application to brazilian data,” *Journal of Applied Econometrics* 14, 27-56.

Cecchetti, S. and G. Debelle (2004), “Has the inflation process changed?,” Working Paper, Bank for International Settlements.

Clarida, R. J. Gali and M. Gertler (2000), "Monetary policy rules and macroeconomic stability: Evidence and some theory," *Quarterly Journal of Economics* 115, 157-180.

Cogley, T. and T. Sargent (2001), "Evolving post-World War II inflation dynamics," NBER Macroeconomics Annual, 16.

Corvoisier, S. and B. Mojon (2005), "Breaks in the mean of inflation: how they happen and what to do with them," ECB Working Paper 451.

Dias, D and C. Marques (2004), "Using mean reversion as a measure of persistence,"

Dickey, D and W. Fuller (1979), "Distribution of the estimators for autoregressive time series with a unit root," *Journal of the American Statistical Association* 74(366), 427-431.

Elliott, G., T. Rothenberg and J. Stock (1996), "Efficient tests for and autoregressive unit root," *Econometrica* 64, 813-836.

Evans and Wachtel, 1993

Gadea, M. D. and L. Mayoral (2006), "The persistence of inflation in OECD countries: a fractionally integrated approach," *International Journal of Central Banking*, March (4), 51-104.

Gadzinski, G. and F. Orlandi (2004), "Inflation persistence for the EU countries, the euro area and the US,"

Goodhart and Hofmann, 2003

Hansen, B. E. (1991), "Tests for parameter instability in regressions with I(1) processes," *Journal of Business and Economic Statistics* 10, 321-335.

Harvey, D. L., S. J. Leybourne and A. M. R. Taylor (2006), "Modified tests for a change in persistence," *Journal of Econometrics* 134, 441-469.

Kim, J. (2000), "Detection of change in persistence of a linear time series," *Journal of Econometrics* 95, 97-116.

Kim, J., J. Franc and R. Amador (2002), "Corrigendum to "Detection of change in persistence of a linear time series," *Journal of Econometrics* 109, 389-392.

Kim, J.Y., J. Belaire Franch and R. Amador (2002), 'Corringendum to detection of change in persistence of a linear time series'. *Journal of Econometrics* 109, 389-392.

Levin and Piger, 2004

Lumsdain, R. and D. Papell (1997), "Multiple breaks and the unit root hypothesis," *The Review of Economics and Statistics*, 79(2), 212-218.

Marques (2004)

Ng, S. and P. Perron (2001), “Lag lengh selection and the construction of unit root tests with good size and power,” *Econometrica* 69(6), 1519-1554.

O’Reilly, G. and K. Whelan (2004), “Has Euro-area inflation persistence changed over time?,” Working Paper 335, European Central Bank.

Orphanides, A. (2003), “The quest of prosperity without inflation,” Journal of Monetary Economics 50, 633-663.

Park (1990)

Park and Choi (1988)

Phillips, P. C. B. and P. Perron (1988), “Testing for a unit root in time series,” *Biometrika* 75, 335-346.

Phillips and Xiao (1998)

Pivetta, F. and R. Reis (2006), “The persistence of inflation in the United States,” *Journal of Economic Dynamics and Control*, forthcoming.

Pivetta, F. and R. Reis (2007), “The persistence of inflation in the United States,” *Journal of Economic Dynamics and Control*,” 31, 1326-1358.

Ravenna, 2000

Taylor, J. (2000), ”Low inflation, pass-through, and the pricing power of firms,” *European Economic Review* 44(7), 1389-1408.

Vogelsang (1998)

Willis, J. (2003), “Implications of structural changes in the US economy for pricing behavior and inflation dynamics,” *Economic Review*, First Quarter 2003, Federal Reserve Bank of Kansas City.