The Hypothesis of a Unit Root in OECD Inflation Revisited

Diego Romero-Ávila and Carlos Usabiaga

Abstract

This paper investigates the hypothesis of a unit root in inflation for 13 OECD countries over the period 1957-2005, taking into account cross-sectional dependence and multiple mean shifts. We conduct unit root testing with the more powerful unit root tests with cross-dependence proposed by Smith *et al.* (2004) and a bootstrap version of the panel stationarity test of Hadri (2000), which provide inconclusive evidence on the time series properties of OECD inflation rates. To shed some light on this issue, we employ the recently developed panel stationarity test of Carrión-i-Silvestre *et al.* (2005) that assumes a highly flexible trend function by incorporating an unknown number of breaks in level. Overall, our confirmatory analysis renders clear-cut evidence in favor of regime-wise stationarity. Furthermore, the breaks in inflation detected are closely associated with macroeconomic shocks and changes in monetary policy.

Key words: OECD Inflation, Panel Stationarity Test, Panel Unit Root Test, Structural Breaks, Cross-Sectional Dependence, Bootstrap Distribution.

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1. Introduction

Since the seminal work by Nelson and Plosser (1982), researchers have paid considerable attention to the presence of stochastic trends in macroeconomic variables. That work influenced the way macroeconomists think about secular trends and short-run fluctuations. Since the presence of a unit root in a macroeconomic variable implies that shocks are permanent, short-run fluctuations can no longer be explained by the traditional view as simple deviations around a deterministic trend. Rather, the trend function itself can be subject to permanent shocks leading to fluctuations in the underlying series.

In this paper we focus on the time series properties of the inflation rate, which is one of the macroeconomic variables drawing more attention among researchers in economics. The existence of a high degree of persistence in inflation has important implications with regard to macroeconomic modeling and economic policy formulation. For example, a unit root behavior in inflation would not be consistent with the sticky price model analyzed by Taylor (1979) or the Phillips curve model of Calvo (1983), for which the price level is trend-stationary and nonstationary, respectively. In addition, in the rational expectations version of the model of Cagan (1956) –see Sargent (1983)–, the stability of the growth rate of money supply requires the mean-stationarity of inflation in steady state. Stationarity in inflation also appears consistent with the hypothesis of a "natural rate of inflation", while the acceleracionist view implies an ever-increasing level of inflation in order to keep unemployment below its natural rate. Therefore, in the case where the unemployment rate is kept below the natural rate, the inflation would be characterized as a unit root.

The presence of a unit root in inflation also affects the intertemporal allocation of savings and investment in line with Fisher (1930) as well as the Fisher-type effects in the transmission process of monetary policy. Another related area where a unit root in inflation matters is standard consumption theory, where the Euler equation implies that stable consumption growth is congruent with stationarity in real interest rates. Finally, a unit root in inflation is also relevant for the convergence analysis based on cointegration techniques –see, for instance, Artis and Nachane (1990) for European Monetary System countries.

Much of the empirical literature on the analysis of the time series properties of inflation rates has used univariate unit root tests of the augmented Dickey-Fuller (1979, ADF) type. To name a few studies, Nelson and Schwert (1977), McDonald and Murphy (1989), Ball and Cechetti (1990), Johansen (1992) and Brunner and Hess (1993), for different countries and time periods, fail to reject the unit root hypothesis in inflation. Following the insights of the seminal work by Perron (1989), Culver and Papell (1997) apply the sequential break unit root test of Perron and Vogelsang (1992) to monthly inflation series of 13 OECD countries for the period 1957-1994. They provide evidence of stationarity in inflation for only four countries at the 5% level or better. In addition, Culver and Papell (1997) try to raise statistical power by applying the panel unit root test of Abuaf and Jorion (1990), which enables them to reject the unit root hypothesis for both the whole panel and smaller panels with as few as three countries. Since Culver and Papell (1997) failed to allow for cross-sectional correlation which could lead to dramatic size distortions,[1] Lee and Wu (2001) applied a bootstrap version of the panel unit root test of Im *et al.* (2003, IPS hereafter) along with the Seemingly Unrelated approach by Taylor and Sarno (1998). Their findings continued lending support to mean reversion in OECD inflation rates.

However, as stressed by Taylor and Sarno (1998) and Karlsson and Lothgren (2000), heterogeneous panel unit root tests, even with cross-dependence, are likely to reject the joint nonstationarity null when there is a single stationary series in a system otherwise nonstationary. Indeed, the application in this paper of the five panel unit root tests with cross-dependence proposed by Smith *et al.* (2004) leads us to reject the null of joint nonstationarity in favor of the alternative hypothesis that at least one of the individual series is stationary. As a result, following Hadri (2000), we should confirm the validity of these findings by employing a panel test that takes the opposite null hypothesis of joint stationarity. Interestingly, this null hypothesis is strongly rejected with a bootstrap version of the test of Hadri (2000). Therefore, the evidence from panel statistics with cross-dependence is clearly inconclusive.

In similar spirit to Perron (1989), we argue that the failure to allow for structural breaks in mean inflation may be responsible for the evidence of nonstationarity obtained with Hadri's test. To account for this possibility, we take the analysis one step further by employing the panel stationarity test of Carrión-i-Silvestre *et al.* (2005, CBL hereafter), which jointly allows for multiple regime changes and cross-sectional dependence through bootstrap methods. There are not only statistical reasons to allow for structural change in the data generation process of inflation series, but also important economic reasons. As stressed by Rapach and Wohar (2005), there is extensive evidence of structural breaks in the mean inflation rate of OECD countries due to monetary policy changes and other macroeconomic shocks like the oil crises of the 1970's. The evidence provided by the CBL test indicates that only after jointly accounting for multiple mean shifts and cross-dependence do our findings support mean stationarity in OECD inflation rates over the postwar era. In the next section we present in more detail the results of the analysis of the time series properties of OECD inflation, and section 3 concludes.

2. Empirical results

We extend the inflation rate series initially computed by Culver and Papell (1997) on the basis of monthly consumer price index data. The original period, which starts in February 1957 and ends in September 1994, is updated through to June 2005. The list of OECD countries analyzed are Belgium, Canada, Finland, France, Germany, Italy, Japan, Luxembourg, the Netherlands, Norway, Spain, the United Kingdom and the United States. The source of the data is the *International Financial Statistics* (International Monetary Fund).

Since traditional panel unit root tests which are derived under the assumption of cross-sectional independence are well known to be subject to large size distortions, we need to test whether this assumption holds in practice. To shed some light on the presence of cross-sectional dependence in inflation innovations we use the CD statistic of Pesaran (2004), which is based on the average of pair-wise correlation coefficients of Ordinary Least Squares residuals. For each country we compute the residuals from ADF regressions whose optimal lag-order is determined using the general-to-specific procedure suggested by Ng and Perron (1995) with a maximum lag-order of eight. Since the CD statistic takes on a value of 15.677, we are able to strongly reject the null hypothesis that output innovations are cross-sectionally independent at the 1% level. This result is plausible and reflects the high degree of cross-sectional dependence induced by trade links, international capital movements and common shocks hitting the sample of industrialized countries under scrutiny.

To account for this cross-sectional dependence, we will apply bootstrap methods to obtain the critical values of the panel statistics of Smith *et al.* (2004). The results from the computation of these more powerful panel unit root tests are reported in Table 1. The first two tests are the standard t-bar statistic ($\Psi_{\bar{t}}$) and the Lagrange Multiplier (LM) panel statistic (Ψ_{LM}) of IPS. They also develop panel versions of some powerful modifications of univariate ADF t-statistics such as the Max t-statistic and the Min LM-statistic, which are both based on forward and reverse ADF regressions, as well as the weighted symmetric (WS) test. These panel statistics are denoted by Ψ_{Max}, Ψ_{Min} and Ψ_{WS}.[2] The p-values of the five tests are computed employing 20,000 bootstrap replications in order to control for general forms of cross-sectional dependence as well as for finite-sample bias. We define a block size equal to 30 and a maximum lag-order for whitening the residuals equal to eight.[3] The evidence provided by the bootstrap version of the five panel unit root t-statistics of Smith *et al.* (2004) lends little support to the presence of a unit root in inflation, as they all reject the null of nonstationarity at the 1% significance level.

[Insert Table 1 about here]

However, as noted by Taylor and Sarno (1998) and Karlsson and Lothgren (2000), the null hypothesis of joint nonstationarity is likely to be rejected with heterogeneous panel unit root tests due to the presence of as few as a single stationary series. Therefore, we should complement these results with the application of the panel stationarity test of Hadri (2000), which takes the opposite null hypothesis of joint stationarity. This allows us to provide confirmatory evidence of stationarity in OECD inflation rates.

The results from the panel stationarity test of Hadri (2000) are reported in Table 2.[4] Panel A presents the univariate KPSS tests upon which the panel test is based. Remarkably, we are able to reject the null of stationarity for the 13 countries at the 5% significance level or better. In line with the evidence from univariate tests, we strongly reject the stationarity null with the Hadri test, irrespective of the assumption regarding cross-sectional dependence. Therefore, the evidence from the panel unit root tests of Smith *et al.* (2004) and the panel stationarity test of Hadri (2000) appears contradicting, as the former points to stationarity in inflation while the latter leads to the opposite conclusion.

[Insert Table 2 about here]

However, as stressed by Perron (1989), the evidence of nonstationarity from Hadri's test could be due to the failure to allow for structural breaks in the data generation process of inflation rates, which may lead to misinterpret mean stationarity with structural breaks as a unit root. To shed some light on this issue, we apply the recently developed panel stationarity test with multiple breaks of

CBL, which allows for multiple level shifts, thereby accommodating general forms of cross-dependence through bootstrap methods.

This test is a generalization for the case of multiple structural breaks of the panel stationarity test of Hadri (2000). Let $y_{i,t}$ be the stochastic process which under the null hypothesis of stationarity is characterized by:

$$y_{i,t} = \alpha_i + \sum_{k=1}^{mi} \gamma_{i,k} DU_{i,k,t} + \varepsilon_{i,t} \tag{1}$$

where *i=1…N* and *t=1…T* stand for the number of panel members and time periods, respectively. The dummy variable for the changes in level is given by $DU_{i,k,t}$ such that: $DU_{i,k,t} = 1$ for $t > T^i_{b,k}$ and *0* otherwise, with $T^i_{b,k}$ denoting the *kth* break location for the *ith* individual for $k = 1 \ \dots \ m_i$, $m_i \geq 1$. To determine the location (λ_i) and the number of breaks, we apply the methodology proposed by Bai and Perron (1998). The breaks location is chosen on the basis of the global minimization of the sum of squared residuals using a trimming value of 0.15, and the optimal number of breaks is selected using the sequential procedure based on pseudo F-statistics. After determining the vector $\hat{\lambda}_i$ for each *i*, we compute the normalized test statistic as follows:[5]

$$Z(\hat{\lambda}) = \sqrt{N}\left(\eta(\hat{\lambda}) - \bar{\xi}\right)/\bar{\varsigma} \ \xrightarrow{d} \ N(0,1) \tag{2}$$

where $\eta(\lambda)$ is the average of univariate KPSS tests with breaks and $\bar{\xi}$ and $\bar{\varsigma}^2$ are computed as averages of individual means and variances of univariate KPSS tests. Since the assumptions of cross-independence and asymptotic normality are unlikely to hold, we compute the bootstrap distribution of the panel statistic following Maddala and Wu (1999).

Table 3 reports the results from the univariate and panel KPSS tests with multiple level shifts. In order to control for finite-sample bias resulting from the use of asymptotic critical values, we will compute finite-sample critical values for the individual KPSS tests with multiple breaks by means of Monte Carlo simulations using 20,000 draws. Beginning with the individual KPSS test statistics, it is remarkable that all countries under consideration have experienced at least one change in their mean inflation rate. For the Netherlands and the United Kingdom we are able to reject the null of regime-wise stationarity at the 1% level, for Finland we reject the null at the 5% level, and for Belgium and Japan at the 10% level. With the panel KPSS test –shown in Panel B–, which assumes asymptotic normality and cross-sectional independence, we are able to reject the null of regime-wise stationarity at the 1% level. However, these conclusions are completely overturned when we compare the panel KPSS statistic with the bootstrap critical values –shown in Panel C–, which control for general forms of cross-sectional dependence as well as for finite-sample bias. In this case, we fail to reject the null of regime-wise stationarity even at the 10% level, thus supporting the view that inflation rates can be best described as regime-wise stationary.[6] These results stand in stark contrast to those obtained with Hadri's test, thus rendering clear-cut support to Perron (1989)'s argument about the need for considering segmented trends when investigating the nonstationarity properties of macroeconomic series.[7]

Taking as a whole, our analysis indicates that considering cross-sectional dependence alone yields inconclusive results, and it is only after taking into account both cross-sectional dependence and multiple breaks that our analysis renders evidence supportive of joint stationarity in OECD inflation rates.[8]

[Insert Table 3 about here]

Looking at the number of breaks, two countries –Canada and the United States– are shown to have four mean shifts, six countries –Finland, France, Italy, Japan, Spain and the United Kingdom– exhibit three breaks, four countries –Belgium, Germany, Luxembourg and Norway– present two breaks and the Netherlands only one break. This provides strong evidence of the need for controlling for multiple level shifts in inflation due to the occurrence of infrequent permanent shocks to OECD inflation over the past five decades. In addition, we can observe some clustering of the break dates. Among the 35 breaks identified, seven breaks occurred in each of the periods 1964-1970 and 1972-1973, 12 breaks are detected during the period 1981-1985, and seven breaks during the period 1988-1992. It is also interesting to note that all the breaks occurring in the two first clusters exhibit a significantly positive sign. This reflects the wide-spread tendency for inflation rates to rise over the mid-late sixties and early seventies due to burgeoning economic activity – caused in part by expansionary economic policy– and the first great oil shock, respectively.[9] Conversely, all the breaks detected in the eighties and early nineties show a negative sign, which highlights the prevailing tendency for inflation rates to fall due to the change in monetary policy stance towards inflation targeting and the commitment adopted by European Union member states to fulfill the Maastricht criteria aimed at achieving nominal convergence.

Our highly flexible specification of the trend function that incorporates multiple mean shifts in inflation rates has enabled us to provide a detailed account of the most important infrequent changes in inflation rates over the past decades. These findings, though more detailed, are in line with those from previous studies. Culver and Papell (1997) find that more than half of the countries experienced a sharp rise in inflation coinciding with the oil-price shock of 1973-1974. For the United States the timing of the break is identified in the early eighties, thus reflecting the downward shift in inflation caused by Volcker's monetary stabilization implemented in 1979. For the remaining countries, the breaks take place during the mid-eighties coinciding with the reversion of oil prices to pre-shock levels. Corvoisier and Mojon (2005) also investigate the presence of multiple breaks in the mean of inflation for a sample of 22 OECD countries over the past 40 years. Their timing of breaks is closely related to ours as they find three main clusters of breaks:[10] the first one related to the wide-spread rise in inflation during the late-sixties and early seventies, a second one occurring in the first half of the eighties and a third one in the early nineties, both reflecting the shift in monetary policy stance towards more stringent measures to fight inflation.[11] That stage culminated in the adoption of inflation targeting during the early nineties in countries such as Australia, Canada, Finland, France, Spain, Sweden, the United Kingdom and the United States.

In addition, Benati and Kapetanios (2003) also investigate the occurrence of mean shifts in inflation for 18 OECD countries by using a univariate unit root test that allows for multiple structural breaks. In line with previous findings, they identify three waves of breaks associated with the early seventies, early eighties and early nineties, all broadly coinciding with the major macroeconomic shocks and institutional changes described above. Finally, it is interesting to point out that Boschen and Weise (2003) provide evidence that US inflation episodes turn out to be a triggering factor for subsequent outbreaks in inflation across other OECD countries.

4. Conclusions

The benchmark for this article is that, employing both the panel unit root tests with cross-dependence of Smith *et al.* (2004) as well as a bootstrap version of the panel stationarity test of Hadri (2000), the evidence regarding the time series properties of OECD inflation rates appears contradicting. Since the failure to account for structural change in the mean inflation rate could be responsible for the findings of nonstationarity with the test of Hadri (2000), we have deployed the panel stationarity test of CBL which incorporates multiple mean shifts. Our findings strongly reject the presence of a unit root in inflation, thus lending firm support to the existence of regime-wise stationarity for our panel of countries. Therefore, our analysis indicates that considering cross-sectional dependence alone yields inconclusive results, and it is only after accounting for both cross-sectional dependence and multiple breaks that our analysis does provide support for joint stationarity in OECD inflation rates.

Furthermore, our results indicate a clear pattern of inflation dynamics common to all countries: inflation increased in the sixties and early seventies and then fell during the eighties and nineties. The timing of the breaks estimated in the analysis concurs with those obtained in previous studies: a first wave of breaks over the mid-late sixties associated with the rise in inflation caused by expansionary economic policy leading to burgeoning economic activity; a second wave of breaks in the early seventies coinciding with the wide-spread rise in inflation due to sharp increases in energy prices; a third wave in the early eighties associated with the drop in inflation caused by the more decisive fight against inflation following Volcker's disinflationary policy in the United States; and finally a fourth wave of breaks in the early nineties coinciding with a further drop in inflation rates due to the wide-spread adoption by central banks of inflation targeting rules in addition to the effort made by European Union member states to nominally converge towards the Maastricht criteria.

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TABLES

Table 1: Panel unit root tests without breaks

	Tests	p-value
$\Psi_{\bar{t}}$	-4.658***	0.000
Ψ_{Max}	-4.613***	0.000
Ψ_{LM}	25.474***	0.000
Ψ_{Min}	24.947***	0.000
Ψ_{WS}	-4.722***	0.000

The p-values for these panel unit root tests are computed employing 20,000 bootstrap replications and defining a block size equal to 30. The maximum lag order is set at 8. *** implies rejection of the null hypothesis at 1%.

Table 2: Panel KPSS stationarity test without structural breaks

Panel A: Country-specific test

Country	Univariate KPSS test
Belgium	0.882***
Canada	1.031***
Finland	1.882***
France	1.106***
Germany	0.937***
Italy	0.578**
Luxembourg	1.505***
Netherlands	1.368***
Norway	1.763***
Japan	2.841***
Spain	1.761***
United Kingdom	0.931***
United States	0.532**

Panel B: Panel KPSS test without breaks assuming cross-sectional independence

	Test	p-value
Panel KPSS Test	27.824***	0.000

Panel C: Bootstrap distribution (%)

	1	2.5	5	10	90	95	97.5	99
Panel KPSS Test	-2.080	-1.891	-1.710	-1.468	1.845	2.886	3.896	5.442

The finite sample critical values for the univariate KPSS test obtained from Sephton (1995) are 0.739 (1%), 0.463 (5%) and 0.347 (10%). The bootstrap critical values for the panel KPSS test are obtained through bootstrap methods with 20,000 replications following Maddala and Wu (1999). *** and ** imply rejection of the null hypothesis at 1% and 5%, respectively.

Table 3: Panel KPSS stationarity test with multiple structural breaks

Panel A: Country-specific test

Country	Univariate KPSS test	m_i	$\hat{T}^i_{b,1}$	$\hat{T}^i_{b,2}$	$\hat{T}^i_{b,3}$	$\hat{T}^i_{b,4}$	Finite sample critical values		
Belgium	0.112*	2	1970(12) +	1985(4) –			0.098	0.118	0.169
Canada	0.056	4	1964(10) +	1972(6) +	1982(6) –	1991(6) –	0.056	0.066	0.091
Finland	0.100**	3	1972(12) +	1982(5) –	1991(5) –		0.072	0.086	0.118
France	0.063	3	1973(3) +	1983(10) –	1991(11) –		0.073	0.087	0.117
Germany	0.073	2	1969(11) +	1982(6) –			0.106	0.131	0.198
Italy	0.056	3	1972(7) +	1984(2) –	1996(5)–		0.071	0.084	0.113
Luxembourg	0.073	2	1969(11) +	1983(12) –			0.102	0.124	0.181
Netherlands	0.386***	1	1982(10) –				0.153	0.188	0.263
Norway	0.059	2	1969(12) +	1988(3) –			0.097	0.115	0.159
Japan	0.098*	3	1969(12) +	1977(5) –	1984(10) –		0.088	0.109	0.163
Spain	0.047	3	1973(3) +	1984(1) –	1992(2) –		0.072	0.086	0.117
United Kingdom	0.178***	3	1973(8) +	1981(4) –	1992(5) –		0.073	0.088	0.120
United States	0.056	4	1965(10) +	1973(1) +	1981(9) –	1991(1) –	0.057	0.067	0.090

Panel B: Panel KPSS test with multiple breaks assuming cross-sectional independence

	Test	p-value
$Z(\hat{\lambda})$	6.677***	0.000

Panel C: Bootstrap distribution (%)

	1	2.5	5	10	90	95	97.5	99
$Z(\hat{\lambda})$	0.583	1.046	1.474	1.970	6.682	7.478	8.259	9.198

$Z(\hat{\lambda})$ denotes the panel KPSS test with multiple breaks developed by CBL assuming heterogeneity in the estimation of the long-run variance. The finite sample critical values are obtained through Monte Carlo simulations with 20,000 replications. The bootstrap distribution for $Z(\hat{\lambda})$ is obtained through bootstrap methods with 20,000 replications following Maddala and Wu (1999). ***, ** and * imply rejection of the null hypothesis at 1%, 5% and 10%, respectively. +/– represent an upward shift and a downward shift, respectively. The maximum number of breaks is set at 5.

ENDNOTES:

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[1] O'Connell (1998), Maddala and Wu (1999), Strauss and Yigit (2003) and Banerjee *et al.* (2005) provide overwhelming evidence supporting these claims.

[2] All the five tests take as null a unit root in all individuals versus the alternative of stationarity for at least one individual panel member. The value of $\Psi_{\bar{t}}$, Ψ_{Max} and Ψ_{WS} must be compared with the critical values from a lower-tailed standard normal distribution, while Ψ_{LM} and Ψ_{Min} with the critical values from an upper-tailed standard normal distribution.

[3] The results remain unchanged for a block size equal to 100 and other maximum lag-orders different from eight.

[4] This test is computed as the average of univariate Kwiatkowski *et al.* (1992, KPSS hereafter) stationarity tests. When constructing these statistics, we obtain an autocorrelation and heteroskedasticity consistent estimate of the long-run variance of the residuals using the procedure developed by Kurozumi (2002) with the Quadratic Spectral kernel and fixed bandwidth.

[5] The null hypothesis implies regime-wise stationarity for all *i*, versus the alternative of nonstationarity for some *i*. The value of the $Z(\hat{\lambda})$ statistic must be compared with the critical values from the upper tail of the standard normal distribution.

[6] Therefore, the bootstrap distribution appears to dramatically shift to the right of the upper tail of the standard normal distribution due to size distortions caused by the presence of error cross-sectional dependence as evidenced by the test of Pesaran (2004).

[7] We checked for the robustness of our results obtained from the panel stationarity test with multiple breaks by reducing the sample size of our panel of 13 countries. However, we should note that when we analyze the full sample, which includes the five countries (Belgium, Finland, the Netherlands, Japan and the United Kingdom) for which we could reject the stationarity null with the univariate KPSS tests with multiple breaks, we are facing the most possible way of rejecting the null of regime-wise stationarity in favor of nonstationarity in OECD inflation rates. Not surprisingly so, when we drop from the full sample these five countries one at a time, in small groups, or all

together, the evidence continues supporting the rejection of the unit root hypothesis. Similar evidence arises from other country combinations such as the panel with the G7 countries. To conserve space, we do not report the results from these robustness checks, which are available from the authors upon request.

[8] Our main findings also hold for the time span investigated by Culver and Papell (1997). These unreported results are available from the authors upon request.

[9] Boschen and Weise (2003) find that the main reason behind the inflation episodes of the mid-late sixties was the tendency for policy makers to target excessively high rates of economic growth.

[10] See Corvoisier and Mojon (2005, p. 42) for a summary table of the main results regarding the location of the breaks in mean inflation from several studies. Interestingly, the timing of the breaks widely coincide with ours.

[11] According to these authors, the early and mid eighties were characterized by marked changes in the institutional environment as well as in the implementation of monetary policy that moved towards a more decisive fight against inflation. As they note, countries like France, Italy, Portugal, Spain, the United Kingdom and the United States were all subject to these changes in monetary policy stance. Our results may closely reflect these institutional changes as we find breaks during the period 1981-1985 in all the countries under analysis except for Norway that exhibits a downward shift in inflation in 1988.