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Are Inflation Rates Stationary in 11 Mediterranean Countries? Evidence from Univariate and Panel Unit Root Tests

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Abstract

In this study, we employ unit root tests that allow for heterogeneous structural breaks developed by Lee and Strazicich (2003), Narayan and Popp (2010), and Im et al. (2010) to examine the stationarity of inflation rates in 11 Mediterranean countries over the period 1960-2011. The empirical findings support the stationarity of inflation rates when structural breaks are allowed. Thus, most shocks to inflation rates appear to be temporary and inflation rates show mean reversion. In addition, the results indicate some important policy implications.

Keywords: *Inflation, Monetary Policy, Stationarity, Unit Root, Structural Breaks*

JEL Code Classification: *C23, E31, E52*

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1. Introduction

Since the seminal paper of Nelson and Plosser (1982), who investigated integrational properties of important macroeconomic variables of US, the investigation of time series properties of macroeconomic variables has become a central research area in economics discipline. Unfortunately, there is no consensus on this issue in the empirical literature.

Especially, analyzing the unit root process in main macroeconomic aggregates, such as unemployment rate, real exchange rate, per capita real GDP, energy consumption or production is of importance since they shed some light on the validity of relative macroeconomic theories. For instance, the stationarity of per capita real GDP, unemployment rate, and real exchange rate support the theories of stochastic income convergence, natural rate of unemployment, and purchasing power parity, respectively. There are many applied studies concerning these theories.

In this paper, we focus on time series properties of inflation rate, which is one of the main macroeconomic variables drawing more attention among researchers in economics. The Fisher effect, the traditional capital asset pricing model, the accelerationist hypothesis, monetary policies about controlling money supply growth, the expectations augmented Phillips curve model, and the unemployment hysteresis are some of the relevant research areas in economics related to inflation rate. In addition, the following reasons lead economists to examine the integration properties of inflation rate.

First, given that inflation largely influences spending and saving decisions, the rate of inflation is usually taken into account in the conduct of monetary policy. Central banks rely on their knowledge of time series properties of inflation when conducting monetary policy and inflation rate targeting because the cost of disinflation is different when inflation shows an I (0) or I (1) behavior. As noted by Arize et al. (2005, p.462), inflation is unpopular due to the real cost it imposes on the economy and society; hence, the level of inflation is generally taken into consideration in the conduct of monetary policy. Nonstationary inflation rate implies that shocks to inflation have a permanent effect, and thus it is associated with a high cost for the policies of disinflation. However, stationary inflation rate will incur a lower cost for the monetary authorities in conducting monetary policies (Cecchetti and Debelle, 2006). Thus, knowledge on integration properties of inflation rate is essential for policy makers to control inflation.

Second, “inflation is typically regarded as a key variable in many economic models, whose validity hinges critically on whether inflation is I(0) or not” (Basher and Westerlund, 2007). For instance, the validity of one of the most important finance theories, namely Fisher effect, depends on the existence of non-stationary inflation rate. The Fisher equation does hold in the case that inflation rate has I (1) process and is cointegrated with the nominal interest rate in order for the real interest rate to be stationary. Furthermore, another important macroeconomic theory, the expectations

augmented Phillips curve model, where wages and prices share a long-run relationship, requires inflation rate to be non-stationary. Also, the accelerationist hypothesis indicates nonstationary inflation rate. In this theory, authorities have to accept an ever increasing inflation level to keep unemployment level below its natural rate. Also, according to the version of rational expectation behavior developed by Cagan (1956), a stable money growth is accompanied by a stationary inflation rate unless there exist bubbles. In other words, rational expectations hypothesis proposes that stable growth of money supply implies stationary inflation rate.

Third, stationarity of inflation rate is important in the selection of right econometric modeling as well. For instance, in the case of stationary inflation rate, a vector autoregressive model (VAR) in levels can be estimated to reduce the chances of generating spurious results with unreliable policy decisions. However, it is more convenient to study a system of inflation rates containing nonstationary variables by applying a cointegration or vector error correction (VEC) model (see. Arize et al., 2005).

Fourth, the properties of integration of inflation rate are informative for convergence of inflation rates among a group of countries. In the case of nonstationary inflation rates, an inflation convergence within the European Monetary System (EMS) might be identified as a cointegrating relation between inflation rates in Germany and each of the other EMS countries. In addition, testing convergence in inflation rates of new member states of the European Union (EU) is crucial to determine whether or not they have met the inflation criterion, as defined in the Maastricht Treaty (Im et al., 2010).

Taking all these above mentioned reasons into consideration, this study aims at contributing to the literature in two aspects. First, to the best of my knowledge, this is the first study that analyzes the integrational properties of inflation rate in Mediterranean countries. Second, we use recently developed univariate and panel unit root tests of Narayan and Popp (2010) and Im et al. (2010). There are no earlier studies in the related literature applying these unit root tests to the inflation rate series, except for Im et al. (2010)

The remainder of paper is organized as follows. Section 2 includes the literature review. In section 3, empirical strategy, i.e. the methodology of unit root tests used, is explained. Section 4 describes the data and reports the main results. Finally, in Section 5, we suggest some policy implications and conclude the study.

2. Literature Review

As stated by Narayan and Narayan (2010), “in comparison with the literature on unit root tests in real GDP and real exchange rate, the literature on the unit root hypothesis of the inflation rate is scarce.” However, there are important studies in the literature. We can classify them into seven strands.

The first strand of studies employs conventional univariate unit root tests to decide whether inflation rate is stationary or not. Lai (1997), for instance, who employed modified Dickey-Fuller Test, developed by Park and Fuller (1995) to the inflation rates of G7 countries, found inflation rates to be stationary. Charemza et al. (2005) examined the stationarity of 93 worldwide inflation rates by conventional augmented Dickey and Fuller (ADF, 1979, 1981) test and compared the results to those obtained by allowing for symmetric stable innovations. The findings indicate that the results change significantly in favor of nonstationarity if innovations are treated as draws from a symmetric stable Paretian distribution with infinite variance. Implementing Dickey and Fuller (1979, DF hereafter) unit root test in analyzing the stationarity of inflation rates in 40 countries, Ball et al. (1990) showed that nonstationarity null hypothesis couldn't be rejected for 38 out of 40 countries. Cook (2005) examined the integrated nature of U.S. inflation rates using DF and the rank-based DF unit root tests. The former test indicated the rejection of unit root null hypothesis, whereas the latter did not.

The second strand of studies applies panel unit root tests. Lee and Wu (2001), for instance, employed a bootstrap version of panel unit root test of Im et al. (2003) along with the seemingly unrelated approach of Taylor and Sarno (1998) in analyzing the stationarity of inflation rates in 13 OECD countries. They found overwhelming evidence of mean reversion. Culver and Papell (1997) analyzed the stationarity of inflation rates in 13 OECD countries using univariate and panel unit root tests. The results from the ADF, KPSS (Kwiatkowski et al., 1992) and the sequential break tests indicated that inflation rates followed unit root processes in most countries under study. However, the results from panel data provided strong evidence against nonstationarity. Another scholar, Osterholm (2004) took evidence of stationarity in inflation rates of USA by employing three panel unit root tests, namely IPS unit root test (Im et al., 2003), MADF test (Taylor and Sarno, 1998) and the Johansen (1988) likelihood ratio test. Tsong et al. (2012), who used panel unit root tests that allow for cross-sectional dependence and the covariate point optimal test, found strong evidence of mean reversion in the inflation rates of 19 OECD countries. Otero (2008), using univariate KPSS unit root test and Hadri (2000) panel unit root test, analyzed the stationarity of inflation rates in 13 OECD countries. The KPSS unit root test results generally indicated that inflation rates of 13 OECD countries could be best described as I (1) process, while Hadri panel unit root test implied stationarity.

The third strand of studies takes into account structural breaks while conducting unit root process. The univariate unit root tests that don't take into account structural breaks generally indicate inflation rates being nonstationary, whereas unit root tests with structural breaks generally imply inflation rates to be stationary. Caporale and Paxton (2013), for instance, investigated the stationarity of inflation rates for 5 Latin American countries using conventional ADF unit root test and Bai and Perron's (1998, 2003) multiple structural break test over the period 1980 to 2004. The results from the ADF unit root test indicated that stationarity held for only three countries. However, when structural breaks were allowed, all five countries turned to have

stationary inflation rates. Benati and Kapetanios (2002) found that unit root null could be strongly rejected for most series by employing a newly developed unit root test allowing for up to m structural breaks to the 23 inflation series from 18 countries. Levin and Piger (2003) applied both classical and Bayesian econometric methods to characterize the dynamic behavior of inflation rates for 12 industrial countries by estimating a univariate autoregressive model for each series and considering the possibility of structural break. The results indicated that allowing for a break induced stationarity in inflation rates.

In this strand, there are also studies using panel unit root tests with structural breaks. Narayan and Narayan (2010), for instance, examined whether inflation rates for 17 OECD countries could be modeled as a stationary process using univariate and panel KPSS unit root tests that allow for at most five structural breaks developed by Carrion-i Silvestre et al. (2005). The results from the KPSS univariate tests pointed out that inflation rates were stationary for 10 out of 17 countries. Furthermore, panel KPSS test rendered strong evidence favorable to the stationarity of inflation rates in panels consisting of countries, which were declared nonstationary by univariate tests. For comparison purposes, Narayan and Narayan (2010) also applied panel LM unit root test of Im et al. (2005) and got strong evidence against nonstationarity. Applying a battery of recent panel unit root tests on the data of Culver and Papell (1997), Basher and Westerlund (2007) found that stationarity of inflation rate held regardless of cross sectional dependence and structural change. Romero-Avila and Usabiaga (2009) utilized three panel unit root tests developed by Smith et al. (2004), Hadri (2000), and Carrion-i Silvestre et al. (2005) for 13 OECD countries. However, the results from these tests were contradictory. Hadri (2000) test indicated nonstationarity, whereas Smith et al. (2004) test provided little support to the presence of unit root. However, panel KPSS test gave strong evidence for joint stationarity.

In addition, in this strand, there is a group of studies applying univariate and panel LM unit root tests that allow for structural breaks as in our study. For example, Im et al. (2010) conducted univariate and panel LM unit root tests with two structural breaks for testing stationarity of inflation rates in 22 OECD countries. The results from univariate LM test indicated that 16 out of 22 countries had stationary inflation rates. They then employed panel LM unit root test and got results against nonstationarity. Another study implementing LM unit root test belongs to Lee and Chang (2008). They examined trend stationarity of inflation rates in 11 OECD and Asian countries and had results in support of stationarity.

In the fourth strand, studies conduct nonlinear approach in analyzing unit root process in inflation rates. Henry and Shields (2004), for instance, applied bootstrap Wald test approach developed by Caner and Hansen (2001) to the inflation rates of US, Japan, and U.K. The results implied that inflation rates in the U.K. and Japan were well described as a two-regime threshold unit root process. However, for the US, the threshold was not significant and shocks to inflation appeared to be infinitely

persistent. In this line, Ho (2009) accepted the unit root process for 19 OECD countries employing the nonlinear IV statistic that accounts for cross-sectional dependence. Another scholar, Zhou (2013) utilized unit root test that allows for nonlinearity to examine the stationarity of inflation rates in 12 European countries. Nonlinearity held for 8 out of 12 countries, and among them, 6 countries appeared to have a nonlinear stationarity in their inflation rates. Arize (2011) applied linear and nonlinear unit root tests, such as DF-GLS test developed by Elliot et al. (1996), conventional ADF unit root test, and also the KSS nonlinear unit root test of Kapetanios et al. (2003) to the inflation rates from 34 African countries. The DF-GLS, ADF and the KSS tests did not reject the null of nonstationarity in 17, 13, and 25 cases, respectively. Gregoriou and Kontonikas (2009) used ADF, Ng and Perron (2001, NP hereafter) and the non-linear ADF unit root tests (Kapetanios et al., 2003) to the inflation rates of 5 OECD countries that adopt inflation targeting. The results from the ADF unit root test showed that unit root null couldn't be rejected in all countries. Furthermore, NP test implied that unit root null could only be rejected at the 5% significance level in two countries. However, the results from the non-linear ADF test showed unit root process in all cases.

The fifth strand of studies analyzes the stationarity of inflation rates as a necessary first stage while testing the Fisher hypothesis. To name a few, Malliaropulos (2000), applying ADF and Zivot and Andrews (1992, ZA hereafter) unit root tests to the inflation rate in testing Fisher effect hypothesis for US, got results favorable to nonstationarity. However, ZA test indicated inflation rate being stationary. Furthermore, Atkins and Chan (2004) found that nominal interest rates and inflation rates in the Canada were stationary around a deterministic trend with two breaks in the framework of Fisher hypothesis. Furthermore, while searching the validity of Fisher effect in US, Crowder and Wohar (1999) applied the ADF unit root test and found that inflation rates being nonstationary. Clemente et al. (2004), who employed unit root test with breaks to the inflation rates of G7 countries in testing the validity of Fisher hypothesis, had evidence that inflation rate was represented as a broken trend stationary variable.

The sixth strand of studies utilizes fractional integration analysis to evaluate the stationarity of inflation rates. Employing fractional integration analysis to the inflation rates of 50 developing countries, Arize et al. (2005), for instance, found that inflation rate had nonstationarity in most countries under consideration. Bos et al. (1999) employed fractional integration test to the G7 inflation rates and got results in support of long memory or unit root process. In another paper, Bos et al. (2001), employing fractionally integrated moving average model, showed that US postwar inflation had a long memory, with an order of integration around 0.3. Gadea et al. (2004) utilized DF, PP, NP, and the KPSS unit root tests in searching the persistence of inflation rates by fractional integration. The results from the DF, PP, and the NP unit root tests rejected the unit root null hypothesis in all countries.

Finally, the last strand consists of studies analyzing inflation or price level convergence among a group of countries via examining stationarity of inflation differentials or testing for the cointegration relations among inflation rates. Convergence does hold if inflation rates are stationary. Holmes (2002), for example, using monthly data over the period 1972-1999, investigated strong and weak convergence among the major EU economies. The univariate ADF unit root test indicated that unit root null was rejected for only a few cases, whereas t bar test suggested by Im et al. (2003) implied stationarity as a whole. Drine and Rault (2006), who tested inflation convergence between Euro Zone and Central and East European, found strong evidence of convergence using panel LM unit root tests. Das and Bhattacharya (2004) attempted to examine price convergence across Indian regions employing panel unit root tests taking cross-sectional dependence into account. The results pointed to relative price level convergence among various regions in India.

3. Empirical Strategy

It is often argued that univariate unit root tests have low power with short time spans of data and, therefore, fail in rejecting the unit root null should be interpreted with caution. Due to this low power of conventional univariate unit root tests, there were failures to provide consistent answers to inflation dynamics in the previous studies. Researchers have generally pursued two general approaches to improve the low power of unit root tests. One of them is to allow for structural breaks in unit root tests. The second one is to design unit root tests in the panel framework. Panel unit root tests have an advantage since they allow researchers to use information in both cross-section and time series dimensions. However, "… any panel unit root test that does not allow for breaks suffers from the same loss of power that is experienced in the univariate unit root test setting" (Im et al., 2010). It is thus crucial to allow for structural breaks in both univariate and panel unit root tests.

In this line, Perron (1989) was the first to show that ignoring existing structural breaks when testing for a unit root can lead to a significant loss of power. Perron (1899) argued that the conventional ADF test has low power to reject the unit root null when the true data generating process is stationary around a broken linear trend. In this respect, if there is a structural break which is not modelled, then regardless of whether the null is a non-stationarity or stationarity, the results are likely to be spurious. Therefore, researchers have started to recognize the importance of allowing for breaks in unit root testing process by following Perron. Zivot and Andrews (1992), for instance, offered a unit root test that selected the break point endogenously from the data. In this line, Lumsdaine and Papell (1997) extended ZA test to allow for two breaks in both the level and trend of the series. However, all these endogenous tests were criticized since they have the limitation that the critical values are derived while assuming no break under the null hypothesis (Hooi and Smyth, 2005). In addition, when applying ADF-type endogenous break unit root tests, it might be concluded that a series is trend stationary while in fact it is non-stationary with breaks and this situation leads spurious rejections. Depending on these

criticisms, Lee and Strazicich (2003, 2004) developed a new unit root tests mentioned below. Thus, by allowing for structural breaks, some researchers have begun to conclude that inflation is stationary.

Given the aforementioned reasons and the following events, we decided to take into account the structural breaks in the stationarity analysis of inflation rates. World War I (1914 to 1918), the Great Depression (1929 to 1933), World War II (1941 to 1945), the first oil crisis (1973-1974), the Iranian revolution in 1978, a period of high inflation during the late 1970s, and a moderate economic growth and a low inflation period for western industrial countries in the late 1980s and early 1990s, the Iraq's invasion of Kuwait in 1990, and the Asian financial crisis over the time span 1997-1999 might have created some important breaks in the time-paths of inflation rate series, as noted by Lee and Chang (2008:2528).

3.1. Narayan and Popp's (2010) Unit Root Test

Narayan and Popp (2010) consider the following data generating process (DGP).

$$y_t = d_t + u_t \quad (1)$$

$$u_t = \rho u_{t-1} + \varepsilon_t, \quad (2)$$

$$\varepsilon_t = \Psi^*(L)e_t = A^*(L)^{-1}B(L)e_t, \quad (3)$$

where y_t is the inflation rate and has two components, a deterministic component, d_t and a stochastic component, u_t, with $e_t \sim iid\ (0, \sigma^2)$. It is also assumed that the roots of the lag polynomials, $A^*(L)$ and B (L), which are of order p and q, respectively, lie outside the unit circle.

They consider two different models. The first one (Model 1) allows for two breaks in level, and the second one (Model 2) allows for two breaks in both the level and trend. The deterministic component, d_t, is defined differently in these two models:

$$d_t^{M_1} = \alpha + \beta t + \Psi^*(L)(\theta_1 DU'_{1,t} + \theta_2 DU'_{2,t}) \quad (4)$$

$$d_t^{M_2} = \alpha + \beta t + \psi^*(L)(\theta_1 DU'_{1,t} + \theta_2 DU'_{2,t} + \gamma_1 DT'_{1,t} + \gamma_2 DT'_{2,t}) \quad (5)$$

with $DU'_{i,t} = 1$ if t>$T'_{B,i}$, zero otherwise, and $DT'_{i,t} = t - TB'_i$ if t>TB'_i, zero otherwise, i=1,2. Here, $T'_{B,i}$ indicates the break dates. The parameters, θ_i and γ_i, are the magnitude of the level and slope breaks, respectively. The inclusion of $\psi^*(L)$ in equations (4) and (5) allows breaks to occur slowly over time. In this case, as in Vogelsang and Perron (1998), it is assumed that the series responds to shocks to the trend function the way it reacts to shocks to the innovation process. This process is defined as the innovational outlier (IO) model.

When the structural models (1)-(5) are merged, the IO-type test regressions to test for the unit root hypothesis for M1 and M2 can be derived. The corresponding reduced forms of structural models represent the test regressions. The test equation for M1 is defined as:

$$y_t^{M_1} = \rho y_{t-1} + \alpha_1 + \beta^* t + \theta_1 D(T_B')_{1,t} + \theta_2 D(T_B')_{2,t} + \delta_1 DU'_{1,t-1} + \delta_2 DU'_{2,t-1} + \sum_{j=1}^{k} \beta_j \Delta y_{t-j} + e_t \quad (6)$$

where $\alpha_1 = \psi^*(1)^{-1}[(1-\rho)\alpha + \rho\beta] + \psi^{*'}(1)^{-1}(1-\rho)\beta$, $\psi^{*'}(1)^{-1}$ being the mean lag, $\beta^* = \psi^*(1)^{-1}(1-\rho)\beta$, $\phi = \rho - 1$, $\delta_i = -\phi\theta_i$ and $D(T_B')_{i,t} = 1$, if $(t = T'_{B,i} + 1)$, i= 1, 2.

Furthermore, the IO-type test regression for M2 is specified as such:

$$y_t^{M_2} = \rho y_{t-1} + \alpha^* + \beta^* t + \kappa_1 D(T_B')_{1,t} + \kappa_2 D(T_B')_{2,t} + \delta_1^* DU'_{1,t-1} + \delta_2^* DU'_{2,t-1} + \gamma_1^* DT'_{1,t-1} + \gamma_2^* DT'_{2,t-1} + \sum_{j=1}^{k} \beta_j \Delta y_{t-j} + e_t \quad (7)$$

where $\kappa_i = (\theta_i + \gamma_i)$, $\delta_i^* = (\gamma_i - \phi\theta_i)$, and $\gamma_i^* = -\phi\gamma_i$, i=1, 2. The test statistic of $\hat{\rho}$, denoted by $t_{\hat{\rho}}$, in equation (6) and (7) is used to test the unit root null hypothesis of ρ =1 against the alternative hypothesis of $\rho \prec 1$.

3.2. Lee and Strazicich (2003) LM Unit Root Test

In LM unit root test, the break points are endogenously determined from the data. Therefore, when the alternative hypothesis is true, spurious rejections are absent (Lee and Chang, 2008). Given that critical values are invariant to the location of the breaks, it is not necessary to simulate new critical values for different combinations of the break points, either (Jewell et al, 2003; Lee and Strazicich, 2003; Lee and Chang, 2008).

In contrast to ADF-type endogenous break tests, the LM unit root test is unaffected by breaks under the null hypothesis (Ozturk and Aslan, 2011) and can be explained by the following data generating process:

$$y_t = \delta' Z_t + X_t, \quad X_t = \beta X_{t-1} + \varepsilon_t \quad (8)$$

where y_t is the inflation rate and Z_t includes exogenous variables and ε_t is an error term that has classical features. In the case of two breaks in the level and trend (model C), Z_t takes the form of $[1, t, D_{1,t}, D_{2,t}, DT_{1,t}, DT_{2,t}]'$. $D_{j,t}$ and $DT_{j,t}$ are the dummy variables that represent structural shifts in level and trend of the series, respectively. In the case that TB_j refers to the dates of two breaks, for j=1,2, $D_{j,t} = 1$ if t> TB_j, zero otherwise. Also, $DT_{j,t}$ =t-TB_j if t>TB_j, zero otherwise. According to LM principle, a unit root test statistic can be obtained from the following regression:

$$\Delta y_t = \delta' \Delta Z_t + \varphi \tilde{S}_{t-1} + \sum_{p=1}^{k} B_p \Delta \tilde{S}_{t-p} + \varepsilon_t \quad (9)$$

where Δy_t and Δz_t are the first- difference values of y_t and z_t, respectively. The detrended value of y_t is represented by $\tilde{S}_{t-1}$. To correct for auto correlated errors,

augmented terms, i.e. $\Delta \tilde{S}_{t-p}$, are included in the regression as in the ADF type unit root test. ε_t is a stochastic disturbance term that has classical assumptions. The stationarity of y_t is of importance to decide whether φ is equal zero. Finally, a t test for $\varphi = 0$ in the regression (9) is used to calculate LM test statistic (denoted by $\tilde{\tau}$) for each country.

3.3. Im et al. (2010) Panel LM Unit Root Test

In addition to univariate tests, we also implement panel LM unit root tests developed by Im et al. (2010) that take account of possible existence of heterogeneous breaks in both the level and trend of the series. Furthermore, it corrects for the presence of cross-correlations in the innovations of the panel by employing the Cross-Sectionally Augmented (CA) procedure of Pesaran (2007).

$$\tilde{S}^{*} = \begin{cases} \frac{T}{TB_1}\tilde{S}_t & \text{for } t \le TB_1 \\ \frac{T}{TB_2 - TB_1}\tilde{S}_t & \text{for } TB_1 < t \le TB_2 \\ \frac{T}{T - TB_2}\tilde{S}_t & \text{for } TB_2 < t \le T \end{cases} \quad (10)$$

The asymptotic distribution of Panel LM test statistic with trend shift depends on the nuisance parameters and ignorance of nuisance parameters dependency will be subject to crucial size distortion. However, Im et al. (2010) remove this dependency of test statistic on the nuisance parameter with the following transformation by depending on Lee and Strazicich (2009).

After that, they replace $\tilde{S}_{t-1}$ with $\tilde{S}^{*}_{t-1}$ in the testing regression (9) and employ the regression for each cross-section unit. Using average of the test statistics, the following t-bar statistics can be computed. Then, panel LM statistic can be computed as the standardized statistic of this average test statistic.

$$\bar{t} = \frac{1}{N}\sum_{i=1}^{N}\tilde{\tau}^{*}_{i} \quad (11)$$

The expected value and variance of $\bar{t}$ under the null hypothesis ($H_0 : \varphi_i = 0$, for all i) are denoted by $E(\bar{t})$ and $V(\bar{t})$. Finally, using estimated values of the average of the means and variances of $\bar{t}$, $\tilde{E}(\bar{t})$ and $\tilde{V}(\bar{t})$, the following panel LM test statistic that has standard normal distribution can be computed.

$$\text{LM}(\tilde{\tau}^{*}) = \frac{\sqrt{N}\,[\bar{t} - \tilde{E}(\bar{t})]}{\sqrt{\tilde{V}(\bar{t})}} \quad (12)$$

4. Data and Main Results

4.1. Data

We use yearly consumer price index (base=2005) for 11 Mediterranean countries, namely Spain, France, Italy, Greece, Malta, Turkey, Syria, Cyprus, Israel, Egypt, and Morocco over the period 1960-2011. The inflation rate is calculated as the logarithmic first difference of consumer price index. The data come from the World Development Indicators as published by the World Bank (2012). The countries under study and time span are dictated by data availability.

4.2. The Results of Conventional Unit Root Tests

In this study, following Im et al. (2010), we allow for two breaks in both the level and trend of the inflation rate series. Therefore, Model C in Lee and Strazicich (2003) and Model 2 in Narayan and Popp (2010) are used. Many scholars also (see Sen, 2003; Lee and Chang, 2008; Lean and Smyth, 2012) indicate that the use of Model C is superior to Model A.

With respect to unit root tests used, we first employ univariate unit root tests with breaks developed by Lee and Strazicich (2003) and Narayan and Popp (2010). In addition, as a benchmark, we report the results of conventional unit root tests, namely ADF (Dickey and Fuller, 1979, 1981), SP (Schmidt and Phillips, 1992), PP (Phillips and Perron, 1988), DF-GLS (Elliot et al., 1996), and MZ_{α}^{GLS} (Ng and Perron, 2001) for no-break case. After that, we conduct panel LM unit root test of Im et al. (2010) for both the level and trend shift case. Furthermore, as a benchmark, we report the results of panel LM unit root test without breaks.

Table 1: Conventional Univariate Unit Root Test Results

Country	ADF	PP	DF-GLS	MZ_{α}^{GLS}	SP($\widetilde{\tau}$)
Spain	-2.627(0)	-2.712 (4)	-1.908 (0)	-5.853(4)	-1.775(3)
France	-2.123(0)	-2.233 (2)	-1.698(0)	-5.259(2)	-1.708(6)
Italy	-1.999(0)	-2.115 (3)	-1.446(4)	-5.234(3)	-1.284(4)
Greece	-1.839(0)	-1.663(7)	-1.670(0)	-3.575(7)	-1.711(0)
Malta	-3.052(1)	-3.781** (0)	-3.040(1)	-17.613**(0)	-3.766***(0)
Turkey	-1.424(0)	-1.262 (5)	-1.498(0)	-3.784(5)	-1.738(0)
Syria	-2.994(0)	-3.0315 (1)	-2.994*(0)	-13.379(1)	-2.097(0)
Cyprus	-3.419*(0)	-3.271* (2)	-1.728(2)	-10.193(2)	-1.538(2)
Israel	-1.781(2)	-2.097 (6)	-1.679(2)	-7.182(6)	-1.669(2)
Egypt	-2.445(1)	-2.773 (2)	-2.212(1)	-9.741(2)	-1.918(8)
Morocco	-1.745(2)	-3.348* (3)	-1.642(2)	-12.896(3)	-2.081(1)

Notes: All tests include constant and trend terms. The numbers in parentheses are the lag orders in the ADF and DF-GLS tests. The lag parameters are selected based on the modified akaike information criteria (MAIC). Newey-West correction as a selection of truncation lags in parentheses is used for the PP and MZ_{α}^{GLS} unit root tests. The numbers in parentheses in SP test are the lagged augmentation terms to control for autocorrelation. The critical values of SP test are -3.63, -3.06 and -2.77 for the 1%, 5%,, and 10% significance levels, respectively. ***, **, and * indicate significance at the 1%, 5%, and 10% significance levels, respectively.

According to the results of the SP and MZ_{α}^{GLS} unit root tests, it seems that all countries, except for Malta, have non-stationary inflation rates series. Furthermore, ADF and DF-GLS tests indicate stationarity for only Cyprus and Syria, respectively. Finally, PP test points inflation rates of Malta, Cyprus, and Morocco to be stationary. However, allowing for structural breaks might change all these results due to aforementioned reasons. We thus conduct Narayan and Popp (2010) and Lee and Strazicich (2003) unit root tests.

4.3. Narayan and Popp's (2010) Unit Root Test Results

The results from Table 2 indicate that 5 out of 11 countries, namely France, Malta, Syria, Egypt, and Morocco have stationary inflation rates at different significance levels. Concerning the first and second break dates, it seems that the first and second oil crises led to crucial structural shifts in the trend-paths of inflation rate series.

Table 2: The results of Narayan and Popp's (2010) test

Country	Test statistic	$\hat{p}$	TB_1, TB_2
Spain	-3.306	2	1965,1976
France	-4.980*	0	1972,1981
Italy	-3.538	0	1976,1979
Greece	-0.7639	0	1973,1988
Malta	-7.204***	0	1978,1981
Turkey	-0.881	0	1979, 1992
Syria	-4.850*	0	1971,1985
Cyprus	-3.719	0	1973,1980
Israel	-1.905	1	1982,1984
Egypt	-5.089*	0	1967, 1994
Morocco	-6.706***	0	1972,1985

Notes: The critical values for endogenous two breaks test are -5.949, -5.181, and -4.789 at the 1%, 5%, and 10% significance levels, respectively for model 2 (M2), with 50.000 replications, when T=50. ***, **, and * display 1%, 5%, and 10 % significance levels, respectively. We use 3 lagged first difference of y_t to control for autocorrelation in equation (7). Also, 10% of data is trimmed for each country.

4.4. Lee and Strazicich's (2003) Unit Root Test Results

While conducting LM unit root test, depending on Lee and Strazicich (2003, 2004) and Lee and Chang (2008), we first start from a maximum of k=8 lagged terms to determine the number of lagged augmentation terms to control for autocorrelation As such, the procedure looks for significance of the last augmented term. We then use the 10% asymptotic normal value of 1.645 on the t-statistic of the last first differenced lagged term. After determining the optimal k at each combination of two break points, the breaks where the endogenous two breaks LM t-test statistic is at a minimum are determined. We thus examine each possible combination of two break points over the time interval (0.1T, 0.9T), while eliminating the endpoints. Here, T stands for the sample size. Table 3 reports the results of LM unit root test of Lee and Strazicich (2003).

Table 3: Individual LM Unit Root Test Statistics

Country	$\hat{k}$	$T\hat{B}$	Individual test statistics	$\hat{d}_1$	$\hat{d}_2$
Spain	8	1983	-4.962**	-0.002(-0.410)	0.021(4.098)***
France	2	1971, 1983	-6.461***	0.030(4.525)***	-0.051(-6.036)***
Italy	4	1971, 1983	-7.509***	0.063(6.184)***	-0.089(-7.469)***
Greece	1	1970, 1984	-5.049	0.049(3.355)***	-0.090(-5.209)***
Malta	0	1980, 1984	-7.125***	-0.022(-1.890)*	0.044(3.625)***
Turkey	6	1978, 1993	-6.215**	-0.244(-5.520)***	0.274(4.937)***
Syria	7	1981, 1994	-6.540***	0.139(4.568)***	-0.165(-5.100)***
Cyprus	8	1972, 1980	-8.641***	-0.085(-7.994)***	0.090(8.697)***
Israel	3	1980, 1985	-8.655***	0.543(5.616)***	-0.355(-4.518) ***
Egypt	7	1983, 1996	-6.898***	0.113(5.401)***	-0.075(-4.650)***
Morocco	5	1971, 1984	-6.606***	0.082(4.927)***	-0.099(-6.346)***

The critical values for the univariate LM unit root test with two breaks

	λ_2		
λ_1	0.4	0.6	0.8
0.2	-6.16, -5.59, -5.27	-6.41, -5.74, -5.32	-6.33, -5.71, -5.33
0.4	- - -	-6.45, -5.67, -5.31	-6.42, -5.65, -5.32
0.6	- - -	- - -	-6.32, -5.73, -5.32

Notes: $\hat{k}$ is the optimal number of lagged first-differenced terms included in the unit root test to correct for serial correlation. $T\hat{B}$ is the estimated break dates. $\hat{d}_1$ and $\hat{d}_2$ are the coefficients of dummy variables under the unit root null in Lee and Strazicich (2003). The numbers of parentheses are the t statistics. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

As shown in Table 3, all countries have stationary inflation rates, except for Greece. When we include two structural breaks in both the level and trend of the inflation rates, the results change from non-stationary to stationary. We first employ LM unit test with two breaks and examine the significance of dummy coefficients depending on the conventional t statistics. If one of them appears insignificant at the 10% significance level or better, we then conduct LM unit root test with one break as in Lee and Strazicich (2004). The results imply that two dummy coefficients are significant in nearly all countries, except for Spain. Given that only one dummy coefficient is significant, we report the results of LM unit root test with one break for Spain.

Regarding the break dates, the first breaks generally correspond to the first half of 1970s in five countries, namely France, Italy, Greece, Cyprus, and Morocco. The international monetary crisis in 1971, the first oil crisis in 1973 and 1974 and the second oil crisis in 1978 caused crucial structural shifts in the trend ways of inflation rate series. The second break dates generally coincide with the early 1980s. The oil glut that started in the early 1980s as a result of slowed economic activity in industrial countries, the international debt crisis in 1982 and the European currency crisis in 1992-1993 seem to be the most important events that caused significant breaks in the time paths of inflation rates.

4.5. Im et al. (2010) Panel LM Unit Root Test Results

After employing univariate unit root tests, we can report the results of panel LM unit root tests. As a benchmark, we first employ panel LM unit root test without break. As seen in Table 4, when structural breaks are ignored, panel LM test indicates inflation rate being non-stationary. However, when two breaks are allowed in both the level and trend of the inflation rates as in Im et al. (2010), we take two different results depending on two test statistics. The first statistic does not correct for cross-correlations, whereas the second one does by applying Pesaran's CA procedure to the panel LM test. According to the results of panel LM unit root test developed by Im et al. (2010), when cross-sectional dependence isn't corrected, inflation rate series follows a stationary process as a whole. However, it seems nonstationary when cross correlation is taken into account.

Table 4: Panel LM Unit Root Tests Results

Panel LM Test Statistic without break	-0.110
Panel LM Test Statistic with two breaks	-3.595***
Panel LM Test CA Statistic with two breaks	-0.994

Notes: The panel LM test employed here belongs to Im et. al (2010). The critical values (with and without breaks) are −2.326, −1.645, and −1.282 at the 1%, 5%, and 10 % significance levels, respectively. *** denotes significance at the 1% level.

To sum up, depending on the results mentioned above, it could be asserted that structural breaks must be taken into account in ascertaining integrational properties of inflation rate. The results from unit root tests without breaks generally indicate that inflation rates follow unit root process. However, allowing for structural breaks seems to change the results from nonstationary to stationary in most cases.

5. Conclusion

In this study, we employ Lee and Strazicich (2003), Narayan and Popp (2010), and Im et al. (2010) unit root tests that allow for structural breaks in examining the integrational properties of inflation rates in 11 Mediterranean countries. For the purpose of comparison, we first apply conventional univariate unit root tests without breaks. The results from the univariate unit root tests without breaks point out that most countries have nonstationary inflation rates. However, the univariate LM unit root tests developed by Lee and Strazicich (2003, 2004) imply that stationarity does hold for 10 individual countries, whereas Narayan and Popp (2010) unit root test shows inflation rates being stationary for 5 countries. When we conduct Im et al. (2010) panel LM unit root test that allows for breaks in both the level and/or trend of the series, inflation rate seems stationary when cross-sectional correlation is ignored, while it implies nonstationarity in the case that cross-sectional dependence is corrected. Depending on those results from the univariate and panel unit root tests with and without breaks, it can be asserted that individual inflation rate series generally show mean reversion (trend stationarity) when structural breaks are

allowed. However, they follow unit root process without breaks. In addition, the results favorable to stationarity suggest the following crucial policy implications.

First, in modeling inflation rates, it is important to take account of unit root tests with structural breaks. The findings confirm that allowing for breaks leads us to accept the stationarity in the most Mediterranean countries under consideration. This finding has implications for modeling inflation rates with other macro variables such as GDP, stock prices, money demand etc. For instance, conventional cointegration analysis may not be appropriate to test money demand models, Fisher effect and the expectation-augmented Phillips-curve models.

Second, the results imply that shocks to inflation have no long lasting effects on the inflation rates of 11 Mediterranean countries. Therefore, monetary authorities of these Mediterranean countries would less costly implement disinflationary policies than those of the countries with nonstationary inflation rates. Furthermore, trend stationarity of inflation rate indicates that inflation rate will return to its trend path over time and it might be possible to forecast future movements in the inflation rate based on its past behavior.

Third, since inflation rates are stationary around a breaking trend, only large shocks such as government policies aimed at changing the fundamentals will have lasting effects on the inflation path, but shocks only cause short-run deviations around a deterministic trend. In this case, aggregate demand policies may not be over-implemented (Lee and Wu, 2001) and administrative policy of a government should not be to adopt excessive interfering targets. As stated by Lee and Chang (2008), models ignoring breaks in the trend path of inflation cannot avoid the wasted costs of interference, which can also increase fluctuations in other macroeconomic variables.

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