Nonlinear trend stationarity of the Real Exchange Rates of Mediterranean countries

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Abstract

The aim of this article is to provide additional evidence on the fulfillment of the Purchasing Power Parity hypothesis in the so-called Mediterranean countries. In order to test for the empirical validity of such hypothesis, we have applied two types of unit root tests. The first group is due to Bierens (1997) who generalizes the alternative hypothesis to nonlinear trend stationariry and, the second is the Leybourne, Newbold and Vougas (1998) approach that uses a nonlinear specification for the intercept and slope in order to detrend the series. The results suggest a pretty weak evidence in favour of the Purchasing Power Parity hypothesis for this group of countries.

Classification J.E.L.: C22, F31

Key words: purchasing power parity, real exchange rate, unit roots, structural change, nonlinearity.

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1 Introduction

During the last decades, a number of authors have studied whether Purchasing Power Parity (hereafter PPP) theory holds. The concept was introduced by Cassel in 1918. Since then, its empirical validity has been tested for different time periods, country-groups and using a variety of econometric techniques. The absolute version of PPP establishes that prices in different countries have to be equal when measured in a common currency, i.e. the nominal exchange rate and the price ratio should share a co-movement along time (cointegrate or share deterministic trends depending on the order of integration of them). This is equivalent to saying that the real exchange rate, defined as

$$Q_t = \frac{E_t P_t}{P_t^*} \tag{1.1}$$

is equal to unity, where Q_t is the real exchange rate, E_t the nominal exchange rate[1], and P_t^* and P_t are respectively the foreign and domestic price indices. Another version less restrictive is known as relative PPP, and implies that the real exchange rate is a constant different to one. PPP holds when the real exchange rate is stationary so that shocks have only transitory effects.

The empirical literature about PPP is very wide. Many different techniques have been used to test for PPP fulfilment, from Ordinary Least Squares and Instrumental Variables (Frenkel, 1978 and Krugman, 1978) to cointegration (Taylor, 1988, 1992; Johansen and Juselius, 1992; and Doganlar, 1999) and nonlinear techniques (Dixit, 1989; Moosa, 1994; Obstfeld and Taylor, 1997; and Sarno, 2000). Although the empirical literature is vast, the evidence is far from conclusive.

Perron and Phillips (1987) and West (1988), among others, suggest that traditional unit root tests may suffer from lack of power when there is a misspecification of the deterministic time trend. If the variables present structural changes, these tests may conclude that the series analyzed are I(1) when in fact they are stationary around a deterministic time trend or broken time trend (Rappoport and Reichlin, 1989 and Perron, 1989, 1990).

Bearing this consideration in mind some authors have applied unit root tests with structural changes in order to test for the order of integration of real exchange rates. Following this approach, the results obtained by Dropsy (1996), Parkes and Savvides (1999), and Montañés and Clemente (1999) support PPP.

[1] Units of foreign currency for a unit of domestic currency.

A broken time trend is a particular case of a nonlinear time trend. Thus, traditional unit root tests, even with structural changes, may incorrectly conclude that the series are I(1) when in fact they are stationary around a nonlinear trend (Bierens, 1997). For instance, Michael, Peel and Nobay (1997) provide proof of the fact that the ADF test applied to a linear model may reject the PPP hypothesis if the DGP is nonlinear. Additionally, Taylor and Peel (2000, p. 35) justify the use of nonlinear modelling claiming that *"As the exchange rate becomes increasingly misaligned with the economic fundamentals, however, one might expect that the pressure both from the market and from policy markets to return to the exchange rate to the neighbourhood of fundamental equilibrium would become increasingly strong".*

In this paper we study PPP fulfilment for the so-called "Mediterranean countries" (Algeria, Cyprus, Egypt, Israel, Jordan, Malta, Morocco, Syria, Tunisia and Turkey). Unlike other papers, we concentrate on the real exchange rate against the European Union (EU). There are two reasons for the adoption of this approach. First, they have a commitment with the EU for the creation of a Free Trade Area by 2010, on the basis of the Euroministerial Conference held in Barcelona in 1995. Since PPP can be understood as a measure of economic integration, it may be worthy to test for such relationship between both zones in order to understand their degree of economic integration. Second, former studies such as Sarno (2000), and Camarero, Cuestas and Ordóñez (2006) have highlighted that PPP does not hold for these countries.

The aim of this paper is to test whether the case of nonlinear deterministic components is a better statistical characterization of the long run behaviour of the real exchange rates for this group of countries. In order to do so we apply Bierens (1997) unit root tests that generalize the alternative hypothesis to stationarity around a nonlinear trend. Also, we have applied Leybourne, Newbold and Vougas (1998) unit root test allowing for a smooth transition from one trend function to another. The difference between these two approaches is that Bierens (1997) approximates the nonlinear deterministic trend by Chebishev polynomials, whereas Leybourne et al. (1998) allow for smooth transition not only in the trend but also in the intercept[2].

To the best of our knowledge, there is no empirical work that analyses this issue using the Bierens tests. Nevertheless, Sollis (2005) applies the Leybourne, Newbold and Vougas (1998) approach to test for PPP for a number of countries against the US dollar finding that this relationship holds for many of them.

[2]See Michael, Nobay and Peel (1997) for the adequacy of smooth transition models vs. threshold models to characterize the long run behaviour of real exchange rate.

The remainder of this paper is organized as follows. In the next section, we summarize Bierens (1997) and Leybourne, Newbold and Vougas (1998) unit root tests. In the third section we present the results of such tests applied to PPP in the Mediterranean countries and, finally, the last section summarizes with the main contributions of this paper.

2 Econometric Methodology

2.1 Unit root test with drift versus non-linear trend stationarity

Bierens (1997) develops a test for the null hypothesis of unit root with drift:

$$H_0 : z_t = z_{t-1} + \mu + u_t \tag{2.1}$$

where μ is the constant term and u_t is a stationary $AR(p)$ process. The author proposes to test for this null hypothesis against the alternative of nonlinear trend stationarity:

$$H_1 : z_t = g(t) + u_t \tag{2.2}$$

where $g(t)$ is a nonlinear trend function. The test is based on the following ADF-type auxiliary regression model, where the nonlinear trend $g(t)$ is approximated by a linear combination of Chebishev polynomials $P_{t,n}^m = (P_{0,n}^*(t), P_{1,n}^*(t), ..., P_{m,n}^*(t))^T$:

$$\Delta z_t = \alpha z_{t-1} + \sum_{j=1}^{p} \phi_j \Delta z_{t-j} + \theta^T P_{t,n}^m + \varepsilon_t \tag{2.3}$$

where n is the number of usable observations and

$$P_{0,n}^*(t) = 1,$$

$$P_{1,n}^*(t) = \frac{t-(n+1)/2}{\sqrt{(n^2-1)/12}},$$

$$P^*_{2k,n}(t) = \frac{P_{2k-1,n}(t) - \alpha_{k,n} - \sum_{j=1}^{k-1} \beta_{k,j,n} P_{2j-1,n}(t) - \gamma_{k,n}(t/n)}{c_{k,n}},$$

where $k = 1, 2, .., [n/2]$, $\alpha_{k,n}$, $\beta_{k,j,n}$ and $\gamma_{k,n}$ are the coefficients of the least square regression of $P_{2k-1,n}(t)$ upon 1, $P_{2j-1}(t)$, $j = 1, ..., k-1$, and t/n respectively and $c_{k,n}$ is a constant used to get $(1/n)\sum_{t=1}^{n}[P^*_{2k,n}(t)]^2$=1. Finally,

$$P^*_{2k+1,n}(t) = P_{2k,n}(t).$$

Bierens (1997) proposes an F test to test the joint hypothesis that, under H_0, α and the last m components of θ are zero:

$$\hat{F}(m) = \frac{\left(\sum_{t=1}^{n} \hat{\varepsilon}^2_{0,t} - \sum_{t=1}^{n} \hat{\varepsilon}^2_{m,t}\right)/(m+1)}{s^2} \tag{2.4}$$

where

$$s^2 = \frac{1}{n-p-m-1} \sum_{t=1}^{n} \hat{\varepsilon}^2_t. \tag{2.5}$$

Since this test does not follow a standard F distribution, Bierens (1997) provides the distribution fractiles based on Monte Carlo simulation.

In addition, the author develops a model-free unit root test $\widetilde{T}(m)$, given that for the F test it is necessary to choose the lag length p in the auxiliary regression and the results may be sensitive to this choice. The model-free unit root test is based on the following regression:

$$\Delta z_t = -\rho z_{t-1} + \lambda_0 + \rho\lambda_1 t + f(t) + u_t \tag{2.6}$$

where ρ lies in the interval $\{0, 1\}$, $f(t)$ is a non-constant deterministic function of time such that $\lim_{n\to\infty}(1/n)\sum_{t=1}^{n} f(t) = 0$, $\lim_{n\to\infty}(1/n)\sum_{t=1}^{n} tf(t) = 0$, and u_t is a zero-mean process that follows the functional central limit theorem. The null hypothesis of a unit root is formulated as:

$$H_0 : \rho = 0, f(t) \equiv 0, \tag{2.7}$$

There are two alternative hypothesis. The first one is linear trend stationarity

$$H_1^{lin} : \rho = 1, f(t) \equiv 0, \tag{2.8}$$

whereas the second alternative is nonlinear trend stationarity

$$H_1^{nlin} : \rho = 1. \tag{2.9}$$

In case of rejection of the null, in order to distinguish between stationarity around a linear or around a nonlinear trend, Bierens (1997) designs the $\widetilde{T}(m)$ test:

$$\widetilde{T}(m) = \frac{\left(\sum_{t=1}^{n} \Delta z_t P_{t,n}^{(1,m)} - \hat{\xi}_1 P_{n+1,n}^{(1,m)} - \hat{\xi}_2 P_{1,n}^{(1,m)}\right)^T \left(\sum_{t=1}^{n} \Delta z_t P_{t,n}^{(1,m)} - \hat{\xi}_1 P_{n+1,n}^{(1,m)} - \hat{\xi}_2 P_{1,n}^{(1,m)}\right)}{(1/n) \sum_{t=1}^{n} \left(z_t - \widetilde{\theta}^{(m)T} P_{t,n}^{(m)}\right)} \tag{2.10}$$

where $\hat{\xi}_1$ and $\hat{\xi}_2$ are the least squares coefficients from regressing $\sum_{t=1}^{n} \Delta z_t P_{k,n}^*(t)$ on $P_{k,n}^*(n+1)$ and $P_{k,n}^*(1)$, and $P_{t,n}^{(i,m)} = (P_{i,n}^*(t), ..., P_{m,n}^*(t))^T$. As this test does not have a standard limiting distribution, Bierens (1997) provides the most important fractiles of the distribution for $m = 3, ..., 20$. Left side rejection would imply linear trend stationarity whereas right side rejection implies nonlinear trend stationarity.

Thus, the main advantage of $\widetilde{T}(m)$ over $\hat{F}(m)$ is that the former permits the distinction between stationarity around a linear and nonlinear trend. However, in $\widetilde{T}(m)$ we assume that the lag length of the auxiliary regression (2.3) is zero[3].

2.2 Smooth transition regression models

Leybourne, Newbold and Vougas (1998) propose a unit root test applied to three logistic smooth transition regression models in an attempt to model structural change as a smooth transition between different regimes rather than an instantaneous structural break:

$$y_t = \alpha_1 + \alpha_2 S_t(\gamma, \tau) + \nu_t \tag{2.11}$$

$$y_t = \alpha_1 + \beta_1 t + \alpha_2 S_t(\gamma, \tau) + \nu_t \tag{2.12}$$

$$y_t = \alpha_1 + \beta_1 t + \alpha_2 S_t(\gamma, \tau) + \beta_2 t S_t(\gamma, \tau) + \nu_t \tag{2.13}$$

[3] The ADF-type regression becomes a DF-type regression.

where ν_t is a stationary process with zero mean and $S_t(\cdot)$ is the nonlinear function which controls the transition between regimes. The authors define $S_t(\cdot)$ as a logistic smooth transition function for a sample size T:

$$S_t(\gamma, \tau) = (1 + exp\{-\gamma[t - \tau T]\})^{-1}, \gamma > 0. \tag{2.14}$$

Model A (equation (2.11)) approximates the nonlinear deterministic component as a transition in the intercept of a non-trending series, Model B (equation (2.12)) does it by a transition in the intercept of a trending time series and, finally, Model C (equation (2.13)) uses a transition in the intercept and slope of a trending series (Leybourne et al., 1998).

The above mentioned models can be used to formally test the order of integration of the variables, taking into account the different specification of the deterministic component,

$$\begin{aligned} H_0 &: y_t = \mu_t, \mu_t = \mu_{t-1} + \varepsilon_t, \mu_0 = \psi \\ H_1 &: \text{Model A, Model B or Model C} \end{aligned}$$

and

$$\begin{aligned} H_0 &: y_t = \mu_t, \mu_t = \kappa + \mu_{t-1} + \varepsilon_t, \mu_0 = \psi \\ H_1 &: \text{Model B or Model C} \end{aligned}$$

where ε is assumed to be an I(0) process with zero mean.

To apply the unit root tests Leybourne et al. (1998) propose a procedure that involves two steps. In the first step, the models A, B or C are estimated by Nonlinear Least Squares and the residuals saved. In the second step, the DF test is applied to the residuals. The null distributions of the tests are approximated using Monte Carlo simulation methods.

3 Empirical Results

3.1 Data

The data used in the empirical analysis are the log of nominal effective exchange rates (e_t), defined as the price of the national currency in terms of the foreign currency and the log of the price differential relative to the EU (p_t), computed as national Consumer Price Index minus foreign prices. The log of the real effective exchange rate is then calculated as $e_t + p_t$. The data have been taken from the *International Financial Statistics*, IMF. The nominal effective exchange rates and foreign prices have been calculated specifically for each country, using as weights the proportion of trade with their respective EU trade partners. These weights have been obtained from the *Direction of Trade Statistics Yearbook,* IMF. The frequency of the data is quarterly and spans from 1979:1 to 2002:4. In the case of Tunisia the sample starts in 1987:3.

In figure 1 we display the graphs of the series of the real exchange rates for the Mediterranean countries. The graphical analysis shows that the path of the real exchange rates does not look like following a linear trend, hence suggesting the possibility of nonlinear deterministic components in the series.

3.2 Nonlinear unit root tests

We analyse first the results of the Bierens' (1997) tests. As described above, the $\hat{F}(m)$ test is calculated from the ADF regression (2.3) where the lag length p has been chosen using the Akaike information criterion (AIC). In addition, we also apply the $\widetilde{T}(m)$ test that is based on regression (2.6). In this case it is not necessary to choose the lag length, as $p = 0$ by definition. Bierens (1997) shows that both tests suffer from important size distortions. Accordingly we have computed the critical values using Monte Carlo simulations based on 10,000 replications of a Gaussian $AR(m)$ process for Δx_t. The parameters and error variances are equal to the estimated $AR(m)$ null model, where the order p of the ADF regression has been selected by the AIC and the initial values are taken from the actual data. In table 3, we present the results of the $\hat{F}(m)$ and $\widetilde{T}(m)$ tests. As pointed out by Bierens (1997), there is not a unique way of choosing the value of m: a low value could be not enough to approximate the nonlinear trend, whereas a large value for m might imply low power because of the estimation of redundant parameters. For that reason, table 3 presents the Bierens' test for different orders of m.

The results for the $\hat{F}(m)$ test suggest that the null of unit root is rejected for Algeria and Egypt (for large values of m), as well as for Morocco (in this case for a low length of m). Although stationarity might be accepted for these three countries, the $\hat{F}(m)$ test does not allow us to distinguish the alternative hypothesis. There are three possibilities:mean stationarity, linear trend stationarity and nonlinear trend stationarity. To complement the analysis, the $\widetilde{T}(m)$ test statistic is also computed. The results are similar to those obtained with the $\hat{F}(m)$ test. Thus we do not fail to reject the null of unit root for Algeria, Egypt and Morocco, when the alternative is nonlinear trend stationarity (right-sided rejection).

In table 4 we present the results of the ADF test for the residuals of the STR models[4]. As pointed out by Taylor and Peel (2000), a transition function like (2.14) implies asymmetric behaviour of the modelled variable, being inappropriate for modelling exchange rate movements. Instead, we use an exponential smooth transition (ESTR) function since the adjustment towards equilibrium is symmetric and does not depend on the sign of the shock. The ESTR function is given by:

$$S_t(\gamma, \tau) = (1 - exp\{-\gamma^2[t - \tau T]^2\}), \gamma > 0. \tag{3.1}$$

The critical values for the DF and ADF tests applied to the residuals of the auxiliar nonlinear regressions are presented in tables 1 and 2, and have been obtained by Monte Carlo simulations over 20,000 replications. The null DGP is been specified as:

$$y_t = \mu_t, \mu_t = \mu_{t-1} + \varepsilon_t, \varepsilon_t \sim NID(0, 1).$$

The Nonlinear Least Squares estimation was computed using the optimization algorithm in the OPTMUM subroutine library of GAUSS. The initial values where obtained using the SIMPLEX algorithm.

According to the results in table 4 and in contrast with previous tests, it is no possible to reject the null of unit root for any of the countries.

This weak evidence in favour of the PPP hypothesis contrasts with the findings of Camarero, Cuestas and Ordóñez (2006) who find evidence of stationarity in half of the Mediterranean countries by using unit root tests with structural changes and for TAR models. This joint evidence suggests that the evidence in favour of PPP stationarity improves once the deterministic structural changes are characterized by instanteneous structural breaks (TAR models) rather than by a smooth transition between different regimes over time.

4 Conclusions

Trying to contribute to the vast literature on PPP, in this paper we have analysed the empirical fulfilment of PPP in the Mediterranean countries using two unit root tests that take into account the possibility of nonlinearities in the deterministic components.

[4]The results for Cyprus and Israel does not evidence the existence smooth transitions either in the intercept or in the slope.

Our results find little evidence of PPP holding in the Mediterranean countries in contrast with a previous study on this topic. On the one hand, by using Bierens' unit root tests PPP holds for Algeria, Egypt and Morocco. On the other hand, by applying Leybourne, Newbold and Vougas (1998) approach, there is no evidence on PPP fulfillment.

Our conclusion is twofold. First, a proper statistical characterization of the deterministic componentes is of crucial importance when testing for PPP to hold. Second, the use of smooth transition models as a means of representing deterministic structural changes in real exchange rates appears to be less appropriate than modelling structural change by instantaneous breaks.

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Table 1: Null critical values for unit root tests against stationarity around a smooth transition: model (A) with smooth drift

$n = 25$	$k = 0$	$k = 1$	$k = 2$	$k = 3$	$k = 4$	$k = 5$
0.100	-4.7516	-4.4666	-4.0151	-3.8256	-3.4610	-3.2961
0.050	-5.2002	-4.8794	-4.3884	-4.2157	-3.8127	-3.6664
0.010	-6.1531	-5.7955	-5.2508	-5.0382	-4.6874	-4.5617
$n = 50$						
0.100	-4.4654	-4.3656	-4.2358	-4.1932	-4.0822	-4.0165
0.050	-4.8021	-4.7295	-4.5817	-4.5043	-4.4053	-4.3192
0.010	-5.5202	-5.3509	-5.2338	-5.0958	-5.0250	-4.9604
$n = 100$						
0.100	-4.4288	-4.3332	-4.0868	-3.9851	-3.8146	-3.7174
0.050	-4.8161	-4.7106	-4.4470	-4.3107	-4.1285	-4.0237
0.010	-5.6059	-5.3661	-5.1243	-4.9559	-4.7413	-4.6694
$n = 200$						
0.100	-4.2374	-4.2149	-4.1269	-4.1038	-4.0364	-4.0321
0.050	-4.5170	-4.4946	-4.4394	-4.3783	-4.3157	-4.3106
0.010	-5.1621	-5.0714	-5.0571	-5.0133	-4.8653	-4.8865

Note: Nominal sizes 0.10, 0.05 and 0.01. k is the order of lags in the ADF regression.

Table 2: Null critical values for unit root tests against stationarity around a smooth transition: model (C) with smooth drift and trend

$n = 25$	$k = 0$	$k = 1$	$k = 2$	$k = 3$	$k = 4$	$k = 5$
0.100	-4.9454	-4.6414	-4.1389	-3.9724	-3.5558	-3.4099
0.050	-5.4175	-5.0999	-4.5700	-4.3857	-3.9443	-3.8390
0.010	-6.3450	-6.1386	-5.4386	-5.2056	-4.7268	-4.6855
$n = 50$						
0.100	-4.5961	-4.4293	-4.1912	-4.1282	-3.9105	-3.8094
0.050	-4.9863	-4.7932	-4.5634	-4.4731	-4.2652	-4.1707
0.010	-5.7703	-5.5672	-5.2794	-5.2029	-5.0229	-4.8374
$n = 100$						
0.100	-4.2889	-4.2300	-4.1070	-4.0616	-3.9456	-3.9092
0.050	-4.6072	-4.5481	-4.4011	-4.3689	-4.2569	-4.2054
0.010	-5.2206	-5.1407	-5.0321	-4.9010	-4.8075	-4.7791
$n = 200$						
0.100	-4.0305	-3.9845	-3.9234	-3.8907	-3.8265	-3.8167
0.050	-4.3371	-4.2920	-4.2239	-4.1990	-4.1363	-4.1112
0.010	-4.9329	-4.8852	-4.8285	-4.8269	-4.7551	-4.7004

Note: Nominal sizes 0.10, 0.05 and 0.01. k is the order of lags in the ADF regression.

Table 3: Bierens (1997) unit root tests

	$\hat{F}(m)$				$\widetilde{T}(m)$			
	$m = 5$	$m = 10$	$m = 15$	$m = 20$	$m = 5$	$m = 10$	$m = 15$	$m = 20$
Algeria	0.74	0.04	0.80	0.98	0.92	0.30	0.94	0.94
Cyprus	0.13	0.61	0.43	0.63	0.40	0.47	0.63	0.53
Egypt	0.80	0.84	0.99	0.99	0.94	0.95	1.00	0.99
Israel	0.03	0.23	0.09	0.55	0.09	0.31	0.65	0.87
Jordan	0.58	0.76	0.88	0.69	0.92	0.91	0.82	0.62
Malta	0.77	0.78	0.84	0.74	0.88	0.92	0.91	0.73
Morocco	0.99	0.99	0.68	0.49	0.99	0.87	0.43	0.20
Syria	0.26	0.31	0.11	0.03	0.16	0.35	0.13	0.12
Tunisia	0.60	0.31	0.26	0.12	0.27	0.42	0.25	0.10
Turkey	0.26	0.18	0.01	0.01	0.15	0.02	0.02	0.09

Note: simulated p-values obtained with EasyReg International by Bierens.

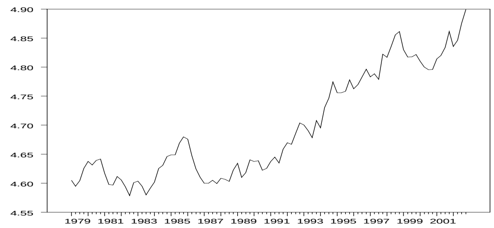
Table 4: ADF test statistics applied to the ESTR models

Country	$\hat{E}$	k	Model
Algeria	-1.6187	0	A
Egypt	-2.9185	4	A
Jordan	-1.7064	3	C
Malta	-4.0187	5	C
Morocco	-3.0117	1	C
Syria	-2.5455	0	C
Tunisia	-2.3816	0	A
Turkey	-4.7652	1	A

Note: $\hat{E}$ is the test statistic for the null hypothesis of unit root of the residuals of the ESTR models. The order k of the ADF regression has been selected by the AIC.

Figure 1: Real Exchange Rates

(a) Algeria

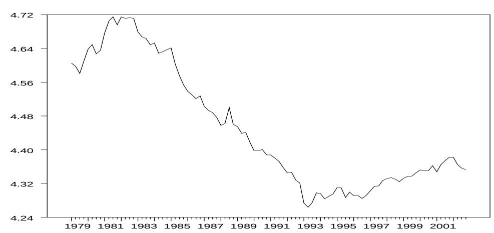


(b) Cyprus

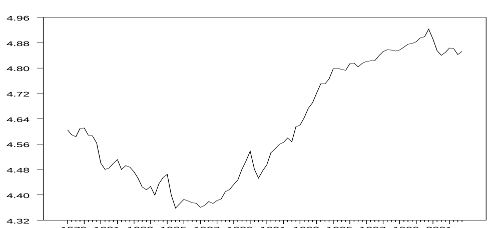
(c) Egypt

(d) Israel

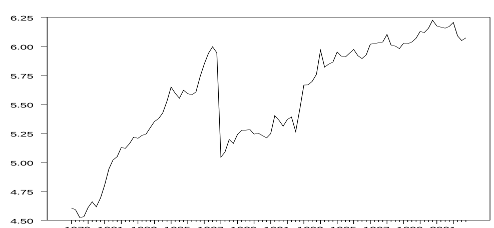
(e) Jordan



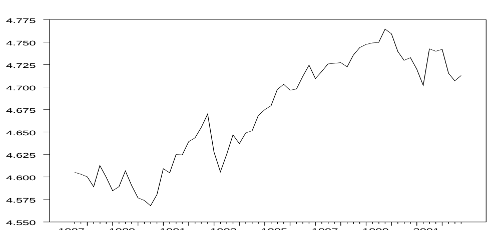
(f) Malta



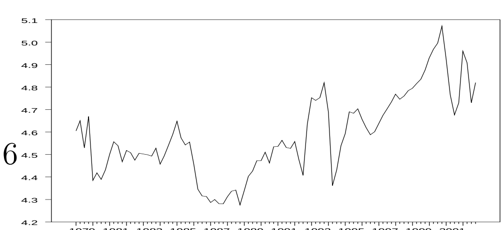
(g) Morocco



(h) Syria



(i) Tunisia



(j) Turkey