A discusión

NONLINEAR TREND STATIONARITY OF REAL EXCHANGE RATES: THE CASE OF THE MEDITERRANEAN COUNTRIES*

Mariam Camarero, Juan Carlos Cuestas and Javier Ordoñez**

WP-AD 2006-27

Correspondence to: Juan Carlos Cuestas, Department of "Fundamentos Análisis Económico", University of Alicante. Campus de San Vicente del Raspeig, Ap. Correos 99, 03080 Alicante, Spain. email: cuestas@merlin.fae.ua.es.

Editor: Instituto Valenciano de Investigaciones Económicas, S.A.
Primera Edición Diciembre 2006
Depósito Legal: V-5352-2006

IVIE working papers offer in advance the results of economic research under way in order to encourage a discussion process before sending them to scientific journals for their final publication.

---

* The authors gratefully acknowledge the financial support from the CICYT and FEDER project SEJ2005-01163, the Bancaja project P1.1B2005-03 and the Generalitat Valenciana Complementary Action ACOMP06/047. Juan Carlos Cuestas also thanks the grant FPU AP2000-2552 by the Spanish Ministerio de Educación y Cultura. The three authors are members of the INTECO research group. The usual disclaimer applies.

** M. Camarero: Universitat Jaume I (Castellón); J.C. Cuestas University of Alicante; J. Ordoñez: Universitat Jaume I (Castellón).

NONLINEAR TREND STATIONARITY OF REAL EXCHANGE RATES:
THE CASE OF THE MEDITERRANEAN COUNTRIES

Mariam Camarero, Juan Carlos Cuestas and Javier Ordoñez

ABSTRACT

The aim of this article is to provide additional evidence on the fulfilment 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 stationarity 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 that the evidence in favour of the Purchasing Power Parity hypothesis increases when we allow for nonlinear alternatives.

J.E.L. Classification: C22, F31

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

1 Introduction

During the last decades, a number of authors have studied whether Purchasing Power Parity (hereafter PPP), a concept introduced by Cassel in 1918 holds. Since then, the empirical validity of PPP 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 their order of integration). 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 less restrictive version 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 on the PPP is very wide. Many different techniques have been used to test for its 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 the deterministic time trend is misspecified. 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 even around a broken time trend (Rappoport and Reichlin, 1989 and Perron, 1989, 1990).

Bearing this considerations in mind some authors have applied unit root tests with structural changes 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.

A broken time trend is a particular case of a nonlinear time trend. Thus, traditional unit root tests, even with structural changes, may incorrectly

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

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 from the market (...)to the exchange rate to return 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, these countries have agreed 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 the PPP can be considered 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 may not hold for some of these countries. In addition, the latter find that the evidence in favour of PPP increases when the unit root tests allow for structural changes.

The aim of this paper is to test whether models allowing for other forms of nonlinear deterministic components are 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

[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.

many of them.

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 our main results.

2 Econometric Methodology

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

The common practice in empirical macroeconomics is to model time series as a unit root rather than trend-stationary processes. However, standard unit root tests are not able to reject the I(1) hypothesis in the presence of breaking deterministic linear trends (Perron, 1989 , 1990). Thus, time series could be, after all, stationary around a (broken) deterministic linear trend. Bierens (1997) has generalized this idea by suggesting the the macro variable may be stationary around a deterministic nonlinear trend. Such trends are meant to capture the evolution of the underlying data generating processes from changes in the economy's structural parameters (Bierens, 2000).

Park and Choi (1988) and Ouliaris, Park and Phillips (1989) first suggest to use ordinary time polynomials in various standard unit root tests, as the Dickey-Fuller test, to capture the presence of breaking deterministic linear trends. Bierens (1997) revised the nonlinear Dickey-Fuller version by replacing the ordinary time polynomials with orthogonal Chebishev time polynomials. The advantage of using the Chebishev polynomials is that they allow to distinguish between stationarity around a linear trend from stationarity around a nonlinear deterministic trend under the alternative hypothesis.

Denote the Chebishev polynomial as $P_{0,t}$ through $P_{m,t}$, where $P_{0,t}$ equals 1, $P_{1,t}$ is equivalent to a linear trend, and $P_{2,t}$ through $P_{m,t}$ are cosine functions. With these polynomials, the augmented Dickey-Fuller test becomes:

$$\Delta z_t = \alpha z_{t-1} + \sum_{j=1}^{p} \phi_j \Delta z_{t-j} + \theta^T P_{t,n}^m + \varepsilon_t \tag{2.1}$$

Bierens (1997) considers the null of unit root with drift against three alternative hypotheses: stationarity around a level, around a linear trend or around a nonlinear trend. This author develops several test statistics for model (2.1): $\hat{t}(m)$ which is the t-statistic on the estimated coefficient $\hat{\alpha}$,

$\hat{A}(m) = \frac{n\hat{\alpha}}{|1-\sum_{i=1}^{p}\hat{\phi}_i|}$, $\hat{F}(m)$ which is F test for the joint hypothesis that $\hat{\alpha}$ and the last m componentes of the parameter vector θ in model (2.1) are zero under the null. When H_0 is rejected, the proper alternative hypothesis will depend on the test statistic involved on and whether there is left-side or right-side rejection (see Table 1). 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:

$$\Delta z_t = -\rho z_{t-1} + \lambda_0 + \rho\lambda_1 t + f(t) + u_t \tag{2.2}$$

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:

$$H_0 : \rho = 0, f(t) \equiv 0, \tag{2.3}$$

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

$$H_1^{lin} : \rho = 1, f(t) \equiv 0, \tag{2.4}$$

whereas the second alternative is nonlinear trend stationarity

$$H_1^{nlin} : \rho = 1. \tag{2.5}$$

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. 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 (as described in Table 1).

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 is zero[3].

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

Table 1: Alternative hypotheses

Test	Left-side rejection	Right-side rejection
$\hat{T}(m)$	LTS	NLTS
$\hat{F}(m)$	-	MS, LTS or NLTS

Note: MS= mean stationarity, LTS= linear trend stationarity, NLTS= nonlinear trend stationarity.

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 as an instantaneous structural break:

$$y_t = \alpha_1 + \alpha_2 S_t(\gamma, \tau) + \nu_t \tag{2.6}$$

$$y_t = \alpha_1 + \beta_1 t + \alpha_2 S_t(\gamma, \tau) + \nu_t \tag{2.7}$$

$$y_t = \alpha_1 + \beta_1 t + \alpha_2 S_t(\gamma, \tau) + \beta_2 t S_t(\gamma, \tau) + \nu_t \tag{2.8}$$

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:

$$S_t(\gamma, \tau) = (1 + exp\{-\gamma[t - \tau T]\})^{-1}, \gamma > 0. \tag{2.9}$$

Model A (equation (2.6)) approximates the nonlinear deterministic component as a transition in the intercept of a non-trending series, Model B (equation (2.7)) does it by a transition in the intercept of a trending time series and, finally, Model C (equation (2.8)) 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 for 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

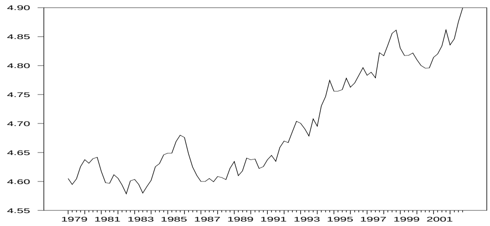
3.1 Data

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 follow, apparently, a linear trend, hence suggesting the possibility of nonlinear deterministic components in the series.

Figure 1: Real Exchange Rates

(a) Algeria



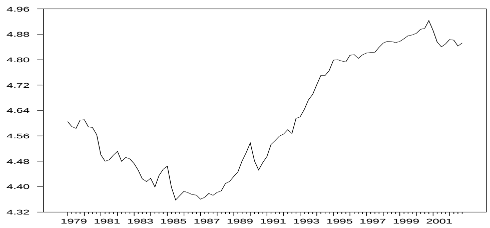
(b) Cyprus

(c) Egypt

(d) Israel

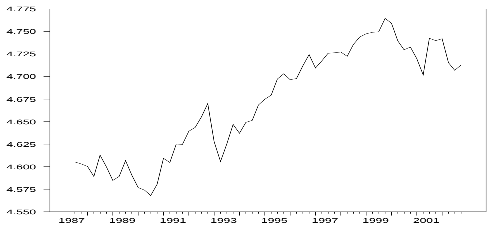
(e) Jordan

(f) Malta



(g) Morocco

(h) Syria



(i) Tunisia

(j) Turkey

3.2 Results for the Mediterranean countries real exchange rates

As a preliminar analysis, first, we display in figure 2 the autocorrelation functions of the RER. The speed of decay of the autocorrelation functions is very slow, implying the presence of a unit root in the series. Additionally, in table 2 we present the results of applying the ADF test to the series of RER. In no case is it possible to reject the unit root hypothesis. This result might be caused by the poor performance of this test when there are nonlinearities in the path of the variable that are not taking into account.

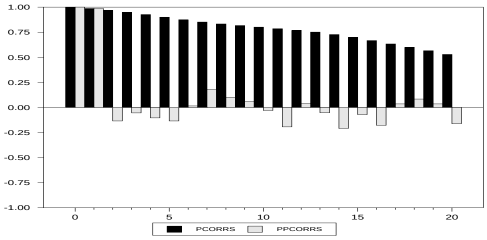
Table 2: ADF test statistics applied to the RER

Country	p-value	k
Algeria	0.661	0
Cyprus	0.692	0
Egypt	0.130	4
Israel	0.060	2
Jordan	0.696	3
Malta	0.978	0
Morocco	0.465	1
Syria	0.336	0
Tunisia	0.564	0
Turkey	0.117	2

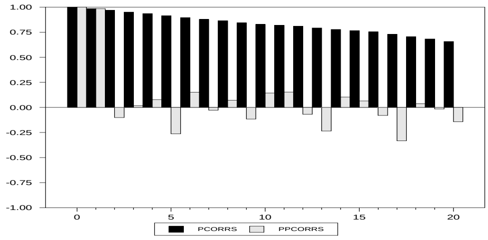
Note: P-values computed by using Mackinnon (1996) critical values. Auxiliary regression for the ADF test with trend and intercept. The order k of the ADF regression has been selected by the AIC.

Now, we analyse the results of the Bierens' (1997) tests. As described above, the $\hat{F}(m)$ test is calculated from the ADF regression where the lag length p has been chosen using the Akaike information criterion (AIC). In addition, we also apply the $\widetilde{T}(m)$ test. 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

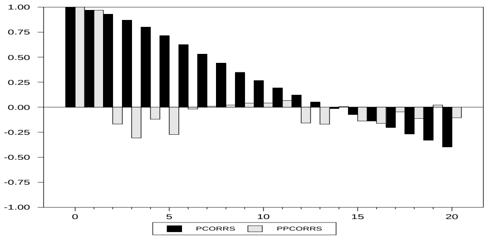
Figure 2: Autocorrelation functions



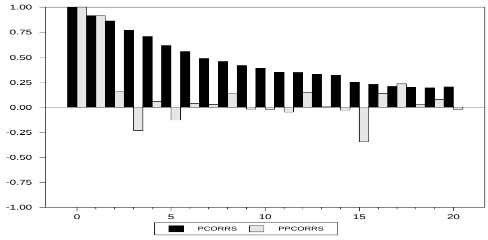
(a) Algeria



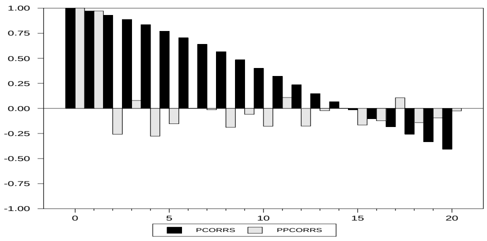
(b) Cyprus



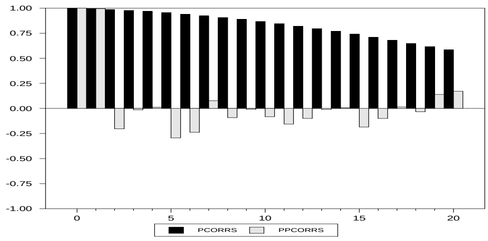
(c) Egypt



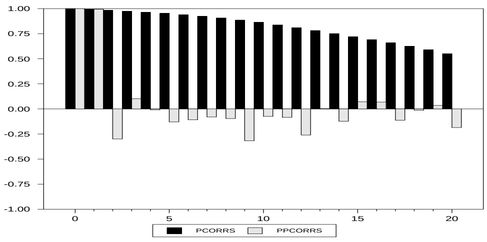
(d) Israel



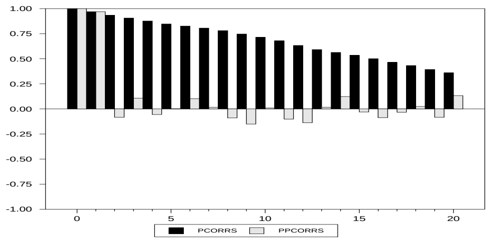
(e) Jordan



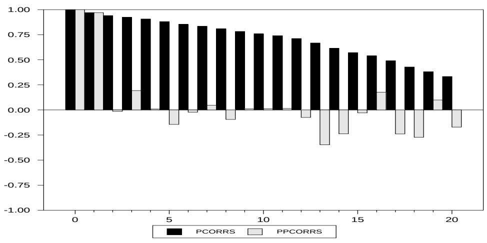
(f) Malta



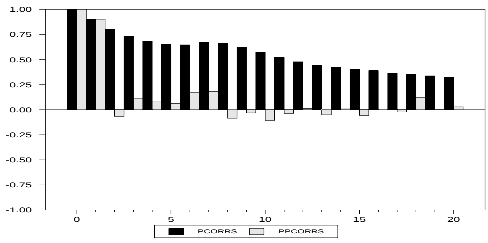
(g) Morocco



(h) Syria



(i) Tunisia



(j) Turkey

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.

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.

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 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 (2.9) 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

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

towards equilibrium is symmetric and does not depend on the sign of the shock. The ESTR function is given by:

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

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.

The critical values for the DF and ADF tests applied to the residuals of the auxiliar nonlinear regressions are presented in tables 5 and 6, and have been obtained by Monte Carlo simulations over 20,000 replications[5]. The null DGP is been specified as:

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

According to the results in table 4, we can reject the null of a unit root for the case of Turkey and Model A (5% level of significance), and at 10% for Malta (model C). Nevertheless, as pointed out by Rodrigues and Rubia (2005), the DF test might suffer from power problems in the presence of ARCH effects in the residuals of the DF regression. In order to check this point, it has been tested the existence of such effects, finding no evidence of ARCH problems in the residuals[6].

These results turn out to be complementary to those found by Camarero, Cuestas and Ordóñez (2006), as PPP was also fulfilled in the cases of Algeria, Egypt and Turkey, whereas in this case additional evidence is found for

[5]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.

[6]Results available from the authors upon request.

Table 5: 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.

Malta when a slow transition is allowed. This joint evidence suggests that the evidence in favour of PPP stationarity improves once the deterministic component is adequately characterized.

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.

Our results complement previous evidence on PPP in the Mediterranean countries. First, using Bierens' unit root tests PPP holds for Algeria, Egypt and Morocco, and thus confirm previous results for Algeria and Egypt. On the other hand, by applying Leybourne, Newbold and Vougas (1998) approach, there is evidence of PPP fulfillment for Malta and Turkey.

Our conclusion is twofold. First, a proper statistical characterization of the deterministic componentes is of crucial importance when testing for

Table 6: 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.

PPP. Second, the use of smooth transition models as a means of representing deterministic structural changes in real exchange rates appears to be appropriate for Malta and Turkey, whereas non-linear alternatives are adequate for Algeria, Egypt and Morocco.

References

Bierens, H. J. (1997): "Testing the unit root with drift hypothesis against nonlinear trend stationarity, with an application to the U.S. price level and interest rate", *Journal of Econometrics*, 81, 29–64.

Camarero, M., J. C. Cuestas and J. Ordóñez (2006): "Purchasing power parity versus the EU in the Mediterranean countries", *Applied Financial Economics*, 16, 157–167.

Cassel, G. (1918): "Abnormal deviations in international exchanges", *Economic Journal*, 28, 413–415.

Dixit, A. (1989): "Hysteresis, import prenetation, and exchange rate pass-through", *Quarterly Journal of Economics*, 104, 205–228.

Doganlar, M. (1999): "Testing long-run validity of purchasing power parity for Asian countries", *Applied Economic Letters*, 6, 147–151.

Dropsy, V. (1996): "Real exchange rates and structural breaks", *Applied Economics*, 28, 209–219.

Frenkel, J. A. (1978): "Purchasing power parity: Doctrinal perspective and evidence from de 1920s", *Journal of International Economics*, 8, 169–191.

Johansen, S. and K. Juselius (1992): "Testing structural hypothesis in a multivariate cointegration analysis of the PPP and the UIP for the UK", *Journal of Econometrics*, 53, 211–244.

Krugman, P. R. (1978): "Purchasing power parity and exchange rates: Another look at the evidence", *Journal of International Economics*, 8, 397–407.

Leybourne, S., P. Newbold and D. Vougas (1998): "Unit roots and smooth transtions", *Journal of Time Series Analysis*, 19, 83–97.

MacKinnon, J. G. (1996): "Numerical distribution functions for unit root and cointegration tests", *Journal of Applied Econometrics*, 11, 601–618.

Michael, P., A. R. Nobay and D. A. Peel (1997): "Transaction costs and nonlinear adjustment in real exchange rates: An empirical investigation", *Journal of Political Economy*, 105, 862–879.

Montañés, A. and J. Clemente (1999): "Real exchange rates and structural breaks: Evidence for the Spanish peseta", *Applied Economic Letters*, 6, 349–352.

Moosa, I. A. (1994): "Testing nonlinearities in purchasing power parity", *Applied Economic Letters*, 1, 41–43.

Obstfeld, M. and M. P. Taylor (1997): "Nonlinear aspects of goods-market arbitrage and adjustment: Hekscher's commodity point revisted", *Journal of the Japanese and International Economics*, 11, 441–479.

Ouliaris, S., J. Y. Park and P. C. B. Phillips (1989): "Asymptotic properties of residual based tests for cointegration", in *Advances in Econometrics and Modelling*, edited by B. Raj, Kluwer, 6–28.

Park, J. and B. Choi (1988): "A new approach to testing for a unit root", CAE Working Paper 88–23, Cornell university.

Parkes, A. L. H. and A. Savvides (1999): "Purchasing power parity in the long run and structural breaks: Evidence from real sterling exchange rates", *Applied Financial Economics*, 9, 117–127.

Perron, P. (1989): "The Great Crash, the Oil Price Shock and the unit root hypothesis", *Econometrica*, 57, 1361–1401.

Perron, P. (1990): "Testing for a unit root in a time series with a changing mean", *Journal of Business and Economic Statistics*, 8, 153–162.

Perron, P. and P. C. B. Phillips (1987): "Does GNP have a unit root? A reevaluation", *Economic Letters*, 23, 139–145.

Rappoport, P. and L. Reichlin (1989): "Segmented trends and non-stationary time series", *Economic Journal*, 99, 168–177.

Rodrigues, P. M. M. and A. Rubia (2005): "The performance of unit root tests under level-dependent heteroskedasticity", *Economic Letters*, 89, 262–268.

Sarno, L. (2000): "Real exchange rate behaviour in the Middle East: A re-examination", *Empirical Economics*, 66, 127–136.

Sollis, R. (2005): "Evidence on purchasing power parity from univariate models: the case of smooth transition trend-stationarity", *Journal of Applied Econometrics*, 20, 79–98.

Taylor, M. P. (1988): "An empirical examination of long-run purchasing power parity using cointegration techniques", *Applied Economics*, 20, 1369–1381.

Taylor, M. P. (1992): "Dollar-Sterling exchange rate in the 1920s: Purchasing power parity and the Norman Conquest of 4.68$", *Applied Economics*, 24, 803–811.

Taylor, M. P. and D. A. Peel (2000): "Nonlinear adjustment, long-run equilibrium and exchange rate fundamentals", *Journal of International Money and Finance*, 19, 33–53.

West, K. D. (1988): "Asymptotic normality when regressors have a unit root", *Econometrica*, 56, 1397–1418.