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REAL EXCHANGE RATES OVER A CENTURY: THE CASE OF THE DRACHMA/STERLING RATE, 1833-1939

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ABSTRACT

Recent studies on real exchange rates advocate the use of long samples in order to reveal the low frequency properties of the processes. The present paper contributes to this strand of the literature by exploiting recently released time series for the drachma/sterling rate for the period 1833-1939. This is an interesting period as it covers different exchange rate regimes and the effects of important historical events. In the paper, the mean-reverting behaviour of the real drachma/sterling exchange rate is initially examined applying univariate unit root tests and then the validity of Purchasing Power Parity (PPP) is tested using cointegration analysis. The results provide support for a weak PPP relationship, which turns out to be robust across different sub-periods characterised by different exchange rate regimes. Adjustment to PPP is reached at a relatively high speed and occurs *via* movements of the nominal exchange rate.

JEL classification: F31; C32; N23; N24.
Keywords: real exchange rates; cointegration; PPP.

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1. Introduction

The behaviour of real exchange rates has been one of the most extensively investigated topics in the financial economics literature during the last three decades or so. The question of interest is whether real exchange rates tend to revert to a constant mean, or, equivalently, whether the purchasing power parity (PPP) doctrine holds as an equilibrium relationship. Early empirical studies on this topic can be categorised in three groups, based on the different types of the tests they apply[1]: The "correlation" studies, which apply traditional regression analysis, the unit root studies, which test for stationarity of the real exchange rates, and the cointegration studies, which test for the existence of a cointegrating relationship connecting nominal exchange rates with foreign and domestic prices. Most correlation studies fail to support PPP, whereas the early unit root and cointegration studies provide mixed evidence.

Failure to find evidence of PPP in the post-1973 flexible exchange rate period is attributed by recent studies to the short sample size of the observations and the low statistical power of the early tests.[2] To address the problem of the short sample size, a number of researchers advocate the use of long historical time series (e.g. Lothian and Taylor, 1996) whereas others use the information of time series data from a large number of countries performing panel econometric analysis (e.g. Papell, 1997). The problem of the low statistical power of the initial tests can be dealt with by applying more advanced econometric techniques. Such techniques include high power unit root tests (e.g. Ng and Perron, 2001), unit root tests which may also account for possible structural breaks (e.g. Cheung and Lai, 1998), panel unit root and panel cointegration tests (e.g. Pedroni, 2001), tests which account for possible non-linearities in the behaviour of the real exchange rates (e.g. Taylor *et al.*, 2001) and multivariate cointegration techniques – mainly the Johansen methodology (e.g. Juselius, 1995).

Studies which use large samples of historical time series sustain that the real exchange rates may revert to PPP over very long periods of time, possibly of a century or more (see *inter alia* Lothian and Taylor, 1996; Taylor, 2002). Following this reasoning, even high frequency data over relatively short periods, such as for twenty or thirty years, may not disclose mean-reversion to PPP, while data over a century or more might do so. Advocates of this strategy state that in order to reveal the long-run or low frequency properties of the real exchange rate processes, we need long horizon data. By now, long span studies constitute a

[1] In this, we follow the categorisation of Froot and Rogoff (1995).
[2] See Taylor and Taylor (2004), Taylor (2006) for recent reviews of the debate and evidence on PPP.

large body of the international finance literature and most of them provide evidence of mean reverting behaviour for the real exchange rates.

In the present paper, we extend this strand of the literature by examining the properties of the drachma/sterling real exchange rate using recently released data (Dertilis, 2005) on the drachma/sterling rate and Greek prices for the period 1833-1939. This is an interesting period as it covers different exchange rate regimes and the effects of a number of important historical events. In addition, even though there exist numerous studies testing for PPP using Greek data for the post-Bretton Woods period, (see *inter alia*, Sideris, 2000; Brissimis *et al.*, 2005) no previous work covers the period under study. Georgoutsos and Kouretas (1992) and Phylaktis (1992) test for PPP between Greece and the US for a short sub-period of the present sample, the period 1923-1925, employing monthly observations. Both studies provide some support for mean reverting real exchange rates.

In the study, we follow the methodological suggestions of the relevant literature. We initially test for stationarity of the real drachma/sterling rate using a battery of univariate unit root tests, which include the high power Dickey-Fuller generalised least squares tests developed by Elliott *et al.* (1996) and the modified geneneralised least square versions of the Phillips-Perron tests developed by Ng and Perron (2001). Then we test for cointegration among the nominal exchange rate and Greek and UK prices and assess the validity of alternative versions of PPP, using the Johansen (1995) methodology.

We also aim to overcome problems associated with long span studies. Long span studies have been criticized (a) in that they mix data from both fixed and flexible exchange rate regimes, (b) to contain serious structural breaks (see e.g. Roggof, 1996) and (c) not to be able to answer the question of whether PPP holds for floating exchange rate regimes (see e.g. Amara and Papell, 2006).[3] Following the suggestions of relevant studies (see e.g. Grilli and Kaminski, 1991; Lothian and Taylor, 1996; Taylor, 2002) we evaluate the robustness of the empirical results by examining the behaviour of the real exchange rate in different sub-periods covering alternative exchange rate regimes. We pay particular attention to the sub-period 1877-1927, which mainly covers a period of floating exchange rates. In the modeling framework which examines the validity of PPP as a long-run relationship, we also account for the effects of structural breaks.

[3] Point (a) is based on the observation that the volatility of the real exchange rates typically shifts according to the exchange rate regime, and this may imply that the exchange rate series have different statistical properties in the different regimes.

The rest of the paper is organised as follows: Section 2 provides a brief literature review of PPP studies which use long horizon data. Section 3 offers some historical information on the economic circumstances which prevailed in Greece during the period under study. Section 4 describes briefly the theory and the specifications of the PPP doctrine and outlines the econometric tests and techniques we apply in the empirical work. Section 5 presents the empirical work and results. The final section summarises and concludes.

2. Long-span studies testing for PPP

Studies using long horizon observations in order to analyse the long-run properties of the real exchange rates represent a large body of the literature on PPP. A number of studies published in the late 1980s and in the 1990s use historical data from the nineteenth and twentieth centuries covering a hundred or more of annual observations mainly for industrial economies. Their main argument is that the short time span is one of the possible reasons which account for the empirical failure of PPP during the floating period (see e.g. Lothian and Taylor, 1996). This idea is further supported by the works of Campbell and Perron (1991) and Hakkio and Rush (1991) who show that in analysing the long-run characteristics of interrelated time series, the length of the series is more important than the frequency of the observations.[4]

Early studies in this strand of the literature examine the mean reverting properties of real exchange rates and find evidence of significant mean reversion: Frankel (1986) uses dollar/sterling real exchange rate data for the period 1869-1984; he rejects the random walk hypothesis in favour of a first order autoregressive process with an autocorrelation coefficient of 0.86. Similar results to Frankel are obtained by Edison (1987), in the framework of an error correction model for the period 1890-1978. Abuaf and Jorion (1990) study a century of dollar-franc-sterling exchange rate data and verify PPP. Lothian (1991) also finds evidence that the yen/sterling and yen/dollar real exchange rates are stationary during the years 1875-1989. Diebold, *et al.*(1991) explore long time spans (ranging from 74 to 123 years) of nineteenth century data covering the gold standard pre-1913 period for six countries –five European countries and the US. They capture fractional integration processes and they find support for PPP as a long-run concept. Grilli and Kaminski (1991) provide evidence for stationarity of the sterling/dollar real exchange rate for the period 1885-1986, but not over

[4] Thus, due to the choice of data at annual frequency and the longer spans, the statistical tests performed in these studies are considered to have reasonably high power to make robust inferences.

specific sub-periods. Glen (1992) reports mean reversion for the real exchange rates of nine countries over the period 1900-1987. Lothian and Taylor (1996) use observations of two hundred years covering the nineteenth and twentieth centuries on dollar/sterling and franc/sterling real exchange rates; they apply unit root tests and reject non-stationarity for the real exchange rates.

More recently, Alves, Cati and Fava (2001) apply fractional cointegration techniques to test for PPP using data from Brazil for the period 1855-1966. They provide some evidence for the relative version of PPP but not for its absolute version. Costa and Crato (2001) use century-long annual time series to examine the behaviour of the real exchange rates between Portugal and the UK, and Portugal and the US; they report evidence for mean reverting real exchange rates. Taylor (2002) uses samples of more than 100 observations starting in the late nineteenth century for twenty real exchange rates with respect to the US dollar. He provides positive evidence for long-run PPP not only for industrialized economies but also for some developing economies. Sabaté *et al.* (2003, 2005) use observations for peseta exchange rates (*vis-à-vis* the UK sterling, the US dollar and the French franc) for the period 1870-1935. They employ a number of unit root tests and reject non-stationarity of the real exchange rates, once structural breaks are accounted for in the empirical analysis. Hasan (2006) uses a data set of more than a hundred observations for Australia, the UK and Canada. He employs cointegration analysis and a number of univariate tests to test for stationarity of the real exchange rates and for the validity of long-run PPP. Overall, his findings support mean reverting behaviour of the real exchange rates and the existence of PPP-type cointegrating relationships.

3. The drachma during the period 1833-1939

The present sample, which covers the years 1833-1939, offers a rich body of data for studying the behaviour of the drachma exchange rate. Over the period under study, exchange rate arrangements varied considerably, ranging from the pure gold standard, to wartime controls, to episodes and periods of floating exchange rates. The various exchange rate regimes are summarised in Table 1.

The sample starts in 1833, when a new currency, the drachma, was introduced. The drachma was issued on a bimetallic standard that valued silver to gold at 15.5:1 and was legally under this bimetallic regime until 1876 with the exception of a couple of episodes of non-convertibility. Over the period 1877-1909, the drachma pursued a flexible exchange rate regime despite numerous efforts of the monetary authorities to revert to a fixed exchange rate regime. Convertibility was restored in 1910, when the drachma joined the gold standard and adopted fixed exchange rate with the French franc. It remained under the gold standard even after the beginning of the World War I, until 1918. Between 1915 and 1918, Greece experienced high wartime inflationary pressures and a depreciation of the drachma which eventually led to the gold standard being abandoned in 1919. Between 1919 and 1927 the drachma was again in a flexible regime; the period is characterised by the destructive war in the Asia Minor, 1921-1923, which resulted in severe economic crisis. Convertibility of the drachma was restored in 1928 and Greece joined the interwar gold exchange standard until 1939 which is the last observation of the present sample.[5]

The period covers five nominal exchange rate regimes and this can be considered as the inevitable cost of increasing the length of the sample size. However, at the same time, the long span provides the basis for more strict examination of the behaviour of the real drachma/sterling exchange rate.[6] The present work aims to overcome problems associated with studies using long horizon data. To this end, analysis is also performed in sub-periods. Considering data sets that cover at least fifty observations, in order to avoid loss of statistical power, empirical work is performed in the following four sub-periods: (i) 1833-1918, (ii) 1833-1927, (iii) 1877-1927 and (iv) 1877-1939. Among the above four sub-periods, the 1877-1927 period is of particular interest, as drachma was under a flexible exchange rate regime during most of the years covered. It can then be considered as a period of a flexible exchange rate regime, with the exception of the 9-year gold standard phase.

4. Theoretical and methodological issues

4.1 Theoretical issues

PPP states that the nominal exchange rate between the currencies of two economies should equal the ratio of the two relevant national aggregate price levels. The basic general specification of the hypothesis in its stochastic log-linear form is:

[5] We do not extend the sample to more recent observations, given that most statistical series in Greece suffer from a serious break during the World War II when effectively no data is available. The period 1941-44 was characterised by hyperinflation and a strong and continuous devaluation of the drachma. In the period April 1941 – October 1944, the price of gold sovereigns increased 523 million times whereas the cost of living in Athens was multiplied by 1.4 billion. In 1944, the fiscal and monetary systems of Greece broke down. In November 1944 a new drachma was issued, worth 50,000 million old drachmae (see, *inter alia*, Alogoskoufis and Lazaretou, 2002). Price indices are also defined and measured differently in the pre- and post- War II periods (see Bank of Greece *Statistical Bulletin*, various issues).

[6] In this, we follow the suggestions of Lothian and Taylor, 1996.

$s_t = \gamma_0 + \gamma_1 p_t + \gamma_2 p_t^* + u_t$ (1)

where p_t, p_t^* indicate the logs of the price levels of the domestic and the foreign economy respectively, s_t the log of the exchange rate denominated in the currency of the domestic economy and u_t is the error term.

Strong PPP is implied by the proportionality restriction (γ_1=1, γ_2 =-1):

$s_t = \gamma_0 + p_t - p_t^* + u_t$ (2)

and states that, whatever the monetary or real disturbances in an economy, under the assumption of instantaneous costless arbitrage, the prices of a common basket of goods in the two countries measured in a common currency will be the same.

However, strong PPP cannot be expected to hold always as an empirical proposition. The prices of a given commodity will not necessarily be equal in different locations, because of transportation costs, possible tariff barriers, information costs and measurement error problems. Thus, the relationship is more likely to have the weak PPP form implied by the restriction of symmetry (γ_1=- γ_2):

$s_t = \gamma_0 + \gamma_1(p_t - p_t^*) + u_t$ (3)

with γ_1 being a constant factor which accounts for assumed constant transportation, information costs and measurement errors. γ_1 is allowed to differ from unity, implying that long-run proportionality between the exchange rate and relative prices may not be exactly one-to-one.

The real exchange rate is defined as the nominal rate deflated by a ratio of foreign and domestic price levels. In logs, the real exchange rate q_t is given by:

$q_t = s_t - p_t + p_t^*$ (4)

Thus, from a statistical point of view, the validity of strong PPP as given by equation (2) reduces to a stationarity test of q_t.

Based on the alternative specifications of PPP, existing studies examine the validity of the doctrine by testing: (i) for stationarity of the real exchange rate, (ii) for evidence of cointegration between the exchange rate and relative prices as implied by (3) and (iii) for the existence of a cointegrating relationship involving the exchange rate, domestic and foreign prices of the form of (1). In the present study, we first test for stationarity of q_t using a battery of unit root tests. We then test for evidence of a cointegrating relationship of the form of (1) and once cointegration is established, we evaluate the validity of strong and weak PPP as described in (2) and (3), respectively.

4.2 Econometric issues

Univariate unit root tests

Initially, the traditional Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests are performed to identify the order of integration of the real drachma/sterling rate series. The selection of the lag truncation (k) for the ADF and PP tests is based on the traditional Akaike (AIC) and Scwartz (SIC) information criteria. In order to avoid size problems, we follow the recommendations of Ng and Perron (2001) and employ ADF and PP tests with a lag truncation based on a modified AIC (MAIC).[7]

However, the DF tests have been criticised because of their limited power. The power of these tests is limited where the root is very close to the unit circle and decreases as deterministic factors are added. More powerful univariate unit root tests are the generalised least squares versions of the DF tests (DF-GLS) due to Elliott *et al*. (1996). The DF-GLS tests seem to be appropriate for the analysis of real exchange rates, since they allow for possible deterministic trends, which could be present in real exchange rate behaviour caused by e.g. Balassa-Samuelson effects.[8] In the DF-GLS tests, the series z_t to be tested is replaced in the ADF regression by $z'_t = z_t - \hat{a}' d_t$,where â' is a GLS estimate of the coefficients on the deterministic trends d_t. In the present work, we perform the DF-GLS tests allowing the series to be (i) demeaned and (ii) demeaned and de-trended and use the AIC, SIC and MAIC criteria to select the lag k. We also calculate the Elliot Rothenberg and Stock (ERS) point optimal test (Elliott *et al.*, 1996).

The PP tests have been criticised to give rise to size distortions in the presence of a significant negative moving average factor. Ng and Perron (2001) have modified the PP tests and constructed four test statistics with good size and power based upon the GLS de-trended data as proposed by Elliot *et al.* (1996) to be used in conjunction with a suitably chosen k. The Ng and Perron (NP) tests (which are also known as M-tests) are the following: MZ_α and MZ_t are the modified versions of Z_α and Z_t tests of the PP tests; the MSB test is a modified

[7] Given that size problems can arise from a near common factor in the moving-average and autoregressive polynomials in the time series ARMA representation, Ng and Perron (2001) emphasise that the selection of the lag truncation (k) plays a crucial role in the size of the unit root tests. Traditional information criteria such as the AIC tend to select a lag truncation that is too low. When, in particular, there are errors with a moving average root close to -1, a higher order augmented autoregression would be necessary to avoid over-rejecting the null hypothesis of a unit root. In order to account for this type of problems they suggest the use of a modified AIC (MAIC) with a penalty factor that is sample dependent.

[8] The DF-GLS tests have been considered ideal for PPP testing (see *inter alia* Taylor, 2002; Cheung and Lai, 1998).

version of the Bhargava test R_t and the MP_T test is the modified version of the Elliot *et al.* (1996) point optimal test (for details, see Ng and Perron, 2001).

The Johansen cointegration technique

To test for cointegration we apply the Johansen (1995) cointegration technique, for a vector x_t such that: $x_t' = (s_t, p_t, p_t^*)$. The technique leads to the estimation of the matrix Π of the long-run responses of the vector and shows that in the event that there exist r cointegrating relationships, Π can be expressed as the product of two matrices α and β' (Π = α β'), where β contains the r cointegrating vectors and α is the loadings matrix. Johansen provides the test statistics to define r and to test for linear restrictions on the parameters of either α or β'. In the present case and in the event that there is evidence for one cointegrating vector, the restriction $\beta = (1, -\beta_1, \beta_1)$ implies the symmetry hypothesis H_1 and provides evidence for the validity of weak PPP – or in other words that the real exchange rate follows a pattern based on market fundamentals – whereas the restriction $\beta = (1, -1, 1)$ implies the proportionality hypothesis H_2 and provides evidence for the validity of strong PPP – or stationarity of the real exchange rate. The Johansen technique also allows for the effects of structural breaks and for possible interactions in the determination of the variables so that no variable has to be considered *a priori* exogenous.

5. Empirical analysis and results

5.1 Univariate Analysis

The analysis employs annual data on the drachma/sterling exchange rate s_t, the cost of living index in Greece p_t, and the cost of living index in the UK p_t^*; the real exchange rate q_t is defined as in (4). The Greek series (the exchange rate and the price index) are taken from Dertilis (2005) whereas the UK price index is taken from Mitchell (1992). All variables are measured as natural logarithms.

First, univariate analysis of the real exchange rate q_t is performed applying the ADF, PP, DF-GLS, ERS point optimal and NP tests. Tables 2 and 3 report the test results for the whole period and the floating exchange rate period 1877-1927, respectively.[9] The truncation lag is selected following the SIC, AIC and MAIC criteria. The results based on the PP tests which allow an intercept and a trend to be included in the autoregressive spectral density

[9] Similar results are obtained by the univariate analysis for the periods (i), (ii) and (iv), but they are not reported for reasons of space.

estimation, are not reported given that the trend did not turn out to be significant at the conventional significance levels.

At first sight, the results seem to vary depending on the test procedure, the truncation lag selected and the estimated period. However, all tests whose lag truncations are based on the MAIC criterion provide evidence for a unit root in the real exchange rate, independently of test procedure and estimation period. Given that the results of the tests based on the MAIC criterion are shown by Ng and Perron (2001) to have higher statistical power, we can maintain that there is no evidence for mean reverting behaviour of the drachma/sterling real exchange rate.

5.2 Cointegration analysis

Full sample analysis

In a second step, cointegration between s_t, p_t* and p_t is investigated using the Johansen approach. Initially, a three-dimensional vector autoregressive (VAR) system is estimated using multivariate least squares. The estimation involves three lags of the variables to obtain non-correlated residuals; hence, the effective estimation period is reduced so as to accommodate the lag structure of the model. In the VAR, the deterministic variables set includes a constant, a trend and a number of dummy variables. The constant and the trend are kept in the system as they turn out to be significant at a 1% level of significance (F-test on retained constant: $F(3, 86)=5.872$; F-test on retained trend: $F(3, 86)=9.037$). The dummies are included to account for specific structural breaks that affected the performance of the Greek economy over the period under study.[10] They are reported in the second column of Table 4.[11] All reported dummies are kept in the system, as they turn out to be significant, whereas their inclusion ameliorates significantly the normality properties of the residuals. Thus specified, the VAR satisfies the statistical assumptions required for the Johansen technique and thus we can proceed with the cointegration analysis.[12] The outcomes of the maximum eigenvalue and trace statistics are reported in the first raw of Table 4. According to both likelihood ratio tests, there is strong evidence for one cointegrating vector in the model.

[10] Allowance for structural breaks is crucial for the correct specification of PPP testing as evidenced by *inter alia,* Sabaté et al. (2003, 2005).

[11] The impulse dummy D48 (taking the value 1 in 1848) accounts for the pause of drachma's convertibility in 1848. The step dummy S1518 (taking the value 1 in 1915-1918) accounts for the effects of the World War I years 1915-1918. The step dummy S2123 (taking the value 1 in 1921-23) accounts for the effects of the 1921-1923 war in Asia Minor; the exchange rate market was characterised by war time controls which resulted in a monetary crisis in 1923. The impulse dummy D25 accounts for a political crisis in Greece in 1925, which caused large fluctuations in the drachma/sterling rate. In 1931 sterling is devalued and the UK leaves the gold standard; the drachma remains in the gold standard, linked to the US dollar: D31 accounts for these effects.

[12] The diagnostic tests do not indicate any serious mis-specification problem. They are not reported here for reasons of space but are available on request.

The estimated coefficients of the cointegrating vector, which are reported in the first raw of Table 5 indicate that it could express a PPP relationship, given that the estimated coefficients take the theoretically expected signs and magnitudes. Nevertheless, formal testing is also required. Table 5 presents the outcomes of the likelihood ratio test statistics for the hypotheses of symmetry (H_1) and proportionality (H_2) concerning the specification of the cointegrating vector. H_1 and H_2 result in test statistics that are asymptotically χ^2 distributed with one and two degrees of freedom, respectively. According to the test outcomes, H_1 is accepted by the data set, whereas H_2 is not. The results provide evidence for the validity of weak PPP in the long run, of the form (standard error in parenthesis):

$$s_t = 1.3\ (0.03)(p_t - p_t^*) \qquad (5)$$

The PPP relationship in different periods

To test further the robustness of the results – in particular, the stability of the cointegrating relationship of the form of (5), which is estimated for the whole period – we test for cointegration in the four sub-periods (i)-(iv) as defined in section 3. The results on the cointegration rank for the four periods are reported in lines 2-5 of Table 4. According to the Johansen test statistics there is evidence for one cointegrating vector for all four sub-samples. In addition, the estimated parameters of the cointegrating relationships – reported in Table 5 – obtain signs and magnitudes which could describe PPP relationships. The estimated coefficients of the four cointegrating vectors are also close in magnitude to those of the full sample. Table 5 reports the results of the formal testing for the validity of hypotheses H_1 and H_2. The results indicate that weak PPP is not rejected for the three out of the four periods (ii) – (iv) at the conventional 5% level of significance, and for period (i) at the 1% level of significance. The empirical evidence thus indicates that weak PPP can be considered as a robust relationship approximating the behaviour of the real drachma/sterling rate for the period under study. Stationarity of the real drachma/sterling rate –or strong PPP- is rejected even at a 1% significance level, for all sub-samples, result, which is consistent with the outcomes of the univariate analysis.

5.3 The dynamic adjustment to PPP

Assuming that (5) represents a reliable specification for the cointegrating vector of the analysis of the full sample, we can go on and perform weak exogeneity tests. These tests are essentially tests for the significance of the cointegrating vector, when used as error correction term in the equations which model the short-run dynamics of the variables. If, for example, the weak PPP between Greece and the UK of the form of (5) enters significantly the equation modelling the short-run dynamics of the nominal drachma/sterling exchange rate s_t, s_t cannot be considered as weakly exogenous with respect to the parameters of the long-run relationship. Weak exogeneity tests are reported in Table 6. The results lead us to comfortably accept the hypothesis of weak exogeneity of foreign and domestic prices, while this hypothesis is rejected for the nominal exchange rate. The results indicate that the nominal exchange rate adjusts in the short run, in a way to restore the equilibrium PPP relationship implied by (5). The results are consistent with the view that nominal exchange rates adjust quickly, even as prices move sluggishly, an assumption common to many international macroeconomic models (Dornbush, 1976; Obstfeld and Rogoff, 1996).

Having indicated that there exists a robust PPP relationship, and that in the short run the nominal exchange rate adjusts to restore equilibrium, we further estimate an error correction model for the drachma/sterling rate, of the form:

$$\Delta s_t = \zeta(s_t - \gamma_1(p - p^*))_{t-1} + \sum_{t=1}^{l} \lambda_{1i}\Delta s_{t-i} + \sum_{t=1}^{l} \lambda_{2i}\Delta p_{t-i} + \sum_{t=1}^{l} \lambda_{3i}\Delta p^*_{t-i} \qquad (6)$$

where l denotes the number of lags involved in the estimation and $\zeta < 0$. The estimation is performed with l=3 to obtain non-correlated residuals. ζ takes the value -0.38 (0.09).[13] The estimate indicates that shocks that take the nominal exchange rate away from equilibrium are corrected at the relatively fast rate of some 38 percent per year. This speed of adjustment implies in turn that the half life of mean reversion is about 1.5 years.

6. Conclusions

A large body of literature on real exchange rate behaviour attributes the empirical rejection of the PPP doctrine to short time spans coupled with the low power of conventional unit root tests and advocate the use of long span data for PPP testing. The present paper makes a contribution to this strand of the literature as it investigates empirically the relevance of PPP between Greece and the UK using observations covering more than a century. The present study is the first one to analyse the properties of a drachma exchange rate using historical drachma series going back to the early 19th century and spanning an interesting period which covers different exchange rate regimes and the effects of a number of important historical events. Methodologically, the study employs a battery of recently developed unit

[13] We do not report the estimated values of the other coefficients in (6), for space reasons.

root tests and a powerful cointegration technique. Given the longer time span, the choice of data at annual frequency and the high power of the tests implemented in the study, the results can be considered as statistically robust. In addition, in order to cope with criticism on long span studies and to strengthen the validity of the results, analysis is also performed in sub-periods, one of which can be considered as a flexible exchange rate regime period.

The empirical work based both on the univariate testing and on cointegration fails to support stationarity of the real exchange rate or the validity of strong PPP. However, the Johansen method indicates that the nominal drachma/sterling rate and Greek and UK prices form a valid cointegrating vector and support the weak form of PPP. The estimated weak PPP relationship turns out to be robust across periods characterised by different exchange rate regimes and to be valid during the flexible exchange rate regime period 1877-1927. Consistent with the sticky price hypothesis, weak exogeneity tests indicate that adjustment to equilibrium comes *via* movements in the nominal exchange rate. In addition, modelling of the short-run dynamics of the nominal drachma/sterling exchange rate in an error correction framework shows that adjustment to equilibrium PPP is reached at the relatively high speed of 38% per annum, which implies a half life of parity reversion of 1.5 years. In general, the results support that, in equilibrium, the drachma/sterling real exchange rate reverts to a PPP-based level, evidence which is in line with studies using long span data for other currencies.

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Table 1: Drachma exchange rate regimes during the period 1833-1939

1833-1876	Fixed bimetallic regime
1877-1909	Flexible regime
1910-1918	Gold standard
1919-1927	Flexible regime
1928-1939	Gold standard

Table 2: Unit root tests for q: 1833-1939

	Intercept		Intercept and trend	
	k lags		k lags	
ADF	1_{SIC}	-2.763	$1_{SIC,AIC}$	-3.436
	3_{AIC}	-2.484		
	5_{MAIC}	-1.991	5_{MAIC}	-1.557
PP	1_{SIC}	-2.907*		
	3_{AIC}	-2.279		
	5_{MAIC}	-1.898		
DF-GLS	1_{SIC}	-2.067*	$1_{SIC,AIC}$	-3.481*
	3_{AIC}	-1.536		
	5_{MAIC}	-1.125	5_{MAIC}	-1.804
ERS	1_{SIC}	3.364	$1_{SIC,AIC}$	4.028
	3_{AIC}	6.072		
	5_{MAIC}	10.121	5_{MAIC}	12.096

NP tests

Intercept					Intercept and trend				
k	MZ_α	MZ_t	MS_B	MP_T	k	MZ_α	MZ_t	MS_B	MP_T
1_{SIC}	-8.83*	-2.1*	0.23	2.80	$1_{SIC,\ AIC}$	-25.32**	-3.47**	0.13**	4.06*
3_{AIC}	-4.86	-1.54	0.31	5.06					
5_{MAIC}	-2.89	-1.18	0.41	8.43	5_{MAIC}	-7.66	-1.82	0.23	12.21

Note: *and ** indicate rejection of the null hypothesis at the 5% and 1% level of significance, respectively.

Table 3: Unit root tests for q: 1877-1927

	Intercept		Intercept and trend	
	k lags		k lags	
ADF	1_{SIC}	-3.575**	1_{SIC}	-3.817*
	4_{AIC}	-2.676	4_{AIC}	-3.243
	8_{MAIC}	-1.224	3_{MAIC}	-2.202
PP	$1_{SIC, AIC}$	-3.684		
	8_{MAIC}	-1.365		
DF-GLS	1_{SIC}	-3.213**	1_{SIC}	-3.869**
	5_{AIC}	-1.767	4_{AIC}	-3.076
	8_{MAIC}	-0.323	3_{MAIC}	-2.233
ERS	1_{SIC}	1.093**	1_{SIC}	2.872**
	4_{AIC}	0.679**	4_{AIC}	0.001**
	8_{MAIC}	19.89	3_{MAIC}	7.862

NP tests

Intercept					Intercept and trend				
k	MZ_α	MZ_t	MS_B	MP_T	k	MZ_α	MZ_t	MS_B	MP_T
1_{SIC}	-23.0**	-3.39**	0.14**	1.06**	1_{SIC}	-31.72**	3.96**	0.124*	2.98**
4_{AIC}	-37.1**	-4.30**	0.11**	0.66**	4_{AIC}	-730**	-191**	0.002**	0.001**
8_{MAIC}	-1.26	-0.794	0.627	19.33	3_{MAIC}	-11.39	-2.35	0.206	8.169

Note: *and ** indicate rejection of the null hypothesis at the 5% and 1% level of significance, respectively.

Table 4: The Johansen procedure results: Testing for the cointegration rank

Period	**Dummies**	**Maximal Eigenvalue**			**Trace statistic**		
		r = 0	r = 1	r =2	r = 0	r = 1	r =2
Full sample	D48, D25, D31, S1518, S2123	48.64**	10.01	0.004	58.65**	10.02	0.004
1836-1918	D48, S1518	24.37*	5.65	2.152	32.18	7.807	2.154
1836-1927	D48, D25, S1518, S2123	34.98**	5.443	0.0004	40.42*	5.444	0.0004
1877-1927	D25, S1518, S2123	31.77**	2.073	0.912	34.75*	2.459	0.912
1877-1939	D25, D31, S1518, S2123	35.08**	7.229	1.85	44.15**	9.08	1.85
	Critical values at 95%	23.8	16.9	3.7	34.6	18.2	3.7

Note: * and ** indicate rejection of the null hypothesis at the 5% and 1% level of significance, respectively.

Table 5: The Johansen procedure results: Restriction Testing

Sample periods	**Estimated β vectors**			**H_1** ($\beta_1 = -\beta_2$)	**H_2** ($\beta_1=-1, \beta_2=1$)
	s	p	p*	χ^2 *(1)*	χ^2 *(2)*
Full sample	1	1.152	-1.273	0.60084	41.502 **
1836- 1918	1	2.151	-1.089	6.1011*	11.151**
1836-1927	1	1.245	-1.271	0.03101	13.584 **
1877-1927	1	1.282	-1.239	0.0677	10.431**
1877-1939	1	1.205	-1.234	0.02651	20.842 **

Note: * and ** indicate rejection of the null hypothesis at the 5% and 1% level of significance, respectively.

Table 6: Weak exogeneity tests

Full sample 1836-1939	χ^2 *(2)*
$H_1 \cap$ w. exogeneity of p	2.090
$H_1 \cap$ w. exogeneity of p*	4.139
$H_1 \cap$ w. exogeneity of s	33.931**

Note: * and ** indicate rejection of the null hypothesis at the 5% and 1% level of significance, respectively.

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