Purchasing Power Parity Revisited
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ABSTRACT

This paper presents a selective survey on some recent empirical attempts to test the validity of Purchasing Power Parity (PPP) to explain exchange-rate movements in the main currencies, as well as a review of the econometric methodology used in such tests. Finally, we offer some encouraging results regarding the forecastability of exchange rate using PPP.

JEL Codes: F31, C32, C33

Key words: Exchange rates, Purchasing Power Parity, Cointegration

1. INTRODUCTION

*“I go to seek a great perhaps”
François Rabelais*

The concept of purchasing power parity (PPP) has long played a prominent role in theoretical and empirical research. The originators of the PPP hypothesis were Spanish scholars of the sixteenth century who taught at the University of Salamanca [see, e. g., Grice-Hutchinson (1952, 1978) and Officer (1982)]. These theologians and jurists, interested also in international commercial activity, proposed the quantity theory of money that, combined with the medieval analysis of foreign exchange rate that ease (scarcity) of money gave it a low (high) value against foreign exchange, lead to the formulation of the PPP hypothesis in a context of radical changes in economic conditions, due to the streams of gold and silver from the New World. Latter, Sweden and France in the second part of the eighteen century, and England in the early nineteenth century, moved from a fixed-rate metallic standard to a floating-rate regime, arising a controversy over the cause of the falling external value of the domestic currency, defending the so-called bullionists a PPP point of view. Following the bullionist period, the PPP hypothesis remained dormant in the literature until the First World War, where severe episodes of hyperinflation and dislocated exchange rates in the belligerent countries stimulated once more the renewed interest in the PPP. Cassel (1922) named this hypothesis, being largely responsible for the popularity enjoyed by the PPP in the 1920s. At the end of the Second World War, a new wave of interest in the PPP hypothesis emerged, when once again attempts were made to determine exchange rates following the wartime suspension of trade and convertibility, leading the move to flexible exchange rates in the early 1970s to yet another intellectual upturn to the PPP hypothesis.

In the last decade or so, important developments in econometrics and the increasing availability of data sets have stimulated the empirical work on PPP. The aim of this paper is to provide a selective survey on some recent empirical attempts to test the validity of PPP to explain exchange-rate movements in the main currencies, as well as the econometric methodology used in such tests.

The paper is organised as follows. Section 2 discusses the basic concepts of the PPP hypothesis, not only examining the absolute and relative PPP versions and the interpretations of PPP, but also considering the choice of the appropriate price index in a PPP calculation and the factors responsible for deviations of actual exchange rates from PPP rates. In Section 3, we focus on the econometric methodology used in the empirical evaluation of the PPP, providing an up-to-date

survey of the burgeoning literature on testing, estimation and model specification in the presence of integrated variables. Section 4 examines some of the empirical evidence which tests the validity of PPP. In Section 5, we offer some encouraging results regarding the forecastability of exchange rate using PPP. Finally, Section 6 provides some concluding remarks.

2. THE PURCHASING POWER PARITY HYPOTHESIS

*“Under the skin of any international economist lies a deep-seated belief in some variant of the PPP theory of exchange rate”
Dornbusch and Krugman (1976, 540)*

2.1. The absolute and relative PPP versions

In its absolute version, PPP theory establishes a relationship between the exchange rate, S_t (expressed as the home currency price of a unit of foreign exchange), and the ratio of domestic and foreign prices (P and P^* respectively), so that

$$S_t = P_t / P_t^* \quad (1)$$

Its implication is that the higher the domestic price level relative to the foreign price level, the higher must be the exchange rate in order to retain purchasing power parity between domestic and foreign currencies.

Factors such as costs of gathering and processing information, transport costs and other obstacles to trade (in particular tariffs and quotas), and market imperfections can limit spatial arbitrage and therefore account for deviations from absolute PPP. Furthermore, it is more than likely that the weights used in the computation of a price level could differ across countries.

For these reasons, a less restrictive relationship between prices and exchange rates is considered. This is the relative PPP, which asserts that the percentage rate of change of the exchange rate will equal the differential between the percentage rates for change of price levels at home and abroad. That is

$$\Delta s_t = \Delta p_t - \Delta p_t^* \quad (2)$$

where Δ denotes first difference and lower-case letters denotes logarithms. Therefore, $100 \Delta s_t$ is the percentage change in S_t since $\Delta s_t = \log(S_t / S_{t-1}) = \log(1 + \Delta S_t / S_{t-1}) \simeq \Delta S_t / S_{t-1}$.

We usually refer to the percentage changes in the price level (Δp_t) as the rate of inflation. Hence, from equation (1), if the domestic inflation rate exceeds the foreign inflation rate, a domestic currency depreciation (i.e., an increase in Δs_t) is required to sustain purchasing power parity between domestic and foreign currencies. Similarly, if the foreign inflation rate exceeds the domestic inflation rate, this will be associated with a domestic currency appreciation (i.e., a reduction in Δs_t). In summary, equation (2) is less strict than equation (1) in allowing domestic and foreign prices (expressed in domestic currency) to differ from each other but still sustains the assumption that these deviations allowing domestic and foreign prices (expressed in domestic currency) will not grow or diminish persistently over time. The relative purchasing power of domestic money *vis-à-vis* foreign money will therefore be fixed over time, with exchange rate changes Δs_t assuring such parity.

The relative version of PPP has a further advantage over absolute PPP in that, as long as the weights used to define the domestic and foreign price indices remain constant over time, then, the two weighting schemes do not need to be the same.

It should be noted that, if absolute PPP holds, then relative PPP will also hold. But if absolute PPP does not hold, then relative PPP still may hold. This is because the level S may not equal P/P^* , but the change in S could still equal the inflation differential, $(\Delta p - \Delta p^*)$.

2.2. Interpretations of PPP

There are three major approaches to PPP: the arbitrage, the expectations, and the monetary approach.

The arbitrage version of PPP was the first well-developed theory of the determination of exchange rates. The basic idea is that exchange rates tend to settle at the level where the purchasing power of a given currency is the same, or at parity, in all countries.

Consider an homogeneous commodity i produced both at home and abroad. Let P_i and P_i^* represent the price of that commodity at home and abroad, stated in home and foreign currency, respectively, and S the exchange rate. Then ignoring information and information and transaction costs and assuming integrated competitive markets, with effective arbitrage, the price of commodity i should be

the same in all locations when quoted in the same currency, say the home currency, i. e.:

$$P_i = P_i^* S \quad (3)$$

This is commonly referred to as the “law of one price”.

If equation (3) does not hold, it would be profitable for the arbitrageurs of commodities to buy the commodity in the country in which it is cheaper and sell it in the country would eventually eliminate the discrepancy between the two sides of equation (3), restoring the equality.

Assume now that the domestic and foreign economies produce a range of η commodities and that the law of one price holds for each of the η commodities. Let P and P^* be a price level at home and abroad quoted in the respective currencies, where $P = g(P_1, P_2, \dots, P_i, \dots, P_\eta)$ and $P^* = g^*(P_1^*, P_2^*, \dots, P_i^*, \dots, P_\eta^*)$. Then by using identical weights in constructing each country's price level (i.e. assuming that the homogeneous-of-degree-one $g(\cdot)$ and $g^*(\cdot)$ functions are the same, we obtain

$$P = SP^* \quad (4)$$

and on rearranging the terms of the exchange rate, we get equation (1). Hence when the price indices in both countries are identical, the law of one price justifies absolute PPP.

The arbitrage approach can also be used to argue that competitive trade will tend to ensure that movements in exchange rate will be such as to compensate for differences in national inflation rates (i.e., relative PPP).

Whereas the arbitrage approach to PPP concentrates solely on trade in commodities, the expectations approach integrates parity conditions in the commodity and financial (bond) markets. This approach which is also known as the “efficient market approach” (see Roll, 1979), is based on the Fisher hypothesis and on the assumption of uncovered interest parity.

The Fisher hypothesis postulates that a country's nominal interest rate should equal its real interest rate plus the expected rate of inflation. Thus,

$$i = r + \Delta p^e \quad (5a)$$

$$i^* = r^* + \Delta p^{*e} \quad (5b)$$

where i is the nominal interest rate, r is the real interest rate, Δp^e is the expected change in the natural logarithm of the price level ($\Delta p^e = \log P_{t+1}^e - \log P_t = p_{t+1}^* - p_t$), and an asterisk (*) denotes a foreign variable.

Uncovered interest parity requires that the nominal interest differential between a domestic currency investment and a foreign currency investment is equal to the expected change in the logarithm of the exchange rate ($\Delta s^e = \log S_{t+1}^e - \log S_t = s_{t+1}^e - s_t$).

$$\Delta s^e = i - i^* \quad (6)$$

But international investors are concerned with real, not nominal, returns on their assets. In attempting to maximise the real return on their assets, they transfer capital from a country with a lower interest rate to one with a higher real rate. Therefore, abstracting from transaction costs, riskiness of returns, and taxation, this arbitrage process results in the equalisation of real interest rates across countries:

$$r = r^* \quad (7)$$

By subtracting (5b) from (5a), using (7) and (6), and rearranging, we obtain

$$\Delta s^e = \Delta p^e - \Delta p^{*e} \quad (8)$$

Equation (8) provides a relative PPP theory in which all variables take on their expected value rather than the current value.

If we assume certainty about the future (ie., $p_{t+1}^e = P_{t+1}$, $P_{t+1}^{*e} = p_{t+1}^*$ and $s_{t+1}^e = s_{t+1}$), then

$$\Delta s = \Delta p - \Delta p^* \quad (9)$$

which is an expression of relative PPP.

Equation (9) can be rewritten as follows,

$$s_{t+1} - p_{t+1} + p_{t+1}^* = s_t - p_t + p_t^* \quad (10)$$

In the terminology of the efficient market literature, equation (10) means that all the information relevant for determining the real exchange rate next period is already fully reflected in the current real exchange rate.

When the future is uncertain, if we assume that expected values in equation (8) are formed rationally, we obtain

$$\Delta s = \Delta p - \Delta p^* + \varepsilon \quad (11)$$

or

$$\Delta s - \Delta p + \Delta p^* = \varepsilon \quad (12)$$

where ε is a composite white-noise error from rational expectations. From equation (12) we see that deviations from PPP (i.e., the real exchange rate) may be characterised as a martingale process or, more particularly, as a random walk (see Roll, 1979).

The monetary approach to PPP emphasises relative money conditions. This approach assumes some sort of neutrality of money to hold at least in the long run. That is, a change in the money supply in one country, with no change in the other country, induces proportional changes in the nominal variables of that country, including the exchange rate. PPP can be viewed from this perspective as an implication of this neutrality proposition

2.3. The price index issue

The formulation of the PPP theory in equations (1) and (2) does not specify which price measurement should be used in the computation. Since most published measures of price are in the form of indices, the controversy about the choice of the adequate measures in the literature on PPP is conducted mainly in terms of price indices, rather than in terms of price levels.

The various interpretations of PPP are relevant for the selection of appropriate price indices. For those who consider arbitrage as the motivating force behind the PPP relationship, the logical choice is the price index of traded goods: In contrast, the monetary approach to PPP requires the use of a broad price index, encompassing a very large number of goods, both traded and non-traded.

Regarding this point Cassel (1928, p 37) states the need to use “a general index figure representing as far as possible the whole mass of commodities marketed in the country”, and Keynes argued that if the price levels taken into account were only those of commodities entering into international trade, then the theory was “little more than a truism” (Keynes, 1923, 1971, p. 75).

Four alternative price indices have traditionally been used as possible candidates for the comparison of the PPP equations (1) and (2): consumer price indices (CPIs), gross domestic product (GDP) deflators, wage rate indices (WRIs), and wholesale price indices (WPIs). The first three include a broad group of goods and services, while the last one represents a sort of compromise with the arbitrage interpretation of PPP due to the large share of tradeable goods.

The most commonly used price indices for PPP calculations are CPIs. The periodic publication of data on CPI behaviour for almost every country is an advantage of this index. However, it can be subjected to direct distortions stemming from price controls. The GDP deflator is not subject to such distortions and is thought to provide a good indicator of changes in competitiveness in production (see Officer, 1982 and Barro, 1983). Some authors (see, e. g., Artus, 1978 and Artus and Knight, 1984) prefer to use unit labour costs, since it is argued that relative labour costs are more stable than relative goods price (Artus, 1978; Officer, 1982). Nevertheless, the WRIs also have some drawbacks (they are highly one factor of production, and they are only available for some time frequency—generally on a yearly basis). The WPIs are also subject to criticism. In addition to Keynes’s criticism that relative price parities calculated from these indices come close to the actual exchange rate (due to the inclusion of highly homegeneous traded goods whose prices tend to be equated across countries when expressed in a common currency), resulting in a spurious verification of the theory, the use of only tradeable goods raises other problems. First, prices of tradeable goods may be set in the short run to maintain competitiveness in world markets, regardless of overall domestic prices and cost levels. Second, computations of PPP based on WPIs (as well as other indices) may be distorted by the use of different weights across countries. Third, practically all tradeable goods can be considered to be differentiated by country of production, if only because of differences in quality, delivery terms, etc. There is no reason to expect the law of one price to apply to them even in the long run.

2.4. Deviations from PPP

From the statistical point of view, the fact that actual price indices are calculated from individual prices of only a sample of commodities rather than all commodities in the economy (Pigou, 1922, pp. 67-68), and the possibility of different weighting schemes in different countries arising from differences in tastes, economic structures and accounting practices (Katseli-Papaefstratiou, 1979. p. 5) can restrict the validity of the PPP theory.

Regarding the economic reasons for deviation from PPP, we can distinguish between the short run and the long run. In the short run, the existence of transportation and information costs can make arbitrage difficult or even impossible. More fundamentally, exchange rates and commodity prices are determined in different kinds of markets. The prices are not as flexible as financial asset prices (the exchange rate is the price of two moneys). This different speed of adjustment between exchange rates and prices can explain short-run deviation from PPP. However, the fact that expectations play a much smaller role in goods and services markets (apart from primary commodities) than in the foreign exchange market implies that "... in periods during which there is ample "news" [i.e., unanticipated changes] which causes large fluctuations in exchange rates there will also be large deviations from purchasing power parities" (Frenkel, 1983, p. 27). The nature of adjustment back to the norm will depend on the degree to which news is seen as indicating permanent or transitory change (see Booth *et al.*, 1985).

In the long run, problems such as the productivity bias can be important: Balassa (1964) and Samuelson (1964) argue that different sectoral rates of productivity growth change real costs and relative prices, and therefore bring about divergences in PPP. The relative price level is high in high productivity (high income) countries compared with low productivity (low income) countries, and it rises rapidly in fast-growing countries compared with slow-growing countries. Although cross-section evidence in favour of this productivity-bias hypothesis is strong (see Balassa, 1964, and Officer, 1976), Hsieh (1982) shows that it is confirmed by time series tests¹.

Factors such as the existence of price contracts and/or rationing, changes in the structure of relative prices in the domestic and foreign economies, monopolistic and oligopolistic forces, product differentiation, trading restrictions (e.g. tariffs and

¹ Hsieh (1982) suggests that this occurs because country-specific factors such as taste vary relatively little over time.

quotas on imports), changes in consumers' preferences away from the home country's goods towards the foreign country's goods, and a natural resource discovery can also account for persistent deviations from PPP.

3. ECONOMETRIC ISSUES

“At the present stage of development in Economics it is probably an advantage to have different groups looking at the same problem from different viewpoints, so that their conclusions can be compared and possibly then form the basis for a new compressive model”
Granger (1990, 1)

3.1. Time series econometrics

Economic theory generally deals with equilibrium relationships. Most empirical econometric studies are an attempt to evaluate such relationships by summarising economic time series using statistical analysis.

To apply standard inference procedures in a dynamic time series model we need the various variables to be stationary, since the majority of econometric theory is built upon the assumption of stationarity, meaning a process whose means and variances are constant over time. However, in applied research we usually find integrated variables, which are a specific class of non-stationary variables with important economic and statistical properties: the variance increases over time and successive observations are highly interdependent. These are derived from the presence of unit roots which give rise to stochastic trends, as opposed to pure deterministic trends, with innovations to an integrated process being permanent instead of transient.

Statisticians have been aware for many years of the existence of integrated series and, in fact, Box and Jenkins (1970) argue that a non-stationary series can be transformed into a stationary one by successive differencing of the series. Therefore, from their point of view, the differencing operation seemed to be a prerequisite for econometric modelling both from an univariate and a multivariate perspective.

After the seminal paper by Engle and Granger (1987), cointegration techniques have been a dominant force in applied macroeconomics (see, e. g, McKenzie, 1997). The key motivation for using the cointegration analysis is to avoid spurious regression results. In addition, cointegration techniques play a useful role in identifying meaningful long-run economic relationships among non-stationary variables.

The literature on cointegration and unit roots is surveyed in Dolado, Jenkinson and Sosvilla-Rivero (1990) or Hendry and Juselius (2000, 2001).

As it is well known, a previous step in cointegration analysis consists of testing the order of integration of the variables. Therefore, we start by reviewing in subsection 3.1.1 several alternative tests for the existence of unit roots. Subsection 3.1.2 introduces the concept of cointegration and surveys several tests to determine the existence of long-run equilibrium relationships.

3.1.1. Unit root tests

Several statistical tests for unit roots have been developed to test for stationarity in time series. The most commonly used to test that a pure AR(1) process (with or without drift) has a unit root are the Dickey-Fuller (DF) statistics. These test statistics were proposed by Dickey and Fuller (1979).

They consider the three following alternative data generating processes (DGP) of a time series:

$$y_t = \rho_n y_{t-1} + \varepsilon_t \quad (13)$$

$$y_t = \mu_c + \rho_c y_{t-1} + \varepsilon_t \quad (14)$$

$$y_t = \mu_{c\tau} + \gamma_\tau + \rho_{c\tau} y_{t-1} + \varepsilon_t \quad (15)$$

where $\varepsilon_t \sim iid(0, \sigma_\varepsilon^2)$, t is a time trend and the initial condition, y_0 is assumed to be a known constant (zero, without loss of generality). For equation (13), if $\rho_n < 1$, then the DGP is a stationary zero-mean AR(1) process and if $\rho_n = 1$, then the GDP is a pure random walk. For equation (14), if $\rho_c < 1$, then the DGP is a stationary AR(1) process with mean $\mu_c / (1 - \rho_c)$ and if $\rho_n = 1$, then the GDP is a random walk with a drift μ_n . Finally, for equation (15), if $\rho_{c\tau} < 1$, then the DGP is a trend-stationary AR(1) process with mean $\frac{\mu_{c\tau}}{1 - \rho_{c\tau}} + \gamma_{c\tau} \sum_{j=0}^t [\rho_{c\tau}^j (t - j)]$ and if $\rho_{c\tau} = 1$, then the GDP is a random walk with a drift changing over time.

The tests are carried out by estimating the following equations:

$$\Delta y_t = (\rho_n - 1)y_{t-1} + \varepsilon_t \quad (13')$$

$$\Delta y_t = \beta_{0c} + (\rho_c - 1)y_{t-1} + \varepsilon_t \quad (14')$$

$$\Delta y_t = \beta_{0c\tau} + \beta_{1c\tau}t + (\rho_{c\tau} - 1)y_{t-1} + \varepsilon_t \quad (15')$$

The tests are implemented though the usual t -statistic on the estimated $(\rho-1)$. They are denoted τ , τ_μ and τ_τ , respectively. Given that under the null hypothesis this test statistic does not have the standard t distribution, Dickey and Fuller (1979) simulated critical values for selected sample sizes. More extensive critical values are reported by MacKinnon (1991, 1994).

Hitherto, we have assumed that the DGP is a pure AR(1) process. If the series is correlated at higher order lag, the assumption of white noise disturbance is violated. Dickey and Fuller (1979) have shown that we can augment the basic regression models (13')-(15') with p lags of Δy_t :

$$\Delta y_t = (\rho_n - 1)y_{t-1} + \sum_{i=1}^p \alpha_i \Delta y_{t-i} + \varepsilon_t \quad (13'')$$

$$\Delta y_t = \beta_{0c} + (\rho_c - 1)y_{t-1} + \sum_{i=1}^p \alpha_i \Delta y_{t-i} + \varepsilon_t \quad (14'')$$

$$\Delta y_t = \beta_{0c\tau}t + \beta_{1c\tau}t + (\rho_{c\tau} - 1)y_{t-1} + \sum_{i=1}^p \alpha_i \Delta y_{t-i} + \varepsilon_t \quad (15'')$$

The tests are based on the t -ratio on $(\hat{\rho}-1)$ and are known as “Augmented Dickey Fuller” (ADF) statistics. The critical values are the same as those discussed for the DF statistics, since the asymptotic distributions of the t -statistics on $(\hat{\rho}-1)$ is independent of the number of lagged first differences included in the ADF regression. Regarding the lag length selection, p should be sufficiently large to remove serial correlation in the residuals. Here we can make use the Akaike information criterion (AIC) or the Schwarz Bayesian information criterion (BIC). Alternatively, we can follow Hall (1994) general to specific sequential rule, starting with a large value of p (p_{max}), testing the significance of the last coefficient and reducing p iteratively until a significant statistic is encountered.

An alternative approach to dealing with autocorrelation has been presented by Phillips (1987) and Phillips and Perron (1988). Rather than including extra lags of Δy_t (as in the Augmented Dickey Fuller test), they suggest amending these statistics to allow weak dependence and heterogeneity in ε_t . Under such general conditions, a wide class of DGP's for ε_t such as most finite order ARIMA(p,0,q) models, can be allowed. The procedure consists of computing the DF statistics and the using some non-parametric adjustment of τ , τ_μ and τ_τ in order to eliminate the dependence of their limiting distributions on additional nuisance parameters

stemming from the ARIMA process followed by the error terms. Their adjusted counterparts are denoted $Z(\tau)$, $Z(\tau_\mu)$ and $Z(\tau_\tau)$ respectively.

For regression model (13'), Phillips and Perron (PP) define

$$Z(\tau) = (\hat{V} / \hat{V}_{Tm})\tau - 0.5(\hat{V}_{Tm}^2 / \hat{V}^2) \left\{ \hat{V}_{Tm}^2 (T^{-2} \sum_{i=2}^T (y_{t-1}^2)) \right\}^{-1/2}$$

where T is the sample size and m is the number of estimated autocorrelations; $\hat{V}$ and τ are, respectively, the sample variance of the residuals and the t -statistic associated with $(\rho_c - 1)$ from the regression (13'); and $\hat{V}_{Tm}^2$ is the long-run variance estimated as

$$\hat{V}_{Tm}^2 = T^{-1} \sum_{t=1}^T \varepsilon_t^2 + 2T^{-1} \sum_{s=1}^l w_{sm} \sum_{t=s+1}^T \hat{\varepsilon}_t \hat{\varepsilon}_{t-s}$$

where $\hat{\varepsilon}$ are the residuals from the regression (13') and where the triangular kernel

$$w_{lm} = [1 - s(m+1)], l = 1, \dots, m$$

is used to ensure that the estimate of the variance $\hat{V}_{Tm}^2$ is positive (see Newey and West, 1987).

For regression model (14'), the corresponding statistic is

$$Z(\tau_\mu) = (\hat{V} / \hat{V}_{Tm})\tau_\mu - 0.5(\hat{V}_{Tm}^2 / \hat{V}^2) T \left\{ \hat{V}_{Tm}^2 \sum_2^T (y_t - \bar{y}_{-1})^2 \right\}^{-1/2}$$

where $\hat{V}$ and $\hat{V}_{Tm}^2$ are defined as above, but with residuals from equation (14'); $\bar{y}_{-1} = (T-1)^{-1} \sum_2^T y_{t-1}$, and τ_μ is the t -statistic associated with $(\rho_n - 1)$ from the regression (14').

Finally, for regression model (15') we have

$$Z(\tau_\tau) = (\hat{V} / \hat{V}_{Tm})\tau_\tau - (\hat{V}_{Tm}^2 - \hat{V}^2) T^3 \left\{ 4\hat{V}_{Tm} [D_{xx}]^{1/2} \right\}^{-1}$$

where $\hat{v}$ INCRUSTARINCRUSTAR and $\hat{v}_{Tm}^2$ are defined as above, but with the residual obtained from the estimation of (3'). τ_t is the t -statistic associated with $(\rho_{cr} - 1)$ from the regression (15'). D_{xx} is the determinant of the regressor cross product matrix, given by

$$D_{xx} = \left[T^2(T^2 - 1)/12 \right] \sum y_{t-1}^2 - T \left(\sum ty_{t-1} \right)^2 + \\ + T(T+1) \sum ty_{t-1} \sum y_{t-1} - \left[T(T+1)(2T+1)/6 \right] \left(\sum y_{t-1} \right)^2$$

The Phillips and Perron statistics have the same limiting distributions as the corresponding DF and ADF statistics, provided that $m \rightarrow \infty$ as $T \rightarrow \infty$, such that $m/T^{1/4} \rightarrow 0$.

Both, the ADF and PP tests, take a unit root as the null hypothesis. Kwiatkowski, Phillips, Schmidt and Shin (1992) provide an alternative test (known as the KPSS test) for testing the null of stationarity against the alternative of a unit root. This method considers models with constant terms, and either with or without a deterministic trend (their v_τ and v_μ statistics, respectively).

Formally, the KPSS test is given by:

$$LM = \frac{\sum_{t=1}^T \Phi_t^2}{\hat{\sigma}_e^2} \quad (16)$$

where $\Phi_t^2 = \sum_{i=1}^t e_i$ is the running partial sum of the residuals and $\hat{\sigma}_e^2$ is the estimated residual variance from the regression:

$$y_t = \alpha + \varepsilon_t \quad (17)$$

for the model without trend, or

$$y_t = \alpha + \beta t + \varepsilon_t \quad (18)$$

for the model with trend.

The critical values for the KPSS tests are given in Kwiatkowski, Phillips, Schmidt and Shin (1992). Recently, it have been argued that confirmatory analysis (i. e., applying ADF or PP unit rot tests in conjunction with KPSS stationarity tests)

may in some cases lead to a better description of the series, improving upon separate use of each type of tests (see, e. g., Maddala and Kim, 1998). If the ADF/PP tests reject the null while the KPSS test fails to do so, the results of both tests are consistent, suggesting that a given series is stationary. Alternatively, if the ADF/PP tests fail to reject the null while the KPSS test does reject it, both approaches give consistent results, and one may conclude in this case that the series is not stationary. Finally, both ADF/PP and KPSS tests fail to reject the respective nulls or both reject their nulls, the results are inconclusive.

Finally, given that conclusions drawn from unit root tests may well be sensitive to structural breaks in the underlying stochastic process a number of tests have been proposed for unit roots under structural change (see, e. g. Maddala and Kim, 1998). Following Perron (1989, 1997), we can allow for the possibility of a one-time structural change in the trend function occurring at time T_b . Three situations are considered: a change in the intercept, a change in both the intercept and the slope, and a change in the slope. Regarding the transition to the new trend path, and following Perron (1989), two models are evaluated: the “additive outlier model” (AOM) and the “innovational outlier model” (IOM). While the AOM specifies that the change to the new trend function occurs instantaneously (with no further effect on future observations), in the IOM that change takes place gradually (feeding back into the process dynamics).

In the case of the IOM, the unit-root test is performed using the t -statistic for testing $\alpha = 1$ in the following regressions:

$$\text{IOM-1: } y_t = \mu + \beta t + \theta DU_t + \delta D(T_b)_t + \alpha y_{t-1} + \sum_{i=1}^p c_i \Delta y_{t-i} + \varepsilon_t \quad (19)$$

$$\text{IOM-2: } y_t = \mu + \beta t + \theta DU_t + \gamma T_t + \delta D(T_b)_t + \alpha y_{t-1} + \sum_{i=1}^p c_i \Delta y_{t-i} + \varepsilon_t \quad (20)$$

$$\text{IOM-3: } y_t = \mu + \beta t + \gamma DT_t + \alpha y_{t-1} + \sum_{i=1}^p c_i \Delta y_{t-i} + \varepsilon_t \quad (21)$$

where $D(T_b)_t = 1$ if $t = T_b + 1$ (0 otherwise); $DU_t = 1$ if $t > T_b$ (0 otherwise); and $DT_t = (t - T_b)$ if $t > T_b$ (0 otherwise). In equation (19) we allow for a one-time change in the intercept of the trend function, while in equation (20) we allow for both a change in the intercept and in the slope of the trend function, whereas in equation (21) there is a change in the slope of the trend function.

Regarding the AOM, the following two-step procedure is used. First, the series is detrended using the following regressions:

$$\text{AOM-1: } y_t = \mu + \beta t + \theta DU_t + \tilde{y}_t \quad (22)$$

$$\text{AOM-2: } y_t = \mu + \beta t + \theta DU_t + \gamma DT_t + \tilde{y}_t \quad (23)$$

$$\text{AOM-3: } y_t = \mu + \beta t + \gamma DT_t + \tilde{y}_t \quad (24)$$

where $\tilde{y}_t$ is accordingly defined as the detrended series. As can be seen, in equation (22) we allow for a one-time change in the intercept of the trend function, in equation (23) we allow for both the change in the intercept and the slope of the trend function to take place simultaneously, and in equation (24) we allow for a change only in the slope of the trend function.

For models IOM-1 and IOM-2, the test is then performed using the t -statistic for testing $\alpha = 1$ in the regression:

$$\tilde{y}_t = \alpha \tilde{y}_{t-1} + \sum_{j=0}^p d_j D(T_b)_{t-j} + \sum_{i=1}^p c_i \Delta \tilde{y}_{t-i} + \varepsilon_t \quad (25)$$

while for model IOM-3, the second step is of the form:

$$\tilde{y}_t = \alpha \tilde{y}_{t-1} + \sum_{i=1}^p c_i \Delta \tilde{y}_{t-i} + \varepsilon_t \quad (26)$$

Note that in regressions (19) to (26), the break date (T_b) and the truncation lag (p) are treated as unknown. Therefore, to carry out the test procedure, one need to consider a method to choose T_b and k . In order to select the break date endogenously, we can consider the procedure whereby T_b is selected as the value, for all possible break points, which minimises the test statistic for testing $\alpha=1$ in the appropriate autocorrelation specification (see Zivot and Andrews, 1992). Regarding the truncation lag parameter (p), we use a general-to-specific recursive approach based on the value of the t -statistic on the coefficient associated with the last lag in the estimated autocorrelation (see Perron, 1989).

3.1.2. Cointegration

Consider two time series y_t and x_t which are both $I(d)$ (i. e., they have comparable long-run properties). In general, any linear combination of y_t and x_t will be also $I(d)$. If, however, there exists a vector $(1, -\beta)'$, such that the combination

$$z_t = y_t - \alpha - \beta x_t \quad (27)$$

is $I(d-b)$, $b>0$, then Engel and Granger (1987) define y_t and x_t as cointegrated of order $(d-b)$ [or $(y_t, x_t)' \sim CI(d, b)$] with $(1, -\beta)'$ called the cointegrating vector. Notice that a constant term has been included in (19) in order to allow for the possibility that z_t may have a non-zero mean.

The concept of cointegration tries to mimic the existence of a long-run equilibrium to which an economic system converges over time. In the case of PPP, from equation (1) we have:

$$s_t = \alpha + \beta(p - p^*)_t + \varepsilon_t \quad (28)$$

where s_t is the logarithm of the spot exchange rate at time t and p_t (p_t^*) is the logarithm of the domestic (foreign) price level. Therefore, $z_t = s_t - \alpha - \beta(p - p^*)_t$ can be interpreted as the equilibrium error (i. e., the distance that the exchange rate is away from equilibrium at any point of time).

Engle and Granger also show that if s_t and $(p - p^*)_t$ are cointegrated $CI(1,1)$, then there must exist an error correction model (ECM) representation of the following form:

$$\Delta s_t = \mu + \sum_{i=0}^p \phi_i \Delta s_{t-i} + \sum_{j=1}^q \gamma_j \Delta(p - p^*)_{t-j} + \theta z_{t-1} + \xi_t \quad (29)$$

where ξ_t is a sequence of independent and identically distributed random variables with mean zero and variance σ_ξ^2 . Furthermore, they prove the converse result that an ECM generates cointegrated series.

Note that the term z_{t-1} in equation (21) represents the extent of disequilibrium between levels of s and $(p - p^*)$ in the previous period. The ECM states that changes in s_t depend not only on changes in $(p - p^*)_t$, but also on the extent of disequilibrium

between levels of s and $(p - p^*)$. Therefore, the ECM could be seen as capturing the dynamics of the system whilst incorporating the equilibrium suggested by economic theory (see Hendry, 1995).

Based upon the concept of cointegration and on its closely related concept of ECM representation, Engle and Granger (1987) suggest a two-step estimation procedure for dynamic modelling which has become very popular in applied research. In the cases of PPP, if s_t and $(p - p^*)_t$ are both $I(1)$, then the procedure goes as follows:

- i) First, in order to test whether the series are cointegrated, the cointegrating regression (28) is estimated by ordinary least squares (OLS) and it is tested whether the cointegrating residuals' $\hat{z}_t = s_t - \hat{\alpha} - \hat{\beta}(p - p^*)_t$ are $I(0)$.
- ii) Finally, the residuals $\hat{z}_t$ are entered into the ECM (29), where now all the variables are $I(0)$ and conventional modelling strategies can be applied.

Regarding the first step, Engle and Granger (1987) suggest seven alternative tests for determining if z_t is stationary. The two most widely used are the Durbin-Watson statistic for the cointegrating regression (CRDW) and the ADF statistic for the cointegrating residuals (CRADF).

The DW statistic for equation (20) will approach zero if the cointegrating residuals contain an autoregressive unit root, and thus the test rejects the null hypothesis of non-cointegration if the CRDW is significantly greater than zero.

The CRADF statistic is based upon the OLS estimation of

$$\Delta \hat{z}_t = \eta_1 \hat{z}_{t-1} + \sum_{i=1}^p \eta_{2i} \Delta \hat{z}_{t-i} + \varepsilon_t \quad (30)$$

where again p is selected on the basis of being sufficiently large to ensure that ε_t is a close approximation to white noise. The t-ratio statistic on $\hat{\eta}_1$ is the CRADF statistic.

Engle and Granger (1987, p. 269) present the critical values for the CRDW and the CRADF statistics generated from Monte Carlo simulations of 100 simulations for a bivariate case as PPP in equation (28) (i. e., for one dependent variable and one independent variable in the cointegrating regression). Engle and

Yoo (1987) produce expanded critical values for CRDW and CRADF statistics for 50, 100 and 200 observations, and for systems of up to five variables. Finally, MacKinnon (1991) provides an approximation formula for computing critical values for all small samples, while MacKinnon (1994) expands his methodology to calculate both asymptotic and finite sample critical values.

As can be seen, both CRDW and CRADF statistics test the null of no-cointegration against the alternative of cointegration. It has been often argued that cointegration would be a more natural choice, but there are only a few tests for the null of cointegration. Shin (1994), for example, propose a test based on the cointegrating residuals which is an extension of the LM test of KPSS univariate stationarity. The procedure consists of introducing past and future values of $\Delta(p - p^*)_t$ in equation (28), so that the cointegrating regression becomes

$$s_t = \alpha + \beta(p - p^*)_t + \sum_{i=-n}^n \varphi_i \Delta \hat{z}_{t-i} + \varepsilon_t \quad (31)$$

Applying OLS to the modified regression (31) will yield efficient estimates (see Saikkonen, 1991). A version of (16) may be constructed with the residuals $\hat{\varepsilon}_t$ from (31).

A test for absolute PPP consists in testing the joint hypothesis $H_0: (\alpha, \beta) = (0, 1)$ in equation (28). Stock (1987) has shown that if two I(1) series are cointegrated, then the OLS estimates from equation (28) provide “super-consistent” estimates of the cointegrating vector (α, β) . Nevertheless, the joint dependence of most aggregate time series and their non-stationarity invalidates the routine application of many standard statistical procedures in equation (28). Phillips and Hansen (1990) present a class of Wald tests which are modified by semiparametric corrections for serial correlation and for endogeneity. The resulting test statistics (termed fully-modified Wald tests) have limiting χ^2 distributions and therefore allows inference to proceed conventionally (for an explanation see, e. g., Ngama and Sosvilla-Rivero, 1991). Notice that in equation (28) we are imposing that p_t and p_t^* have the same coefficient (β), allowing us to combine the domestic and the foreign price levels forming a differential in prices. Alternatively, we would have assumed that the coefficients are different, so testing PPP should be based on estimates of the following equation :

$$s_t = \alpha + \beta p_t + \beta^* p_t^* + \varepsilon_t \quad (28')$$

In this case, a test for absolute PPP consists in testing the joint hypothesis $H_0: (\alpha, \beta, \beta^*) = (0, 1, -1)$ in equation (28'). However, a distinction is often made between the test that β and β^* are equal and of opposite sign (the so-called “symmetric condition”) and the test that they are equal to unity and minus unity, respectively (the so-called “proportionality condition”).

As an alternative two-step Engle-Granger procedure, Banerjee *et al.* (1986) propose a single-step dynamic model approach, based upon the t -ratio of the coefficient on the error-correction term θ in the following ECM for the PPP hypothesis:

$$\Delta s_t = \mu + \sum_{i=0}^p \phi_i \Delta s_{t-i} + \sum_{j=1}^q \gamma_j \Delta(p - p^*)_{t-j} + \theta[s_{t-1} - \alpha - \beta(p - p^*)_{t-1}] + \xi_t \quad (32)$$

The t -ratio is denoted the ECM statistic. Kremers, Ericsson and Dolado (1992) find that the ECM statistic can generate more powerful tests than those based upon the DF statistic applied to the residuals of a static cointegrating relationship such as (28). More generally, if we think that $(p - p^*)_t$ may be only weakly exogeneous to the parameters of interests, Banerjee, Dolado and Mestre (1998) recommend estimating the following (unrestricted) ECM regression by OLS:

$$\Delta s_t = \mu + \sum_{i=0}^p \phi_i \Delta s_{t-i} + \sum_{j=1}^q \gamma_j \Delta(p - p^*)_{t-j} + \theta_1 s_{t-1} + \theta_2 (p - p^*)_{t-1} + \xi_t \quad (33)$$

When θ_1 exceeds the critical values (provided in Banerjee, Dolado and Mestre, 1998), the null hypothesis of non-cointegration is rejected.

The parameters in the cointegrating regression (28) may not be constant through time. Gregory and Hansen (1996) generalised the usual residual based cointegration tests, allowing for a broader view of cointegration when they consider an alternative hypothesis in which the cointegration vector suffers shift at an unknown time. The new regression model is:

$$s_t = \alpha_1 + \alpha_2 D(t_0) + \beta_1 (p - p^*)_t + \beta_2 (p - p^*)_t D(t_0) + \varepsilon_t \quad (34)$$

where $D(t_0)$ is a dummy variable such that $D(t_0) = 0$ if $0 < t \leq t_0$ and $D(t_0) = 1$ if $t_0 < t \leq T$. The test for cointegration is conducted by testing for unit roots (for instance, with an ADF test) on the residuals $\hat{z}_t = s_t - \hat{\alpha}_1 - \hat{\alpha}_2 D(t_0) - \hat{\beta}_1 (p - p^*)_t - \hat{\beta}_2 (p - p^*)_t D(t_0)$ for each t_0 . Gregory and Hansen

(1996) propose and tabulate the critical values of the test statistic $ADF^* = \inf_{1 \leq t_0 \leq T} \{ADF(t_0)\}$. The null hypothesis of no cointegration and no structural break is rejected if the statistic ADF^* is smaller than the corresponding critical value. In this case, the structural break can be located at time t^* where the *inf* of the ADF test is obtained.

Johansen (1988) and Johansen and Juselius (1990) develop a maximum likelihood estimation procedure that has several advantages on the two-step regression procedure suggested by Engle and Granger. It relaxes the assumption that the cointegrating vector is unique and it takes into account the error structure of the underlying process.

Johansen considers the p -th order autoregressive representation of X_t

$$X_t = \Pi_1 X_{t-1} + \Pi_2 X_{t-2} + \dots + \Pi_p X_{t-p} + \varepsilon_t \quad (35)$$

which, following a similar procedure to the ADF test, can be re-parameterised as

$$\Delta X_t = \tilde{\Pi}'_1 \Delta X_{t-1} + \dots + \tilde{\Pi}'_{p+1} \Delta X_{t-p+1} + \tilde{\Pi}'_p X_{t-p} + \varepsilon_t \quad (36)$$

where $\tilde{\Pi}'_p = -\Pi(1) = -(\Pi_1 + \dots + \Pi_p)$. To estimate $\tilde{\Pi}'_p$ by maximum-likelihood, we estimate by OLS the following regressions

$$\Delta X_t = \Gamma_{01} \Delta X_{t-1} + \dots + \Gamma_{0k-1} \Delta X_{t-k+1} + e_{0t} \quad (37)$$

and

$$\Delta X_{t-p} = \Gamma_{11} \Delta X_{t-1} + \dots + \Gamma_{1k-1} \Delta X_{t-k+1} + e_{1t} \quad (38)$$

and compute the product moment matrices of the residuals

$$\hat{M}_{ij} = T^{-1} \sum_{t=1}^T \hat{e}_{it} \hat{e}_{jt}' ; i, j = 0, 1 \quad (39)$$

To test of the null hypothesis $H_0: \Pi'_p = B\Gamma'$, i.e. there are at most r cointegrating vectors, can be conducted using either of the following two test statistics

$$\lambda_{ir}(r) = -T \sum_{i=r+1}^{\rho} \ln(1 - \hat{\lambda}_i) \quad (40)$$

$$\lambda_{\max}(r, r+1) = -T \ln(1 - \hat{\lambda}_{r+1}) \quad (41)$$

where $\hat{\lambda}_{r+1}, \dots, \hat{\lambda}_{\rho}$ are the $\rho-r$ smallest eigenvalues of $\hat{M}_{10}\hat{M}_{00}\hat{M}_{01}$ with respect to M_{11} , obtained from the determinant

$$|\hat{\lambda}\hat{M}_{11} - \hat{M}_{10}\hat{M}_{00}\hat{M}_{01}| = 0 \quad (42)$$

The statistic in (40), known as the trace statistic, tests the null hypothesis that the number of cointegrating vectors is less than or equal to r against a general alternative. On the other hand, the statistic in (41), known as the maximum eigenvalue statistic, test a null of r cointegrating vectors against the specific alternative of $r+1$. Osterwald-Lenum (1992) offers critical values for both tests using Monte Carlo simulations.

Finally, an alternative approach, which has certain advantages over both the OLS and the maximum likelihood procedures, has been proposed by Stock and Watson (1993). Their dynamic ordinary least squares (DOLS) improves on OLS by coping with small sample and dynamic sources of bias. The Johansen method, being a full information technique, is exposed to the problem that parameter estimates in one equation are affected by any misspecification in other equations. The Stock Watson method is, by contrast, a robust single equation approach which corrects for regressor endogeneity by the inclusion of leads and lags of first differences of the regressors. In addition it has the same asymptotic optimality properties as the Johansen distribution. For the PPP hypothesis the DOLS regression would be as follows:

$$s_t = \alpha + \beta(p - p^*)_t + \sum_{j=-q_1}^{q_2} \varphi_j \Delta(p - p^*)_{t-j} + \varepsilon_t \quad (28'')$$

where q_1 and q_2 are selected to increase at an appropriate rate with T .

Stock and Watson (1993) suggest correcting for serial correlation in (28'') by using generalised least squares GLS (the so-called dynamic GLS: DGLS).

3.2. Panel estimation with non-stationary data

Standard cross-section analysis focuses on long-run average relationships and ignores the time-series variation in the data. But the variation of variables over time contain additional information, which may be particularly valuable in situations where the cross-section variation in the data is relatively limited.

The panel cointegration approach exploits both cross-section as well as time series variation in the data. Compared to individual time series tests, these panel cointegration tests have higher power. In subsection 3.2.1 we review some tests that have been proposed for the existence of unit roots in panel data, while subsection 3.2.2 surveys several alternative tests for panel cointegration.

3.2.1 Panel unit root tests

Before testing for cointegration, we need to confirm whether the variables are actually non-stationary. A variety of procedures for the analysis of unit roots in panel. One of the first tests was the Levin-Lin (LL) test [Levin and Lin (1993) and Levin, Lin and Chu (2002)]. Their test is based on analysis of the following equation

$$\Delta y_{i,t} = \alpha_i + \delta t + \rho y_{i,t-1} + \sum_{j=1}^p \phi_j \Delta y_{i,t-j} + \varepsilon_{i,t} \quad (43)$$

where $i=1, 2, \dots, N$ indexes cross-section units (e. g., countries) and $t=1, 2, \dots, T$ indexes time periods (e. g., years). Corresponding to the maintained hypothesis of common dynamics, the null hypothesis and alternative hypothesis are given by $H_0 : \rho = 0, H_A : \rho < 0$.

Equation (43) can be estimated using the within estimator and the LL test statistic is based on the usual t -statistic:

$$t_\rho = \frac{\hat{\rho}}{\hat{\sigma}_\rho} \quad (44)$$

Levin and Lin (1993) propose two transformations of the t -statistic t_ρ that are asymptotically normally distributed as N and $T \rightarrow \infty$:

$$LL_1 = \sqrt{1.25} t_{\hat{\rho}} + \sqrt{1.55} N \Rightarrow N(0,1) (as \sqrt{N} / T \rightarrow \infty)$$

$$LL_2 = \sqrt{1.25}[t_{\hat{\beta}} - (\sqrt{N}\mu_{1T} / \sqrt{\mu_{2T}}) \Rightarrow N(0,1)$$

with $\mu_{1T} = -\frac{1}{2} - \frac{1}{2}T^{-1}$ and $\mu_{2T} = \frac{1}{6} + \frac{5}{6}T^{-2}$. The two tests LL_1 and LL_2 coincide asymptotically, but in finite samples they will differ.

In, Pesaran and Shin (1997, 2003) (IPS hereafter) extend the LL test to allow for heterogeneity in the value of ρ_i among observations under the alternative hypothesis. The relevant equation is the following:

$$\Delta y_{i,t} = \alpha_i + \delta_i t + \rho_i y_{i,t-1} + \sum_{j=1}^p \phi_j \Delta y_{i,t-j} + \varepsilon_{i,t} \quad (43')$$

The null and alternative hypotheses are defined as $H_0 : \rho_i = 0, \forall i; H_A : \rho_i < 0, i = 1, 2, \dots, N_1, \rho_i = 0, i = N_1 + 1, N_1 + 2, \dots, N$. Due to the heterogeneity, equation (43') is estimated separately by OLS for each cross-section unit based on T observation. Let t_i ($i = 1, 2, \dots, N$) denote the individual t -statistics for testing unit roots. Then, to test the null of a unit root across all individuals, merely take the average of these t -statistics to obtain the IPS t -bar statistic:

$$\bar{t} = \frac{1}{N} \sum_{i=1}^N t_i \quad (45)$$

Assuming that the cross-section units are independent, IPS propose the use of the standardised t -bar statistic:

$$\Gamma_{\bar{t}} = \frac{\sqrt{N} [\bar{t} - E(t_i / \rho_i = 0)]}{\sqrt{Var(t_i / \rho_i = 0)}} \quad (46)$$

The means $E(t_i / \rho_i = 0)$ and de variance $Var(t_i / \rho_i = 0)$ are obtained from Monte Carlo methods.

IPS also propose an LM-bar test statistic where they compute an average Lagrange multiplier test of the null that the lagged level has no explanatory power ($\rho_i = 0$, for all i) across all individuals. The Monte Carlo results indicate that the t -bar test is somewhat more powerful.

3.2.2 Panel cointegration tests

Consider the following system of cointegrated regressions to test PPP as a long-run equilibrium:

$$s_{i,t} = \alpha_i + \beta(p - p^*)_{i,t} + \varepsilon_{i,t} \quad (47.1)$$

$$(p - p^*)_{i,t} = (p - p^*)_{i,t-1} + \xi_{i,t} \quad (47.2)$$

where as before $i=1, 2, \dots, N$ indexes cross-section units (e. g., countries) and $t=1, 2, \dots, T$ indexes time periods (e. g., years), and where $s_{i,t}$ and $(p - p^*)_{i,t}$ are integrated processes of order one for all i .

The zero mean innovation vector $\zeta_{i,t} = (\varepsilon_{i,t}, \xi_{i,t})'$ satisfies

$$\frac{1}{\sqrt{T}} \sum_{t=1}^{T_T} \zeta_{i,t} \Rightarrow B_i(\Omega), \text{ for all } i \text{ as } T \rightarrow \infty \quad (48)$$

where $B_i(\Omega)$ is a vector Brownian motion with asymptotic covariance Ω .

Kao (1999) derives two types of panel cointegration tests. The first is a DF type test and the second is an ADF type test. Both tests can be calculated from²:

$$\Delta \hat{z}_{i,t} = \rho \hat{z}_{i,t-1} + \sum_{i=1}^p \psi_i \Delta \hat{z}_{i,t-i} + \varepsilon_{i,t} \quad (49)$$

where the residuals $\hat{z}_{i,t} = s_{i,t} - \hat{\alpha}_i - \hat{\beta}(p - p^*)_{i,t}$ are based on the OLS estimation of equation (48). The null and alternative hypotheses are defined as $H_0 : \rho = 1, H_A : \rho < 1$. Kao (1999) propose four DF-type statistics. The first two DF statistics are based on assuming strict exogeneity of the regressors, while the remaining two allow for endogeneity of the regressors. In addition, Kao (1999, pp. 6-11) proposes an ADF test statistic. Finally, the DF statistics, which allow for endogeneity, and the ADF statistic involve deriving some nuisance parameters from the long-run conditional variances Ω . The asymptotic distributions of all tests converge to a standard normal distribution $N(0,1)$ as $T \rightarrow \infty$ and $N \rightarrow \infty$.

² In the case of DF tests all $\psi_i=0$.

Using system (47) and allowing for heterogeneity of the long-run variance matrix Ω_i for each i and heterogeneity of the slope parameters across all cross-section units i , Pedroni (1999) presents a total of seven tests of the null of no cointegration, of which four involve pooling on the within dimension and three on the between dimension. The first category of tests uses the following specification for the null and alternative hypotheses; $H_0: \rho = 1, H_A: \rho < 1$, while the second category uses $H_0: \rho_i = 1, H_A: \rho_i < 1$ for all i , therefore requiring to compute N autorregressive coefficients ρ_i by using equation (48) for each i th cross-section unit. Pedroni's statistics also require estimating some nuisance parameters from the long-run conditional variances Ω_i . Each of the seven test statistics can be rescaled so that it is distributed as a normal distribution. The appropriate factors for these tests are given in Pedroni (1999).

McCoskey and Kao (1998) also develop a residual-based panel test, but in contrast to Kao's tests and Pedroni's tests, takes cointegration as the null hypothesis. The test is given by

$$LM = \frac{1}{N} \sum_{n=1}^N \left\{ \frac{\frac{1}{T^2} \sum_{t=1}^T \Phi_{it}^2}{\sigma_z^2} \right\} \quad (50)$$

where $\Phi_{it} = \sum_{j=1}^t \hat{z}_{ij}$ is the running partial sum of the residuals and $\sigma_z^2 = 1/NT \sum_{i=1}^N \sum_{j=1}^T \hat{z}_{ij}^2$. The residuals $\hat{z}_{i,t} = s_{i,t} - \hat{\alpha}_i - \hat{\beta}(p - p^*)_{i,t}$ can be estimated using either the dynamics OLS [DOLS, built upon the work of Saikkonen (1991) and Stock and Watson (1993)] or the fully modified OLS [FMOLS, based upon Phillips and Hansen (1990)] estimator, both of which correct for serial correlation and endogeneity of regressors. McCoskey and Kao (1998) show that a standardised version of the statistic converges to a normally distributed random variable under the null hypothesis of cointegration.

4. EMPIRICAL EVIDENCE ON PPP

*“I have no data yet. It is a capital mistake to
theorise before one has data. Insensibly one
begins to twist facts to suit theories,
instead of theories to suit facts”
From a dialogue between Sherlock Holmes and
Dr. Watson in A Scandal in Bohemia,
by Sir Arthur Conan Doyle*

4.1. Reviewing the literature on empirical evidence on PPP

In this section we provide a review of the vast empirical literature testing the validity of the PPP hypothesis. This area has proven fruitful ground for applying different estimation methods to different periods and countries. Therefore, we have selected the most relevant contributions classifying them from the point of view of the econometric methods used in the empirical application.

Firstly, we have considered some representative studies from the seventies, when the econometric methodologies applied in these papers (OLS, Instrumental Variables) did not take into account the statistical properties of the time series. Secondly, we examine the contributions of some of the authors who applied modern methods for dealing with nonstationary time series, using the concept of cointegration. Finally, there are relatively fewer papers test the PPP hypothesis using the newest econometric methodology: panel cointegration.

Table 1 contains the name and abbreviations of the currencies and countries analysed in the empirical papers, while Table 2 shows the different econometric tests and methodologies used in such papers.

In Table 3 we report the empirical application under study, giving very summarised information regarding its relevant features. The first column gives the reference of a particular paper. The second contains the currencies examined. The third specifies the period (or subperiods) under investigation, as well as the characteristics of the sample (monthly, annual or quarterly), and span. The fourth column shows the variables that are used: “NER”³ stands for “nominal exchange

³ In many works, there is also information about Real Exchange Rates (RER). As our aim is to provide information of studies that analyse the equations related to the Nominal Exchange Rates, rather than summarizing those that study stationarity properties of RER, and in order to maintain the concreteness of the table, its presence is not always notified.

rate”, “CPI” means “consumer price index”, “WPI” is the “wholesale price index”, “GDPD” is the “deflator of the GDP”, “MPI” stands for “manufacturing price index”, “IPI” is the “industrial price index”, and “TPI” represents the “price index of traded goods”. In the fifth column we report the econometric methodology used in the empirical evaluation. After a short comment on each paper, we provide the results of the analysis regarding the validity of the PPP hypothesis (i. e., if the empirical evidence found is favourable or not to this hypothesis).

Table 1: Currency Abbreviations

<i>Country</i>	<i>Currency</i>	<i>Abbr.</i>	<i>Country</i>	<i>Currency</i>	<i>Abbr.</i>	<i>Country</i>	<i>Currency</i>	<i>Abbr.</i>	<i>Country</i>	<i>Currency</i>	<i>Abbr.</i>
Algeria	Dinar	DZD	Ethiopia	Birr	ETB	Malaysia	Ringgit	MYR	Singapore	Dollar	SGD
Argentina	Peso	ARS	Europe	Ecu	ECU	Mexico	Peso	MXN	Slovakia	Koruna	SKK
Belgium	Franc	BEF	France	Franc	FRF	Nepal	Rupee	NPR	South Korea	Won	KRW
Bolivia	Boliviano	BOB	Germany	Marc	DEM	Netherlands	Guilder	NLG	Spain	Peseta	ESP
Brazil	Real	BRL	Ghana	New cedi	GHC	New Zealand	Dollar	NZD	Sweden	Krona	SEK
Bulgaria	Leva	BGL	Greece	Drachma	GRD	North Korea	Won	KPW	Sri Lanka	Rupee	LKR
Canada	Dollar	CAD	Hungary	Forint	HUF	Norway	Krone	NOK	Switzerland	Franc	CHF
Chile	Peso	CLP	India	Rupee	INR	Pakistan	Rupee	PKR	Thailand	Baht	THB
Colombia	Peso	COP	Indonesia	Rupiah	IDR	Peru	New Sol	PEN	Turkey	Lira	TRL
Czech Republic	Koruna	CZK	Israel	New Shekel	ILS	Philippines	Peso	PHP	United Kingdom	Pound	GBP
Denmark	Krone	DKK	Italy	Lira	ITL	Poland	Zloty	PLN	United States	Dollar	USD
Dominican Rep.	Peso	DOP	Japan	Yen	JPY	Portugal	Escudo	PTE	Uruguay	Peso	UYU
Egypt	Pound	EGP	Kenya	Shilling	KES	Romania	Leu	ROL	Venezuela	Bolivar	VEB

Table 2: Econometric Methodologies and Tests Abbreviations

<i>Abbr.</i>	<i>Econometric Method/Test</i>
ADF	Augmented Dickey Fuller
CRADF	Augmented Dickey Fuller test for Cointegration
CRDF	Dickey Fuller test for Cointegration
CRDW	Durbin Watson test for Cointegration
CRPP	Phillips Perron test for Cointegration
DF	Dickey Fuller
DGLS	Dynamic Generalised Least Squares
DOLS	Dynamic Ordinary Least Squares
ECM	Error Correction Model
FDOLS	Fully Modified Ordinary Least Squares
GLS	Generalised Least Squares
IPS	Im, Pesaran and Shin
IV	Instrumental Variables
KPSS	Kwiatkowski, Phillips, Schmidt and Shin
LL	Levin-Lin
(M)SB	(Modified) Sargan-Bhargava
2S-OLS	Two Stages Ordinary Least Squares
OLS	Ordinary Least Squares
PP	Phillips Perron
SBC	Schwarz's Bayesian Critetion

Table 3: Empirical Evidence on Purchasing Power Parity

<i>AUTHOR</i>	<i>CURRENCIES</i>	<i>PERIOD</i>	<i>VARIABLES</i>	<i>ECONOMETRIC METHOD</i>	<i>CHARACTERISTICS</i>	<i>RESULT</i>
<i>Krugman (1978)</i>	DEM, CHF, FRF, ITL, GBP and USD	Monthly: 1921-1925 1973-1976	NER, WPI	OLS IV	Studies the autocorrelation and suggests that movements on the exchange rates are due to omitted variables	Not Fav. Fav.
<i>Frenkel (1981)</i>	USD, DEM, FRF, GBP	Monthly: 1921-1925 1973-1979	NER, CPI, MPI, WPI	2S-OLS, IV	Studies both absolute and relative PPP	Fav. (first period and last period for EU). Not fav (during the last period for USD)
<i>Miller (1984)</i>	FRF, GBP, DEM, USD	Quarterly 1973-1980	NER, Indices Divisia?	GLS	Finds that deviations in prices are persistent	Fav (within EU)
<i>Edison and Klovland (1987)</i>	NOK and GBP	Annual 1974-1971	NER, GDPD at market prices	Cointegration (CRADF, CRDW)	Focus on l.r. structural factors such as productivity and terms of trade that could cause the simple PPP to fail and s.r cyclical factors that cause temporary deviations from PPP, and suggest an expanded model	Fav. (PPP holds only in the l.r after having taken into account the effects of changes in real, structural favors like the relative levels of productivity and the terms of trade)
<i>Taylor (1988)</i>	GBP, DEM, FRF, CAD, JPY against USD	Monthly 1973:06-1985:12	NER, MPI	Unit Root (DF, ADF) Cointegration (DF, ADF, DW)	Allowance for measurement error and/or transportation costs	Not Fav.
<i>Taylor and McMahon (1988)</i>	DEM, FRF, USD against GBP	Monthly 1921:01-1925:05 Data for Germany: 1921:02-1923:08	NER, WPI	Unit Root (ADF) Cointegration (CRDW, CRADF)	Results are largely invariant to the choice of normalising variable. The failure is attributable to non-stationary and non-fundamental factors during the last year before Britain's return to the Gold Standard	Fav. Except for USD-GBP
<i>Mikkelsen (1989)</i>	ARS, CLP, MXN, BRL, UYU, and PEN	Quarterly 1948-1988	NER Relative WPI	Unit Root (DF, ADF) Cointegration (CRADF)	e and p in Brazil are I(2) In Peru, e is I(1) while p is I(2)	Fav. for ARS, CLP, MXN, UYU
<i>McNown and Wallace (1989)</i>	CLP, ARS, BRL, ILS against USD	Monthly: Different periods (from 1972 to 1986)	NER, CPI and WPI	Cointegration (CRDF, CRADF, CRDW)	Cointegration when WPI is used, not for CPI. Uses the ECM model to describe the mechanics of adjustment to the l.r. equilibrium	Fav. for CLP, ARS and BRL
<i>Taylor (1990)</i>	GBP, USD	Monthly 1921:01-1925:05	NER, WPI	Unit Root (PP, ADF, Johansen) Cointegration (CRADF, Johansen)	Includes ECM method. By mid 1926, GBP was undervalued against USD by some 2%. Some form of PPP held for the whole of the 1920s float between GBP and USD.	Fav.
<i>Ahking (1990)</i>	USD against GBP	Monthly 1921:01-1925:05	NER, WPI	Unit Root (ADF) Cointegration (Engle and Yoo, 1987)	Tests that the variables contain no deterministic components	Not Fav.

<i>McNown and Wallace (1990)</i>	GBP, CAD, JPY against USD	Monthly Different periods from 1957:03-1986:06	NER, CPI and WPI	Unit Root (DF, ADF) Cointegration (CRDF, CRADF)	Period encompasses fixed and flexible exchange rate regimes. The strongest support for cointegration comes from the fixed rate subperiod. Except for Canada (WPI), cointegration is rejected for the last period of flexible exchange rates	Fav. for fixed periods, not for flexible. Not fav. for GBP
<i>Canarella, Pollard and Lai (1990)</i>	CAD, DEM, JPY, GBP against USD	Monthly 1974:01-1987:12	NER, WPI	Cointegration (CRDF, CRADF, CRDW) Kalman Filter	Introduction of the time-varying parameter Failure on monetary exchange rate models can come from the presence of structural changes	Fav.
<i>Nachane and Chissanthaki (1991)</i>	FRF, BEF, GLR, ITL, BS, JPY, CAD against USA and DEM	Monthly 1973-1985/6 and 1979-1985/6	NER, WPI	Band-spectral analysis Cointegration (CRDF, CRADF, SB)	Using the Wu-Hausman test, verifies if omitted variables or endogeneity of relative inflation differentials cause misspecification in the model. PPP was likely to fare better with the DEM as a base. Formation of the MSE seems to have contributed to exchange rates stability for its members	Half of the cases
<i>Ngama and Sosvilla-Rivero (1991)</i>	ESP against USD and DEM	Monthly, Quarterly: 1977:01/I-1988:12/IV	NER, CPI, WPI	Unit Root (PP) Cointegration (CRADF) FMOLS	Includes Unrestricted ECM and an analysis of Granger-causality	Only Fav. For ESP/DM with WPI
<i>Phylaktis (1992)</i>	GBP, FRF, USD against GRD	Monthly 1923:01-1925:12	NER, Relative Prices	Unit Root (DF, ADF, Johansen) Cointegration (Johansen)	Speed at which long run PPP was reached following a shock was 50% per month	Fav.
<i>Bleaney (1992)</i>	CHF, GBP, ITL, CAD, FRF, JPY against USD	1900-1972 1973-1988	NER, Relative Prices	Cointegration (CRDF, CRADF) ECM Structural Breaks	Also considers RER, studying for unit roots and structural breaks due to regime changes If a heteroscedasticity correction is applied, the results are less favourable to PPP	Fav.: FRF, ITL, GBP, JPY (Not Fav: CHF, CAD) Not Fav. (except for FRF) Not Fav.
<i>Taylor (1992)</i>	USD against GBP	Monthly 1921:01-1925:05	NER, WPI	Unit Root (PP, ADF, Johansen) Cointegration (CRADF, Johansen)	Includes ECM estimation UK prices are $I(0)$, there is a deterministic trend in this series. Overvaluation of GBP when fixed	Fav.
<i>Bleaney (1993)</i>	FRF, DEM, GBP and USD against CHF	Monthly: 1921:02-1924:11 (1923:08 for Germany)	NER, WPI, Cost-of-living Index	Unit root (ADF) Cointegration (CRADF) ECM	Performs two more tests than other studies. ECM turns PPP for GBP to hold	Fav: FRF, DEM Not fav: GBP, USD

<i>Kugler and Lenz (1993)</i>	15 currencies against DEM	Monthly 1973:01-1990:11	NER, CPI	Unit Root (DF, PP) Multivariate Cointegration (Johansen)	Empirical evidence for PPP is as strong within the EMS than for the european countries outside this system. For the rejection of PPP in the case of the USD we could argue that shocks (fiscal policy shocks) could explain permanent changes of the relative prices	Fav.: GBP, ITL, NWK, ATS, PTE, ESP. Not Fav.: USD, CAD, BEF, DKK Mixed: CHF, FRF, JPY, NLG, SEK
<i>Steigerwald (1996)</i>	CAD, FRF, DEM, GBP, ITL, USD (15 pairs)	Annual 1927-1990	NER, CPI	Unit Root (ADF, PP) Cointegration (Johansen, DOLS)	Specifies a general dynamic structure, which provides evidence of PPP for the 14 of 15, instead of 8 of 15 (result obtained with unit-root tests)	Fav.
<i>Maeso (1997)</i>	19 countries against USD	Quarterly 1974:I-1994:III	NER, IPI, CPI	Unit Root (ADF, PP) Cointegration (Johansen)	Less favourable results when IPIs are used. Suggests that PPP should be studied taking into account that causality can come from both directions (NER $\Leftrightarrow$ Prices)	Fav.
<i>Papell (1997)</i>	20 developed countries (relative to DEM and USD)	Monthly, Quarterly: 1973:01/I-1994:09/III	NER, CPI	ADF and Panel Unit Root	Stronger conclusions can be made when panel is larger, for DEM (monthly) rather than USD (quarterly) data.	Fav
<i>Telatan and Kazdagli (1998)</i>	TRL against DEM, FRF, GBP and USA	Monthly 1980:10-1993:10	NER, CPI	Unit Root (ADF and PP) Cointegration (CRADF, CRPP)	Unique economic features of high inflation, structural changes, changes in taste and technology in Turkey	Not Fav.
<i>Jacobson and Nessén (1998)</i>	DEM, GBP, USD, JPY	Annual 1936-96	NER, WPI	Multivariate Cointegration (Johansen)	Use the ECM. Find three long-run, cointegrating relations but none can be interpreted in terms of PPP	Fav.to a weak form of PPP. Reject PPP
<i>Salehizadeh and Taylor (1999)</i>	27 countries (semi-advanced countries, emerging economies and developing nations)	Monthly 1975:01-1997:09	NER, CPI	Unit Root (ADF) Cointegration (Johansen, CRADF)	Simmetry and proporcionality conditions are rejected (all but one)	Fav. (14 of 27)
<i>Christev and Noorbakhsh (2000)</i>	BGL, CZK, HUF, PLN, ROL, SKK, USD, DEM and ECU	Monthly 1990:01-1998:11	NER, CPI	Unit Root (ADF, PP) Cointegration (Johansen, DOLS)	Use ECM Provide an explanation for such behaviour that is consistent with the literature on transition and foreign exchange markets	Weak evidence to support long-run equilibria
<i>Wang (2000)</i>	PHP, THB, IDR SID, MYR, JPY KRW against USD	Monthly 1979:01-1996:12 1973:01-1996:12 1980:01-1996:12	NER, CPI	Unit Root (ADF) Multivariate Cointegration (Johansen)	PPP vector does not exist in the cointegration space (conditions of symmetry and proportionality are rejected) Flexible exchange rate period	Fav.
<i>Bai and Ng (2001)</i>	21 countries (against USD)	Quarterly: 1974:I-1997:IV	RER, NER, CPI	Panel Unit Root (KPSS, MSB)	Model strong cross-section correlation via a factor model	Not Fav.

<i>Pedroni(2001)</i>	GBP, BEF, DKK, FRF, ITL, DEM, NLG, SEK, CHF, CAD, JPY, GRD, PTE, ESP, TRL, NZD, CLP, MXN, INR, KRW against USD	Monthly: 1973:06 1993:11	NER, CPI	Panel cointegration (within-dimension and between-dimension panel FMOLS and DOLS tests)	The approach allows to pose the null hypothesis so that he can test if strong PPP holds consistently for all countries in the panel or not.	Not Fav.
<i>Nagayasu (2002)</i>	17 African countries against USD	Annual: 1980-1994	NER, Relative Prices	Unit Root (individual: ADF and Panel: IPS) Panel Cointegration (FMOLS)	Uses parallel market rather than official exchange rates (the last are fixed against a single currency or a basket of other currencies) Any significant discrepancy between official and parallel exchange rates may serve as a warning sign that official rates are misaligned	Fav. (weak form)
<i>Xu (2003)</i>	CAD, FRF, DEM, ITL, JPY, KRW, NGL, GBP, USD	Quarterly: 1974:I-1997:IV	NER, RER, WPI, CPI, TPI	Unit root (ADF, SBC)	Cointegration for RER Prediction of NER, using PPP relationship. The I.r.PPP is rejected with the restrictions (symmetry and proportionality) imposed a priori on the exchange rate and price data, but unequivocally supported in their absence.	Mixed
<i>Cerrato and Sarantis (2003)</i>	20 developing countries against the USD	Monthly: 1973:01 1993:12	NER, CPI	Panel unit root (Individual: ADF, panel: IPS) Panel cointegration (DOLS, DGLS, McCoskey and Kao, Larsson)	Results imply the absence of persistently over-valued or under-valued black market exchange rates (if this were not the case, it would cause some damaging effects on economic growth and the allocation of resources)	Fav.
<i>Basher and Mohsin (2003)</i>	10 Asian developing countries against the USD (INR, IDR, KRW, MYR, NPR, PKR, PHP, SGD, LKR, THB)	Monthly, Quarterly: 1980:01/I-1999:12/IV	NER, CPI	Unit Root (individual: ADF; panel: LL and IPS) Cointegration (individual: Johansen-Juselius; panel: FMOLS, DOLS)	The between-dimension estimators consistently produce larger estimates than the within-dimension estimators. Analysis of individual countries indicate that the failure of the PPP is not driven by the data from only a few countries. Empirical findings do not support neither the relative nor the absolute versions of PPP.	Not Fav.

4.2. Reviewing the empirical evidence on PPP

It is important to note that most papers attempt to test the relative version of PPP [equation (2)] following two different paths: some authors [Frenkel (1981), Jacobson and Nessen (1998) or Basher and Mohsin (2003)] test both the absolute (or strong) version of PPP [equation (1)] and the relative version [equation 2]. Some other papers directly test the relative version of PPP. From a theoretical point of view, both versions are suitable for being tested, but traditional literature has pointed out that almost always it is not possible to accept the accomplishment of the strong version. Therefore, the wide majority of papers only consider the relative version of PPP.

While there is a large coincidence on the fact that absolute PPP is almost always rejected, the evidence on relative PPP is mixed. In order to make clear what this inconclusive result means, it is necessary to refer to specific countries, periods and econometric methodologies and proceedings.

Pre-cointegration methodologies [Krugman (1978), Frenkel (1981)] tend to support PPP when instrumental variables are used and USD is excluded of the analysis, which would suggest an effect of the economic trading environment on the convergence of relative prices.

Time-series cointegration includes the largest part of reviewed papers. Indeed, PPP testing would seem to provide the perfect context for applying cointegration methods. There are some studies in which the area of interest is Europe [Edison and Klovland (1987)], but in most of them the base currency is the US Dollar [Ngama and Sosvilla-Rivero (1991), Taylor and McMahon (1988) Taylor (1990), among others], or the US Dollar, the Canadian Dollar or the Japan Yen [see, e. g., McNown (1990), Canarella (1990), Taylor (1988), Steigerwald (1996) or Phylatkis (1992)]. Both the pre-cointegration and the time-series cointegration stages of testing PPP combine different price indices (CPI, WPI, MPI, etc.). The authors have also tried to explore if there is a positive relation between the acceptance (rejection) of PPP and the selected price index. This would be in line with the theoretical framework that suggests a higher probability of rejection in the presence of important differences in the basket of goods included in the CPI, as well as the presence of non tradable goods in them. It would seem that the use of WPI would help to correct this problem, and in general, the papers have made an extensive use of the WPI as a measure of prices in the different countries. When GB Pound is considered, specially when using data of the twenties, the results indicate inconclusive evidence, but in general, there is a supportive evidence

on the accomplishment of PPP. Mikkelsen (1989) and McNown et al (1989) extend the area of study to Latin America using recent time series data and obtaining evidence in favour of the PPP hypothesis.

Panel cointegration is the newest econometric technique in testing PPP. Basher and Mohsin (2003), Cerrato (2003) or Xu (2003) have made use of this technique in order to overcome some of the limitations of time-series cointegration. The wider possibilities of the new methodology and the long tradition in testing PPP, encouraged the economists⁴ to increase the areas of analysis, and therefore they provide some evidence for Africa [Nagayasu (2002)], Asia [Wang (2000) or Basher *et al.* (2003)] or Oceania [Pedroni (2001)]. This last paper only provides evidence for the validity of strong version of PPP, and the author concludes rejecting the absolute version of PPP in countries of Europe, Asia, Oceania and America (using data from 1973 to 1993). Bai and Ng (2001), using approximately the same period and test PPP for twenty-one countries, find evidence against PPP. Basher and Mohsin (2003) extend their analysis from 1980 to 1999, considering some Asian countries, and obtain the same result as Pedroni (2001). The other papers tend to find a more supportive evidence on the empirical accomplishment of PPP, which is at least, mixed in favour of PPP⁵, or unqualifiedly favourable to PPP.

⁴ They coincide in using the Consumer Price Index in their analysis.

⁵ For example, Xu (2003) rejects long run PPP when symmetry and proportionality restrictions are imposed, but supports PPP when these restrictions are absent a priori.

5. FORECASTING WITH PPP

*“Prediction is very difficult,
especially about the future”
Niels Bohr*

Despite the paramount modelling effort registered in the last two decades, it is widely recognised that exchange rates are extremely difficult to forecast.

The pessimism about the forecasting ability of exchange rate models has been generally accepted after the publication of the influential paper by Messe and Rogoff (1983). These authors performed a large number of statistical tests, indicating the superiority of the linear out-of-sample forecast of exchange obtained through a simple random walk model compared with the forecast based exchange-rate determination models that used a wider set of economic variables as regressors. This superiority is also clear when the forecast is determined *ex post* (i. e., using real historical values of the explicative variables in the regression). Recently, Cheung, Chinn and García-Pascual (2002) have evaluated the predictability of a wide variety of models that have been proposed over the past decade, and they conclude that these models are still unable to improve a random walk forecast.

Several explanations for the failure of structural models have been suggested, including misspecification of the models and poor modelling of expectations [Frankel and Rose (1995) for a survey]. Recently, interest has been shown in the possibility that non-linearities account for the apparent unpredictability of exchange rates, and some papers have highlighted the importance of non-linear adjustment of the exchange rate to the value implied by fundamentals, including Taylor and Peel (2000) and Clarida *et al.* (2003).

Based on this non-linearities some authors have explored the non-parametric, nearest neighbour forecasting technique (see Fernández-Rodríguez *et al.*, 2003 for a survey). The basic idea behind these predictors, inspired in the literature on forecasting in non-linear dynamical systems, is that pieces of time series sometime in the past might have a resemblance to pieces in the future. In order to generate predictions, similar patterns of behaviour are located in terms of nearest neighbours. The time evolution of these nearest neighbours is exploited to yield the desired prediction. Therefore, the procedure only uses information local to the points to be predicted and does not try to fit a function to the whole time series at once [see, e. g. Fernández-Rodríguez, Sosvilla-Rivero and Andrada-Félix (1997)].

Regarding PPP, Cochran and Defina (1995) examine the usefulness of PPP as a guide to future exchange rate movements. To that end, they use duration analysis to investigate whether deviations from PPP exhibit positive dependence (i.e. as time passes the probability that an exchange rate will return to its PPP level after a deviation occurs). They use monthly data covering January 1974 to December 1992 to compute PPP series for eighteen currencies (all *vis-à-vis* the US dollar). Their results suggest that PPP can not help in forecasting future exchange rate changes, a result consistent with market efficiency.

From 1986, *The Economist* has been published the Big Mac index, based upon the PPP hypothesis. Their "basket" is a McDonald's Big Mac, produced locally to roughly the same recipe in 120 countries. The Big Mac PPP is the exchange rate that would leave hamburgers costing the same in America as abroad. Comparing actual rates with PPPs signals whether a currency is under-or overvalued. Although the Big Mac index is not a perfect measure of PPP (since hamburgers cannot be traded across borders, prices may be distorted by taxes, different profit margins or differences in the cost of non-tradable goods and services, such as rents), several studies have found that the Big Mac PPP is a useful predictor of future movements. In this sense, when Europe's new currency was launched in January 1999, the Big Mac index suggested that the euro was already overvalued at its launch, when virtually everybody predicted that it would rise against the dollar. Moreover, Ong (2003) finds that the Big Mac index has been surprisingly accurate in tracking exchange rates in the long term, although there are some persistent deviations from PPP (in particular emerging-market currencies are consistently undervalued).

On the other hand, Kilian and Taylor (2003) explore whether a nonlinear, exponential smooth transition autoregressive (ESTAR) model based on PPP may help in beating the random walk forecast of exchange rates. They develop a bootstrap test of the random walk hypothesis of the nominal exchange rate, given ESTAR real exchange rate dynamics. Using quarterly data for seven OECD countries covering the 1973.I-1998.IV period, they do find strong evidence of predictability at horizons two to three years, but not at shorter horizons.

Finally, Sosvilla-Rivero and García (2003) assess the empirical relevance of an expectations version of PPP in forecasting the Dollar/Euro exchange rate, based on the differential of inflation expectations derived from inflation-indexed bonds for the Euro area and the USA. To that end, they use daily data covering the 16 September 1998-31 December period for the Dollar/Euro exchange rate and for the inflation expectations. Regarding the inflation expectations, both for the Euro area

and for the United States, they were calculated as the break-even inflation rates using the information provided by ten-year public bond yields, therefore allowing instantaneous processing of all of the information that exchange-rate operators receive concerning the behaviour of prices in the economies under study. The results in Sosvilla-Rivero and García (2003) suggest that, with few exceptions, the PPP-based predictors behave significantly better than a random walk in forecasts up to five days, both in terms of prediction errors and in directional forecast.

6. CONCLUDING REMARKS

*“I’ll give you a definitive maybe”
Samuel Goldwyn*

The main feature of the PPP hypothesis is that, in the long run, the exchange rate between two currencies should move towards the rate that would equalise the prices of an identical basket of goods and services in each country. This hypothesis is one of the oldest and controversial one in economics. As Officer (1982) has pointed out, over the centuries, the PPP hypothesis has been discovered, fallen into disuse, and been rediscovered (this pattern repeated several times). As stated by Samuelson (1964, p. 149), it seems that “[e]ach generation must rekill its phoenixes”. Each episode of support for the hypothesis has been associated with a traumatic development: inflation following price stability and/or a change from a fixed-exchange rate regime to a flexible-exchange rate one. Empirical research on PPP enjoyed a rebirth after the move to flexible exchange rates in the early 1970s, although this re-emergence does not seem to have led to any consensus as to its general empirical validity.

We have reviewed a selection of empirical papers in order to explore the ongoing discussion on PPP. Given the close relation between successive developments in econometric techniques and the evolution of such empirical research, we have also surveyed the econometric methodology used in testing the validity of PPP during the last decades.

Even though we have seen that there are good reasons why PPP should not be expected to hold (the existence of transportation costs, tariffs and other legal barriers to commerce, non tradable goods, different productivity shocks and preferences for goods in different countries and differences in price indices), PPP remains an essential element of open economic macroeconomics. Its simplicity and intuitive appeal, its concreteness (the ingredient of the hypothesis are minimal and basic), and its usefulness (whether to know to what extent it is valid or used to measure deviations from PPP) lead us to conclude with Houthakker (1978, p. 71) that “the complete rejection of PPP is as mistaken as its complete acceptance”. Indeed, the results from the empirical evidence revised in this paper are not conclusive, although many economist would recognised PPP as an important empirical possibility in the long run as Keynes (1923, p. 79).

The most interesting feature in the empirical effort to test the validity of PPP is that this paramount endeavour has contributed to a great development of many

econometric techniques. Indeed, empirical testing of PPP has been continuously changing as these technical innovations were appearing to an extent that we can point out PPP testing as a main stimulating factor behind the growth of time series econometrics. In this process, the increasing availability of data sets covering long periods for the main currencies has played a major role, mitigating the problems associated with small samples. In this sense, the panel cointegration approach, exploiting both cross-section as well as time series variation in the data is called to become an important tool in PPP testing in the years to come as the development of new data sets will allow researchers to investigate both longer and more disaggregated time series. In view of the mildly encouraging results from this latter approach, some optimism about the benefits from implementing new extensions in this area seems justified.

Finally, recent investigations have indicated some evidence of predictability using PPP both at long (two to three years) and short horizons (up to five days), opening new avenues that seem worthy of further research.

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