ΠΑΝΕΠΙΣΤΗΜΙΟ ΠΕΙΡΑΙΩΣ

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"A NEW APPROACH

TO

PURCHASING POWER PARITY "

ΤΡΙΜΕΛΗΣ
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ΠΕΙΡΑΙΑΣ, ΙΟΥΛΙΟΣ 2004

CONTENTS

1. INTRODUCTION

The fundamental notion of the Purchasing Power Parity (PPP) hypothesis is that the exchange rate depends on relative prices and will adjust to reflect changes in the relative price levels of the two countries. In fact, tests for PPP have evolved along with time series analysis as much advancement in time series techniques have been applied in attempts to uncover parity reversion in real exchange rates. Testing for mean-reversion in the real exchange rate is important for many reasons. It is an important constituent of most models of exchange rate determination, being regarded as a long-run equilibrium or an arbitrage condition in goods and assets markets. The real exchange rate dynamics implied in models of inter-temporal smoothing of traded goods consumption (Rogoff, 1992) and cross-country wealth redistribution/transfers (Obstfeld and Rogoff, 1995) makes the PPP hypothesis a meaningful one to examine. The concept is also of interest to policy makers as it serves as a benchmark for computing an equilibirum exchange rate and assessing whether shocks to the real exchange rate dampen over time.

Given its importance in open economy macroeconomic models and for constructing fundamental equilibrium exchange rates, the long-run PPP relationship has been subjected to extensive empirical investigation during the last decade. Especially, the behaviour of real exchange rates over the post-1973 floating exchange rate period is one of the most extensively studied empirical topics in international economics. The most frequently asked question is whether long-run purchasing power parity holds over this period, posed statistically as whether unit roots in real exchange rates can be rejected. Despite the large amount of research on this topic, we do not have a definitive answer to the most basic issues.

The purpose of this study is to evaluate, and hopefully extend, the evidence of PPP from the application of both univariate and panel methods. While the issue of how much evidence against the unit root hypothesis in real exchange rates has been found by these tests remains controversial, there appears to be an emerging consensus that the evidence of long-run PPP is stronger now than it was a decade ago and that as time passes and more data becomes available, the evidence of PPP is strengthening.. The question is if this is purely statistical, caused by an increase in the power of unit root tests from the use of additional data and methods, or if economic factors enter into the explanation.

We investigate this question by conducting both univariate and panel unit root tests on real exchange rates for the period 1973-2003. The tests use quarterly and monthly data for the 21 countries classified from the IMF as industrialised, with the United States dollar as base currencies. For the univariate tests we first use data from 1973-1988 and then add one by one the data for the precedending years in order to see if by increasing the span of the data the univariate test provides more strengthened evidence of PPP. For the panel tests we construct 7 different panels and conduct the panel tests proposed by Levin and Lin (1993), Im, Pesaran and Shin (1997) and Pesaran (2003).

The study is organised as follows. Section 2 discusses the basic concepts of the PPP hypothesis, the absolute and relative PPP versions, the interpretations of PPP and the price index choice in a PPP calculation and the factors responsible for deviations from PPP hypothesis. Section 3 provides a review of the empirical literature on PPP. In Section 4, we focus on the recent econometric developments in the area of testing the PPP hypothesis. In section 5 we provide an up-to-date survey of all the econometric methodologies used in the empirical evaluation of the PPP and conduct our own econometric research. Section 6 gives information about the data and the panels used in our empirical research. Finally, Section 6 provides the concluding remarks.

2. THE NOTION PURCHASING POWER PARITY.

2.1. THE HISTORY OF PPP

The concept of purchasing power parity (PPP) has long played a prominent role in theoretical and empirical research, though the usage of the "term purchasing power parity" (PPP) is relatively recent and was introduced by Cassel in 1918. The concept itself, and discussions relating to the relationship between the exchange rate and prices more generally have a very long history in the economics. 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 Oficer (1982)]. These theologists 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 a PPP point of view.

After this period, the PPP doctrine appeared to be more or less forgotten both in academic and policy cycles until the First World War, where severe episodes of hyperinflation and dislocated exchange rates stimulated a renewed interest in the PPP. Cassel was 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 it was used as a tool of analysis in the in the international policy debate concerning the appropriate level for nominal exchange rates among the major industrialised countries.

Closer to our own time, the change to flexible exchange rates in the early 1970s provoked yet another intellectual upturn to the PPP hypothesis. Since the abandonment of Bretton-Woods Systems in 1971, the EU tried several alternatives (the "snake" and the European Monetary System) before reaching its current shape. Nowadays, the efforts towards monetary and economic stabilisation under the framework of the European Union (EU) have highlighted the significance of inflation and exchange rates in the success of this effort. By the Maastricht Treaty, membership in the Euro required the achievement of five criteria, including inflation convergence and nominal exchange rate convergence within its member states.

The purchasing power parity (PPP) hypothesis shows the relation between nominal exchange rate and price ratio, which implies that real exchange rate is constant over the time. So in our days, the requirement of inflation convergence and nominal exchange rate convergence in the Euro zone, as stated above, raises new interest in the study of the PPP hypothesis.

Moreover, despite its long intellectual history and the fact that some form of PPP appears to be embedded in the way many economists think about the world, the empirical evidence on PPP is very mixed. The aim of this work is to give the definitions and interpretation of the notion "Purchasing Power Parity", review the empirical work on real exchange rates and PPP over the past years with an emphasis on recent developments and test the validity of the PPP hypothesis for the floating period.

2.2. THE ABSOLUTE AND RELATIVE PPP VERSIONS

As articulated by Cassel (1918), the absolute version of PPP postulates that the relative prices (in different currencies and locations) of a common basket of goods will be equalized when quoted in the same currency. The relative version of PPP, emphasizing arbitrage across time rather than across space, is that the exchange rate will adjust to offset inflation differentials between countries. While Cassel understood the possibility that the exchange rate might transitorily diverge from PPP, he viewed the deviations as minor. Modern versions of PPP, recognizing the importance of slow speeds of adjustment, define PPP as reversion of the real exchange rate to a constant mean.

More specifically, in its ***absolute*** version, PPP theory gives the relationship between the exchange rate, S, (expressed as the home currency price of a unit of foreign exchange), and the ratio of domestic and foreign prices (P and P* respectively), as follows:

$$S_t = P_t / P_t^*$$

According to that, the nominal exchange rate should be such that the purchasing power of a unit of currency is exactly the same in the foreign economy as in the domestic economy, once it is converted into foreign currency in that rate.

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 holds when the percentage rate of change of the exchange rate will equal the differential between the percentage rates for change of price levels across the two countries. That is:

$$\Delta s_t = \Delta p_t - \Delta p_t^*, \qquad \text{where } \Delta s_t = \log (S_t / S_{t-1}) \qquad (2)$$

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 associate with a domestic currency appreciation.

Relative PPP is less strict than absolute PPP. It allows domestic and foreign prices (expressed in domestic currency) to differ from each other but still 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.

2.3. PPP APPROACHES – THE PRICE INDEX ISSUE

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 a homogeneous commodity i produced both at home and abroad.

Let P_i and P_i* are the prices of that commodity at home an abroad, stated in home and foreign currency, respectively, and S the exchange rate. Then ignoring 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.:

$$P_i = P_i{*}S \qquad (3)$$

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 that is more expensive. The rise of demand in the country that the product is cheaper would provoke rise in its price and respectively the rise in supply in the other country would reduce the product's price. All this procedure 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 $P = g(P_1,P_2,...,P_n)$ and $P^* = g(P_1^*,P_2^*,...,P_n^*)$. Then by using identical weights in constructing each country's price level, we obtain

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

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,

$$\iota = r + \Delta p^e \tag{5a}$$

$$\iota^* = r^* + \Delta p^{*e} \tag{5b}$$

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 ($\Delta p^e = \log P^e_{t+1} - \log P_t = p^e_{t+1} - p_t$), 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 ($\Delta s^e = \log S^e_{t+1} - \log S_t = s^e_{t+1} - s_t$).

$$\Delta s^e = i - i^* \tag{6}$$

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:

$$r = r^* \tag{7}$$

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

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

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

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.

The formulation of the PPP theory 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". Four alternative price indices have traditionally been generally used:

- consumer price indices (CPIs),
- gross domestic product (GDP) deflators,
- wage rate indices (WRIs), and
- wholesale price indices (WPIs).

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.

2.4. DEVIATIONS FROM PPP

Persistent deviations from PPP can cause not only macroeconomic disequilibrium, but also resource misallocation and income redistribution.

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, another reason that can explain short-run deviation from PPP is the different speed of adjustment between exchange rates and prices. Exchange rates and commodity prices are determined in different kinds of markets. The exchange rate, as an asset price, reacts rather quickly to “news”, while prices of goods reflect new information relatively slowly. This can cause deviations from PPP (Daniel, 1986; Frenkel, 1981). 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.

Finally, factors such as the existence of price contracts and/or rationing, changes in the structure of relative prices in the domestic and foreign economies lead to divergence from PPP. Official intervention in the foreign exchange markets have a great impact on short-run exchange rate movement, the nominal exchange rate moves further away from its equilibrium level and provokes deviation from PPP. Monopolistic and oligopolistic forces, product differentiation, trading restrictions (e.g. tariffs and 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. REVIEWING THE LITERATURE ON PPP

There is a vast empirical literature testing the validity of the PPP hypothesis. We will try to review this literature and explain the conflicting results that have occurred, classifying the studies from the point of view of the econometric methods used in the empirical application.

Before this it is important to note that most papers attempt to test the relative version of PPP 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 and the relative version. 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.

Firstly, we refer to some representative studies from the seventies, when the econometric methodologies applied in these papers (OLS, Instrumental Variables) did not consider the statistical properties of time series. ***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. The most commonly used formal tests for long-run PPP consist of testing whether the real exchange rate has a unit root. If the unit root hypothesis can be rejected, there is evidence of mean reversion of the real exchange, or equivalently, long-run PPP. Much of the literature has found a random walk in real exchange rates and failed to support PPP for the floating exchange rate period (Frankel, 1981; Cumby and Obstfeld, 1984; Enders, 1988).

A prominent explanation for the failure of PPP maintained by Balassa (1964) and Samuelson (1964) emphasizes that empirically, when all countries' price level are translated to dollars at prevailing nominal exchange rates, rich countries tend to have higher price levels than poor countries. The reason for this phenomenon, they conjured, is not simply that rich countries have higher absolute productivity levels than poor countries, but because rich countries are relatively more productive in the traded goods sector. Non traded goods tend to be more service intensive, so there is less room for technological superiority. Productivity differentials between traded and non-traded sectors lead to movements in the relative prices of non-tradeables and real exchange rates. Hsieh (1982) and Asea and Mendoza (1994) using a neoclassical general equilibrium model empirically supported these propositions. However, Engel (1996) and Rogers and Jenkins (1995) find that the relative price of non-tradeables has virtually no effect on US real exchange rate movements.

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.

The flurry of empirical studies employing this methodology were unanimous in their failure to reject the unit root hypothesis (e.g. Enders, 1988; Taylor 1988; Mark, 1990). Whether one uses conventional Augmented-Dickey-Fuller (ADF) tests or variety of more powerful alternatives, the unit root null is rejected only occasionally. This created something of a problem for international macroeconomists, under the skin of whom, according to Dornbush and Krugman (1976) lay “a deep belief in some variant of the PPP theory of the exchange rate”. The collapse of the short-run PPP had been cheerfully accepted and replaced with a belief in long-run PPP. Now the data appeared to be telling that even long-run PPP was a chimera. This led some economists to posit models to explain why the real exchange rate could in fact be non-mean reverting (e.g. Stockman, 1987; Taylor, 1995).

Frankel (1986, 1990) was the first economist to point out the tests typically employed to examine long-run PPP stability of the real exchange rate, if based on data covering just 15 years or so, may have low power to reject the null hypothesis even if it is indeed false. The argument is that that if the real exchange rate is in fact stable in the sense that it tends to revert towards its mean over long periods of time, then examining a relative short period may not yield enough information to be able to detect slow mean reversion towards PPP. This point was further taken up and examined by a number of authors (e.g. Lothian, 1986, 1998a; Froot and Rogoff, 1995; Lothian and Taylor, 1996, 1997) who gave a solution to the problem by utilizing longer span of data. In some of these studies the unit root hypothesis was rejected with a century (or more) real exchange rate data. However, these studies combine data from fixed and floating exchange rate periods, and do not provide direct evidence on PPP under flexible exchange rates. As demonstrated by Shiller and Perron (1985) researchers cannot overcome the problem by increasing the frequency of observation- say from annually to quarterly or monthly- and increase the number of data this way. We need a long span of data in term of years covered in order to increase the power of the tests.

All the work discussed so far considers countries individually. An alternative approach involves using a panel of data; i.e. data on more than one exchange rate. This way, by increasing the number of real exchange rates under consideration the amount of information employed in the tests is increased and consequently the power of the tests raises. Recently a number of researchers have turned to panel data methods in an attempt to find more evidence of long-run PPP in current floating exchange rate data. This has been in part inspired by the work of Levin and Lin (1992) , who showed that where there is not enough time series variation to produce good power in unit root tests, an alternative small amount of cross-section variation can result in substantial improvement.

Abuaf and Jorian (1990) examine a system of ten first order autoregressive rejections for real dollar exchange rates over the period of 1973-1987 where the first order autocorrelation coefficient is constraint to be the same in every case and takin account an contemporaneous correlation among the disturbances, their results show a marginal rejection of the null hypothesis of joint non mean reversion and they interpret this as evidence in favour of the PPP.

This study stimulated a whole strand of literature in which researchers employ various multivariate generalizations of unit root tests in order to increase the test power (e.g. Flood and Taylor, 1996; Wu, 1996; Frankel and Rose, 1996; Coakley and Fuertes, 1997; Lothian, 1994; O'Connell, 1998; Papell, 1998). Although the recent float seamed to be in favour of panel data methods, the empirical evidence has also been mixed. While Lothian(1994); Frankel and Rose(1996) find evidence in favour of mean reversion, Hakkio(1984) is unable to reject the random walk hypothesis with a sample of four exchange rates against the dollar. Jorian and Sweeney (1996) update the Abuaf and Jorian(1990) and find evidence for real exchange rate stationarity. Using a GLS approach, O'Connell (1998) demonstrates that cross correlations between economies have a significant impact on the results and leads to rejection of PPP. Coakley and Fuertes (1997) adopt the IPS test and support real exchange rate stationarity; whereas, Canzoneri et al. (1999) find nonstationarity for real exchange rates. Recently, Lothian (1998), Siddique and Sweeney (1998) and Koedijk et al. (1998) also support mean reversion using panel data.

The wider possibilities of the new methodology and the long tradition in testing PPP, encouraged the economists 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 andAmerica (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 PPP, or unqualifiedly favourable to PPP.

However, in their majority researchers in the above studies coincide in using the Consumer Price Index in their analysis and construct real exchange rates from bilateral dollar exchange rates and CPI, with the United States as the 'foreign' country.

- Empirical evidence on PPP

In this section we provide a detailed report of the sydies that have been performed during the long history of PPP. Firstly, we start from the most recent papers that test the PPP hypothesis using the newest econometric methodology: panel method. Secondly, we examine the contributions of some of the authors who applied methods for dealing with nonstationary time series, using the concept of cointegration. Finally, 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.

We also provide two supplementary tables that will help in the interpretation of the report that follows. 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 author of a particular paper. The second specifies the period (or subperiods) under investigation, as well as the characteristics of the sample (monthly, annual or quarterly), and span. The third contains the currencies examined.. The fourth column shows the variables that are used: “NER” stands for “nominal exhange rate”, ‘‘RER’’ gives information about “Real Exchange Rates”, “CPI” means “consumer price index”, “WPI” is the “wholesale price index”, “GDPD” is the “deflactor 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

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

Table 2: Econometric Methodologies and Tests Abbreviations

Abbr.	***Econometric Method/Test***
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

AUTHOR	PERIOD	DATA	VARIABLES	ECONOMETRIC METHOD	CHARACTERISTICS	RESULT
Basher, Mohsin (2003)	Monthly and Quarterly 1980:01/I 1999:12/ IV	10 Asian developing countries against the USD (INR, IDR, KRW, MYR, NPR, PKR, PHP, SGD, LKR,THB	NER, CPI	Unit Root (Individual:ADF/ Panel: LL and IPS t-bar Cointegration (Individual: Johansen & Juselius (1990)/ Panel: Pedroni (1999)	Failure of PPP in flexible exchange rate regime. Empirical findings do not support neither the relative nor the absolute versions of PPP.	Not Fav.
Cerrato and Sarantis (2003)	Monthy: 1973:01 - 1993:12	20 developing countries against USD (black market exchange rates)	NER, CPI	Panel Unit Root (Individual:ADF/ Panel: IPS Cointegration (DOLS, DGLS, McCoskey and Kao, Larsson)	Absence of persintently over-valued or under-valued black market exchange rates which whould cause damaging effects on economic growth and allocation of resources	Fav.
Lopez Claude (2003)	1973:I - 2001:IV (for Greece 1973:I - 1999:IV)	AUD, ATS, BEF, CAD, DKK, FIM, FRF, DEM, GRD, IEP, ITL, JPY, NLG, NZD, NOK, PTE, ESP, SEK, CHF, UK, USA	NER, CPI	Univariate Estimations of RER/ ADF & DF-GLS tests (using the lag selection the GS and the MAIC procedures). Mulivariate estimation/ ADF-SUR (a version of the LLC test accounting for contemporaneou s correlation) & the DF-SUR-GLS tests.	Combines univariate-multivariate unit root tests, creats a more powerful panel test. Extends the GLS-detrending procedure of Elliot, Rothenberg, and Stock to a panel ADF test using LL hypotheses. The new DF-GLS-SUR test, focuses on the stationarity of the entire panel.	Fav.
Xu (2003)	Quarterly : 1974:I - 1996:IV	DEM, GBP, ITL, CAD, FRF, JPY, KRW,NLG, USD	NER, RER, WPI, CPI, TPI	Unit Root (ADF,SBC)	Cointegration for RER Prediction of NER, using PPP relationship. The 1.r.PPP is rejected with the restrictions (symmetry and proporcionality) imposed a priori on the exchanche rate and price data, but unequivocally supported in their absence.	Mixed

Nagayasu (2002)	Annual: 1980-1994	17 African countries against USD	NER, Relative Prices	Unit Root (Individual:ADF/ Panel: IPS Panel Cointegration (FMOLS)	Uses parallel market rather than official exchange rates (fixed against a single currency or a basquet of other currencies). Discrepancies between official and parallel exchange rates may serve as a warning that official rates are misaligned	Fav. (weak form)
Papell (2001)	Quarterly 1973:I - 1996:IV	AUD, ATS, BEF, CAD, DKK, FIM, FRF, DEM, GRD, IEP, ITL, JPY, NLG, NZD, NOK, PTE, ESP, SEK, CHF, UK against USD	RER, CPI	Univariate Tests of the RERs through ADF test. Panel extention of the univariate ADF test, which accounts for both a heterogeneous intercept and serial correlation.	Univariate and Panel Methods have not produced strong rejections of unit roots. The hypothesis that these non-rejections are due to the episode of large appreciation and depreciation of the USD in the 1980s is tested. With Panel Methods, strongly reject the unit root null for countries that adhere to the typical pattern of the USD's rise and fall.	Mixed
Pedroni (2001)	Monthly: 1973:06 - 1993:11	GBP. BEF, DKK, FRF, ITL, DEM, NLG, SEK, CHF, CAD, JPY, GRD, PTE, ESP, TRL, NZD, CLP, MXN, INR, KRW USD	NER, CPI	Panel cointegration (within-dimension and betweendimensio n 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.
Bai and Ng (2001)	Quarterly 1974:I - 1997:IV	21countries against USD	RER, NER, CPI	Panel Unit Root (KPSS, MSB)	Model strong cross-section correlation via a factor model	Not Fav.
Christev and Norbakhsh (2000)	Monthy: 1990:01 - 1998:11	BGL, CZK, HUF, PLN, ROL, SKK, USD, DEM and ECU	NER, CPI	Unit Root (ADF, PP) Cointegration (Johansen, DOLS)	Use of ECM provides explanation for behaviour that is consistent with the literature on transition and foreign exchange markets	Weak evidence to support long-run equilibria
Wang (2000)	Monthly 1979:01-1996:12, 1973:01-1996:12, 1980:01-1996:12	PHP, THB, IDR SID, MYR, JPY KRW against USD	NER, CPI	Unit Root (ADF) Multivariate Cointegration (Johansen)	PPP vector does not exist in the cointegration space (conditions of symmetry and proportionality are rejected).	Fav.

Fleissing, Strauss (2000)	Quarterly : 1974:1 - 1996:3	19 OECD countries against USD	CPI, Good, Rent, Fuel, Food, Ser, NT	Panel Unit Root Tests (LL, IPS, Maddala and Wu), GLS	Generally supports PPP, the speed of adjustment differs considerably between the six price indices and four different test procedures that are used. The degree of contemporaneous and serial correlation as well as heterogeneity of the series affect stationarity and the speed of mean reversion	Fav.
Culver, Papell (1999)	Quarterly : 1973 - 1996	21 Industrialised counties	NER, CPI	Unit Root (Individual:ADF/ Panel: KPSS) Cointegration: (Engle and Granger, Shin tests)	Investigates PPP with data from the current floating exchange rate period by using tests where stationarity and cointegration are the null, rather than the alternative hypotheses.	Fav.
Salehizadeh and Taylor (1999)	Monthly 1975:01-1997:09	27 countries (semi-advanced countries, emerging economies, developing nations)	NER, CPI	Unit Root (ADF) Cointegration (Johansen, GRADF)	Simmetry and proporcionality conditions are rejected (all but one)	Fav. (14 of 27)
Jacobson and Nessén (1998)	Annual 1936 - 1996	DEM, GBP, USD, JPY	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
Telatan and Kazdagli (1998)	Monthly 1980:10-1993:10	TRL against DEM, FRF, GBP and USA	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.
Papell, Theodoridis (1998)	Quarterly 1973-82, 1973-96	21 Industrialised countries against USD and DEM	NER, CPI	Univariate Unit Root (ADF, PP) Panel Unit Root	As the sample is extended (fist 1973-82, next1973-96) the evidence of PPP strengthens but only with panel methods. Because of the sharp rise and fall of the dollar in the early80s, rejections of the unit root null don't increase as much as the sample size	Fav. Evidence of PPP stronger with DEM as base currency and with panel method

Maeso (1997)	Quarterly 1974:I - 1994:III	19 countries against USD	NER, CPI, IPI	Unit Root (ADF and 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↔Prices)	Fav.
Papell (1997)	Monthly, Quarterly : 1973:01/I - 1994:09/III	20 developed countries (relative to DEM and USD)	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.
Steigerwald (1996)	Annual 1927-1990	CAD, FRF, DEM, GBP, ITL, USD (15 pairs)	NER, CPI	Unit Root (ADF and 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.
Kugler and Lenz (1993)	Monthly 1973:01-1990:11	15 currencies against DEM	NER, CPI	Unit Root (DF and 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
Blaeaney (1992)	1900-1972 1973-1988	CHF, GBP, ITL, CAD, FRF, JPY against USD	NER, Relative Prices	1)Cointegration (GRDF, GRADF) 2)ECM 3)Structural Breaks	If a heteroscedasticity correction is applied, the results are less favourable to PPP	1)Fav.: FRF,ITL, GBP, JPY (Not Fav:CHF, CAD) 2)Not Fav. (except for FRF) 3)Not Fav.

Taylor (1992)	1921:01-1925:05	USD against GBP	NER, WPI	Unit Root (PP, ADF, Johansen) Cointegration(GR ADF, Johansen)	Includes ECM estimation UK prices are I(0). There is a deterministic trend in this serie. Overervaluation of GBP when fixed.	Fav.
Nachane and Chissanthaki (1991)	Monthly 1973-1985/6 and 1979-1985/6	FRF, BEF, GLR, ITL, BS, JPY, CAD against USA and DEM	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
Ngama and Sosvilla-Rivero (1991)	Monthly, quarterly: 1977:01/I 1988:12/IV	ESP against USD and DEM	NER, CPI and WPI	Unit Root (PP) Cointegration (CRADF) FMOLS	Includes unrestricted ECM method and an analysis of Granger-causality.	Only Fav. For ESP/DM with WPI
Ahking (1990)	Monthly 1921:01-1925:05	USD against GBP	NER, WPI	Unit Root (ADF) Cointegration (Engel and Woo, 1987))	Tests that the variables contain no deterministic components	Not Fav.
Canarella, Pollard and Lai (1990)	Monthly 1974:01-1987:12	CAD, DEM, JPY, GBP against USD	NER, WPI	Cointegration (CRDF, CRADF, CRDW) Kalman Filter	Introduction of the timevarying parameter Failure on monetary exchange rate models that come from the presence of structural changes	Fav.
Mc Nown and Wallace (1990)	Monthly: different periods (from 1957:03 to 1986:06)	GBP,CAD, JPY against USD	NER, CPI and WPI	Unit Root (DF, ADF) Cointegration (GRDF, GRADF)	Period of fixed and flexible exchange rate regimes. The stronger support for cointegrtion comes from the fixed rate subperiod. Except for Canada cointegration is rejected for flexible exchange rates	Fav. For fixed, not for flexible. Not Fav. For GBP

Taylor (1990)	Monthly 1921:01-1925:05	GBP and USD	NER, WPI	Unit Root (PP, ADF, Johansen) Cointegration(GR ADF, 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.
Mc Nown and Wallace (1989)	Monthly: different periods (from 1972 to 1986)	CLP, ARS, BRL, ILS against USD	NER, WPI and CPI	Cointegration(CR DF, GRADF, CRDW)	Cointegration when WPI is used, not for CPI. Uses the ECM model to describe the mechanics of adjustment to the 1.r.equilibrium	Fav.for CLP, ARS and BRL
Mikkelsen (1989)	Quarterly 1948-1988	CLP, ARS, BRL, MXN, UYU and PEN	NER, Relative WPI	Unit Root (DF, ADF) Cointegration(GR ADF)	e and p in Brasil are I(2). In Peru, e is I(1) while p is I(2)	Fav.for CLP, ARS, MXN and UYU
Taylor (1988)	Monthly 1973:06-1985:12	GBP, DEM, FRF, CAD, JPY against USD	NER, MPI	Unit Root (DF, ADF) Cointegration(DF, ADF, DW)	Allowance for measurement error and/or transportation costs	Not Fav.
Taylor and McMahon (1988)	Monthly 1921:01-1925:05. Data for Germany 1921:02-1923:08	DEM, FRF, USD against GDP	NER, WPI	Unit Root (ADF), Cointegration(DR ADF, GRDW)	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 USD-GBP
Miller (1984)	FRF, GBP, DEM, USD	Quarterly 1973-1980	NER	GLS	Finds that deviations in prices are persistent	Fav (within EU)
Frenkel (1981)	Monthly: 1921-1925 1973-1979	USD, DEM, FRF, GBP	NER, WPI, CPI, MPI	2S-OLS, IV	Studies both absolute and relative PPP	Fav. (first, last period for EU) Not fav (last period for USD)
Krugman (1978)	Monthly: 1921-1925, 1973-1976	DEM, CHF,FRF, ITL, GBP and USD	NER, WPI	OLS, IV	Studies the autocorrelation and suggests that movements on exchange rates are due to ommited variables	Not Fav. Fav

4. CURRENT ECONOMETRIC DEVELOPMENTS-
Solving the problem of low-power tests.

Tests for evidence of PPP as a long-run phenomenon have often been based on an empirical examination of the real exchange rate. The most commonly used formal tests consist of whether the real exchange rate has a unit root. If the unit root hypothesis can be rejected, there is evidence of mean reversion of the real exchange rate, which means that the real exchange rate is stationary and PPP hypothesis holds.

Alternatively, another way to find evidence of PPP is by applying cointegration methods. Since cointegration between the exchange rate, domestic price level and foreign price level is a necessary condition for PPP, evidence of PPP can be found by testing the null hypothesis of no cointegration among these variables. The staple test statistic used is the augmented Dickey-Fuller (ADF).

Many economists have tested the long-run PPP empirically and were unanimous in their failure to reject the null hypothesis. Some researchers suggest that the widespread failure to find support of PPP in studies is due to the use of relatively short series of data. Tests employed to examine the long-run stability of real exchange rates that are based on data covering 15 years or so have low power to reject the null hypothesis even if it is false. Which means that the failure to reject the unit root hypothesis may be caused by the low power of the unit root tests in small samples.

The intuition behind the lack of power is straightforward. With cointegration or unit root tests, we are examining the long-run or low-frequency properties of a time series. If a researcher has very long run span of data and the series in question is indeed generated by a mean-reverting process, then he/she will have the opportunity to see it returning to the mean several times. Hence, the power to reject the false null hypothesis of non-mean reversion (non-cointegration) will tend to be higher in long spans of data. In the contrary, if a short span of data is available, only a small segment of the series will be observed and consequently, it will be harder for the researcher to see the mean reversion. Hence, the power to reject the hypothesis of non-mean reversion will be lower.

Lothian and Taylor (1997) and Sarno and Taylor (2002) show, on the basis of Monte Carlo experiments calibrated using “typical” results from the empirical real exchange rate literature, that the probability of rejecting the null hypothesis of a random walk real exchange rate when in fact the real rate is mean reverting would be somewhere between 5% and 8%. Alternatively, this implies the probability of never being able to reject the null hypothesis of a unit root is in excess of 92% when we have only 15 years if data available. Sarno and Taylor (2002) also show that if with the benefit of additional 10 years or so of data which are now available, the power of test increases only slightly, to a maximum of around 11% on the most optimistic view of the speed of mean reversion. Further Sarno and Taylor confirm the finding of Lothian and Taylor (1996) that even with a century of data there would be less than fifty-fifty chance of rejecting the unit root hypothesis.

Attempting to Circumvent the Power Problem: Long Spans of Data

One answer to the above problem has been to utilize longer spans of data. Frankel (1986); Lothian and Taylor (1996), among others, reject the unit root hypothesis with a century (or more) of real exchange rate data. These studies however combine data from fixed and floating exchange rate periods and do not provide direct evidence of PPP under the flexible exchange rates.

Long-span studies have been subject to some criticism in the literature. One criticism relates to the fact that because of the very long data spans involved, various exchange rate regimes are typically spanned. Also, real shocks may have generated structural breaks or shifts in the equilibrium real exchange rate (Hegwood and Papell, 1998). In order to provide a convincing test of real exchange rate stability during the post-Bretton Woods period, it is necessary to conduct a test using data for that period alone.

Attempting to Circumvent the Power Problem: Panel Data

An alternative approach, that has been proposed in the recent years, in order to overcome the problem of low-power of unit root tests is the use of panel data methods for empirical analysis. The additional cross-sectional dimension in the panel leads to better power properties of the panel tests as compared to the lower power of the standard individual-specific unit root tests. Testing for unit root, using panel data means utilizing data on more than one real exchange rates. This way the number of real exchange rates under consideration is increased, raising the power of the test by increasing the information employed.

So, panel data method is the new econometric development in the field of PPP hypothesis examination. This new advanced methodology gives the opportunity to the researchers to make a step further in their studies by solving some of the problems if the previous methods. Researchers, in the recent years, have employed in their vast majority this technique and report much stronger rejections of the unit root hypothesis of the real exchange rates.

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

Drawing our information from the review of the numerous studies testing the validity of the PPP hypothesis that we previously cited, in the preceding section we present the most important econometric methodologies that have been employed during the long history of PPP: cointegration and unit root tests.

5.1. 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, -β)', such that the combination

$$z_t = y_t - \alpha - \beta x_t \tag{9}$$

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)'~CI(d,b)] with (1, -β)' called the cointegrating vector. Notice that a constant term has been included 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. An (n ̦1) vector time series yt (e.g. of exchange rates and price indices) is said to be cointegrated if each of its elements individually is integrated of order 1 (denoted I(1)), i.e. nonstationary with a unit root, and there exists a nonzero (n x 1) vector α such that $\alpha' y_t$ is stationary. In this case, a is called the 'cointegrating vector'. Cointegration means that one or more linear combinations of the variables is stationary even though individually they are not (see, e.g. Dickey et al., 1991, and Hamilton, 1994). Since exchange rates and price indices are considered to be nonstationary variables which are frequently I(1) processes, the cointegration test is appropriate in determining PPP as a long-run equilibrium condition.

So, in the case of PPP, we have:

$$s_t = \alpha + \beta(p - p^*)_t + \varepsilon_t \tag{10}$$

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. Which means that it shows 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:

$$\Delta s_t = \mu + \sum_{i=0}^{p} f_i \Delta s_{t-i} + \sum_{j=1}^{q} g_j \Delta(p - p^*)_{t-j} + J\ z_{t-1} + x_t \tag{11}$$

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

Note that the term z_{t-1} 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 is estimated by ordinary least squares (OLS) and it is tested whether the cointegrating residuals $\hat{z}_t = s_t - \hat{a} - \hat{b}(p - p^*)_t$ are I(0).

ii) Finally, the residuals $\hat{z}_t$ are entered into the ECM that was disribed previously, where now all the variables are I(0) and conventional modelling strategies can be applied.

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 ρ-th order autoregressive representation of X_t

$$X_t = \Pi_1 X_{t-1} + \Pi_2 X_{t-2} + \ldots + \Pi_\rho X_{t-\rho} + \varepsilon_t \tag{12}$$

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

$$\Delta X_t = \tilde{\Pi}'_t \, \Delta \mathrm{X}_{t\text{-}1} + \mathbf{K} + \tilde{\Pi}'_{r+1} \Delta \mathrm{X}_{t-r+1} + \tilde{\Pi}'_r \mathrm{X}_{t-r} + e_t \tag{13}$$

where $\Pi'_r = -\Pi(1) \quad (= \text{-}(\Pi_1 + \mathbf{K} + \Pi_r))$. To estimate $\tilde{\Pi}'_r$ by maximum-likelihood, we estimate by OLS the following regressions:

$$\Delta X_t = \Gamma_{01} \Delta \mathrm{X}_{t-1} + \mathbf{K} + \Gamma_{0k-1} \Delta \mathrm{X}_{t-k+1} + e_{0t} \tag{14}$$

and

$$\Delta X_{t-r} = \Gamma_{11} \Delta \mathrm{X}_{t-1} + \mathbf{K} + \Gamma_{1k-1} \Delta \mathrm{X}_{t-k+1} + e_{1t} \tag{15}$$

and compute the product moment matrices of the residuals

$$\hat{\mathrm{M}}_{ij} = \mathrm{T}^{-1} \sum_{i=1}^{T} \hat{e}_{it} \hat{e}'_{jt} \quad , \qquad \text{i,j=0,1} \tag{16}$$

To test of the null hypothesis $\mathrm{H}_0 : \Pi'_r = \mathrm{B}\Gamma'$, i.e. there are at most r cointegrating vectors, can be conducted using either of the following two test statistics:

$$l_{ir}(r) = -T \sum_{i=r+1}^{r} \ln(1 - \hat{l}_i) \tag{17}$$

$$l_{\max}(r, r+1) = -T \ln(1 - \hat{l}_{r+1}) \tag{18}$$

Where $\hat{l}_{r+1}, \mathbf{K}, \hat{l}_r$ are the ρ-r smallest eigenvalues of $\hat{\mathrm{M}}_{10} \hat{\mathrm{M}}_{00} \hat{\mathrm{M}}_{01}$ with respect to $\mathrm{M}_{11,}$ obtained for the determinant

$$\left| \hat{l} \hat{\mathrm{M}}_{11} - \hat{\mathrm{M}}_{10} \hat{\mathrm{M}}_{00} \hat{\mathrm{M}}_{01} \right| = 0$$

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 test using Monte Carlo.

Cointegration tests are weak tests of PPP since they require only that some linear combination of domestic price level (p_t), the foreign price level (p_t*) and the nominal exchange rate (s_t) to be stationary. The null hypothesis here is that of no cointegration between the three series. A stronger test for PPP can be implemented by imposing proportionality(between relative prices and the exchange rate). and symmetry (as between domestic and foreign prices) restrictions upon the coefficients see Edison et al., 1997): Therefore, as a final step we will empirically investigate these conditions by restricting the cointegrating vector, α, to be (1,1, -1) and then testing for the stationarity of the real exchange rate: $R_t=\alpha' y_t=s_t + p_t - p_t^*$, where s, p and p* are the natural logs of S, P and P*, respectively, and variable Rt is expected to be stationary (implying that departures over time from PPP are self-correcting).

Finally, it is important to note that several factors—including transaction costs, existence of hyperinflation, discrepancies and/or interruptions in statistical releases, and differences in price indices across countries—may cause the cointegration-based test of PPP to fail. Consequently, finding support for PPP despite such obstacles should be regarded as very strong evidence confirming the long-run validity of this concept.

5.2. UNIT ROOTS

In this study we will examine the validity of PPP hypothesis by testing the real exchange rate stationarity. Investigation of long-run PPP is closely related to the issue of whether the real exchange rate contains a unit root.

The real exchange rate is defined as:

$$y_{it} = \log(\frac{s_t p_t^*}{p_t}) \tag{19}$$

where y_{it} is the real exchange rate, S_t is the nominal exchange rate in the home currency of the country i per unit of foreign currency f, P_t^* and P_t denote the foreign and domestic consumer price indices.

The purpose of this work is by employing both univariate and panel methods to will see in practice how these econometric methods work and check by the results of our research whether the higher statistical power of panel unit root tests provide stronger evidence of long-run PPP as theory states.

5.2.1 UNIVARIATE UNIT ROOT TESTS

-Generally

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:

$$y_t = r_n y_{t-1} + e_t \tag{20}$$

$$y_t = m_c + r_c y_{t-1} + e_t \tag{21}$$

$$y_t = m_{ct} + g_t + r_{ct} y_{t-1} + e_t \tag{22}$$

where e_t ~ iid(0, s_e^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 ρ_n<1, then the DGP is a stationary zero-mean AR(1) process and if ρ_n =1, then the GDP is a pure random walk. For equation (14), if ρ_c<1, then the DGP is a stationary AR(1) process with mean $m_c/(1-r_{c)}$ and if ρ_c =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 $\frac{m_{ct}}{1-r_{ct}}+g_{ct}\sum_{j=0}^{t}\left[r_{ct}^{j}(t-j)\right]$ 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:

$$\Delta y_t = (r_n - 1) y_{t-1} + e_t \tag{20'}$$

$$\Delta y_t = b_{0c} + (r_c - 1) y_{t-1} + e_t \tag{21'}$$

$$\Delta y_t = b_{0ct} t + b_{1ct} t + (r_{ct} - 1) y_{t-1} + e_t \tag{22'}$$

The tests are implemented though the usual t-statistic on the estimated (ρ-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).

So, 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 :

$$\Delta y_t = (r_n - 1) y_{t-1} + \sum_{i=1}^{p} a_i \Delta y_{t-i} + e_t \tag{20''}$$

$$\Delta y_t = b_{0c} + (r_c - 1) y_{t-1} + \sum_{i=1}^{p} a_i \Delta y_{t-i} + e_t \tag{21''}$$

$$\Delta y_t = b_{0c} t + b_{1ct} (r_{ct} - 1) y_{t-1} + \sum_{i=1}^{p} a_i \Delta y_{t-i} + e_t \tag{22''}$$

The tests are based on the t-ratio on $(\hat{r} - 1)$ 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 $(\hat{r} - 1)$ 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 Schward Bayesian information criterion (BIC). Alternatively, we can follow Hall (1994) general to specific sequential rule, starting with a large value of p ($_{max}$), testing the significance of the last coefficient and reducing *p* iteratively until a significant statistic is encountered.

-PPP in specific

The most common test for PPP is the univariate Augmented-Dickey-Fuller (ADF) test, which regresses the first difference of a variable (in this case the logarithm of the real exchange rate) on a constant, its lagged level and k lagged first differences,

$$\Delta y_t = m + a y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \tag{23}$$

A time trend is not included in equation (2) because such an inclusion would be theoretically inconsistent with long-run PPP. The null hypothesis of a unit root is rejected in favour of the alternative of level stationarity if α is significantly different from zero. Rejection of the unit root null provides evidence of mean reversion, which means that the real exchange rate is stationary and long-run PPP holds. Generally, these tests provide little evidence against unit roots for post-1973 real exchange rates.

In this research we examine the outcome of univariate unit root tests for the floating exchange rate period. The results of the univariate tests are reported in Table 4, Table 5, Table 6. As we have already mentioned we conduct univariate unit root tests for different spans of data (1971-1988 and 1973-2003) in order to investigate if by increasing the number of observations the test provides more strengthened evidence of PPP. Results for period 1971-1988 are reported in Table 4 and for period 1973-2003 are reported in Table 5. The value of t-statistics are in parentheses.

Table 4
Augmented Dickey-Fuller tests: United States Dollar real exchange rates (period 1973-1988)

Country	Monthly μ	Monthly α	Monthly ADF t-statistic	Quarterly μ	Quarterly α	Quarterly ADF t-statistic
Australia				-0.004 (-1.239)	-0.021 (-1.15)	-1.151
Austria	-0.014 (-1.138)	-0.012 (-1.250)	-1.250	-0.054 (-1.127)	-0.047 (-1.243)	-1.243
Belgium	-0.019 (-1.005)	-0.012 (-1.022)	-1.022	-0.077 (-1.171)	-0.047 (-1.188)	-1.188
Canada	-0.001 (-2.059)	-0.009 (-1.263)	-1.263	-0.002 (-1.760)	-1.760 (-1.373)	-1.373
Denmark	-0.007 (-1.200)	-0.008 (-1.121)	-1.121	-0.027 (-1.163)	-0.029 (-1.066)	-1.066
Finland	-0.007 (-1.761)	-0.010 (-1.597)	-1.597	-0.025 (-1.745)	-0.039 (-1.622)	-1.622
France	-0.005 (-1.262)	-0.006 (-1.060)	-1.060	-0.021 (-1.720)	-0.026 (-1.569)	-1.569
Germany	-0.003 (-0.612)	-0.011 (-1.088)	-1.088	-0.011 (-0.699)	-0.038 (-1.138)	-1.138
Greece	-0.004 (-1.909)	0.002 (-0.835)	0.835	-0.008 (-1.027)	-0.001 (-0.135)	-0.135
Ireland	0.002 (-0.929)	-0.009 (-2.037)	-2.037	0.006 (0.890)	-0.030 (-1.967)	-1.967
Italy	-0.015 (-1.925)	-0.004 (-1.644)	-1.644	-0.071 (-2.167)	-0.021 (-1.809)	-1.809
Japan	0.013 (-0.862)	0.005 (-0.767)	0.767	0.073 (1.259)	0.027 (1.152)	1.152
Netherlands	-0.004 (-0.738)	-0.010 (-0.948)	-0.948	-0.021 (-1.030)	-0.044 (-1.193)	-1.193
New Zealand				-0.007 (-1.793)	-0.011 (-0.870)	-0.870
Norway	-0.004 (-0.675)	-0.005 (-0.549)	-0.549	-0.013 (-0.663)	-0.014 (-0.536)	-0.536
Portugal	-0.010 (-3.334)	-0.001 (-0.679)	-0.679	-0.028 (-2.869)	-0.008 (-1.192)	-1.192
Spain	-0.013 (-2.142)	-0.006 (-1.562)	-1.562	-0.040 (-1.997)	-0.017 (-1.462)	-1.462
Sweden	-0.004 (-1.033)	-0.005 (-0.732)	-0.732	-0.014 (-1.009)	-0.016 (-0.723)	-0.723
Switzerland	-0.002 (-0.462)	-0.011 (-1.391)	-1.391	-0.006 (-0.533)	-0.035 (-1.401)	-1.401
UK	0.004 (-1.352)	-0.016 (-2.224)	-2.224	0.014 (1.666)	-0.050 (-2.353)	-2.353

***,**,* denote rejection of the unit root null with statistical significance at 1%, 5%, 10% respectively.

Critical values

	1%	5%	10%
Monthly	-3.465	-2.877	-2.575
Quarterly	-3.538	-2.908	-2.592

$$\Delta y_t = m + a y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t$$

The critical values are from MacKinnon(1996), adjusted for 370 monthly and 124 quarterly objervations

Table 5
Augmented Dickey-Fuller tests: United States Dollar real exchange rates (period 1973-2003)

Country	Monthly μ	Monthly α	Monthly ADF t-statistic		Quarterly μ	Quarterly α	Quarterly ADF t-statistic	
Australia					-0.005	-0.024	-2.209	
					(-2.236)	(-2.209)		
Austria	-0.010	-0.009	-1.875		-0.038	-0.035	-1.889	
	(-1.726)	(-1.875)			(-1.726)	(-1.889)		
Belgium	-0.017	-0.011	-1.320		-0.096	-0.061	-2.345	
	(-1.285)	(-1.320)			-2.317	-2.345		
Canada	-0.001	-0.011	-2.050		-0.003	-0.027	-2.335	
	(-2.315)	(-2.045)			(-2.184)	(-2.335)		
Denmark	-0.009	-0.011	-1.800		-0.043	-0.051	-2.110	
	(-1.805)	(-1.800)			(-2.100)	(-2.110)		
Finland	-0.008	-0.012	-2.296		-0.033	-0.049	-2.574	
	(-2.379)	(-2.296)			-2.615	-2.574		
France	-0.007	-0.009	-1.803		-0.027	-0.035	-2.205	
	(-1.846)	(-1.803)			(-2.255)	(-2.205)		
Germany	-0.002	-0.010	-1.839		-0.007	-0.031	-1.927	
	(-1.001)	(-1.839)			(-1.133)	(-1.927)		
Greece	-0.007	-0.002	-2.264		-0.013	-0.004	-1.559	
	(-4.467)	-2.264			(-2.062)	-0.257		
Ireland	0.002	-0.011	-2.969	**	0.008	-0.035	-2.667	*
	(-1.914)	(-2.969)			1.914	-2.667		
Italy	-0.017	-0.005	-2.710	*	-0.070	-0.022	-2.776	*
	(-2.932)	(-2.710)			-2.996	-2.776		
Japan	-0.002	-0.001	-0.486		-0.002	-0.003	-0.297	
	(-0.288)	(-0.486)			(-0.078)	(-0.297)		
Netherlands	-0.002	-0.007	-1.409		-0.007	-0.022	-1.466	
	(-0.971)	(-1.409)			-1.100	-1.466		
New Zealand					-0.005	-0.017	-2.365	
					(-2.304)	(-2.365)		
Norway	-0.007	-0.009	-1.379		-0.023	-0.028	-1.420	
	(-1.443)	(-1.378)			(-1.481)	(-1.420)		
Portugal	-0.014	-0.005	-4.107	***	-0.035	-0.014	-3.545	***
	(-5.854)	(-4.107)			-4.569	-3.545		
Spain	-0.016	-0.007	-2.766	*	-0.042	-0.019	-2.447	
	(-3.204)	(-2.766)			-2.777	-2.447		
Sweden	-0.006	-0.006	-1.475		-0.019	-0.024	-1.740	
	(-1.716)	(-1.475)			(-1.849)	(-1.740)		
Switzerland	-0.001	-0.010	-2.148		-0.003	-0.032	-2.128	
	(-0.620)	(-2.148)			(-0.694)	(-2.128)		
UK	0.004	-0.016	-3.092	**	0.010	-0.048	-3.098	**
	-2.127	(-3.092)			(-2.122)	(-3.098)		

***,**,* denote rejection of the unit root null with statistical significance at 1%, 5%, 10% respectively.

Critical values

	1%	5%	10%
Monthly	-3.448	-2.869	-2.571
Quarterly	-3.484	-2.885	-2.579

$$\Delta y_t = m + a y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t$$

The critical values are from MacKinnon(1996), adjusted for 370 monthly and 124 quarterly objervations

With the US dollar used as the base currency, as we can derive from the results stated above that the univariate tests provide little support for PPP. More specifically, no support is provided during the first period of examination (1973-1988), where the unit root null cannot be reject for none of the 20 countries tested (18 countries when monthly data are analysed because for Australia and New Zealand there are no monthly observations available). This result indicates, from an empirical point of view, the low power of univariate unit root tests to support the PPP hypothesis for short spans of data that we have already mentioned in our theoretical analysis. This result appears both for monthly and quarterly data which means that attempting to circumvent the problem of low power of univariate unit root tests by increasing the frequency of observation- say from quarterly to monthly is not the appropriate solution.

The empirical research is unable to provide great evidence of PPP even for the second period of examination (1973-2003). As we can see from table 5 the unit root null can be rejected only for five out of 20 real exchange rates in the case of monthly data and for four out of 20 real exchange rates in the case of quarterly data. So, increasing the span of the data does not strengthen the evidence of PPP for the dollar based real exchange rates. This initially may seem surprising, since the number of observations doubles between the sample that ends in 1988 and the one that ends in 2003 but once again proves empirically that the low power of univariate unit root tests accounts for their inability to reject the unit root null for short spans of data .

We also conduct univariate unit root tests adding one year of data each time in order to see how this extra year of data influences the result in favour or against the unit root null. The results of these sequential tests are reported in Table 6. Specifically, the row marked 1988 uses data for 1973-1988, each row adds another year of data, ending with the row marked 2003 which uses data for 1973-2003.

The results of this test show that by increasing the sample the power of the test is not increased significantly. Increasing the span of the data does not strengthen the evidence of PPP for either monthly or quarterly based real exchange rates. The results of the univariate tests indicate that when the US dollar is used as the base currency, the univariate tests provide little support for PPP. The unit root null can be rejected for at most five out of 20 real exchange rates and in seven of the years cannot be rejected for any country, for monthly data. While for quarterly data support is even weaker with rejection of the null for at most four countries.

Table 6
Rejections of the unit root null with univariate ADF tests

Year	US Dollar					
	Monthly			Quarterly		
	1%	5%	10%	1%	5%	10%
1988	0	0	0	0	0	0
1989	0	0	0	0	0	0
1990	0	0	0	0	0	1
1991	0	0	0	0	0	0
1992	0	0	0	0	0	0
1993	0	0	0	0	0	0
1994	0	0	0	0	0	0
1995	0	0	2	0	0	0
1996	0	1	2	0	0	2
1997	0	0	3	0	0	2
1998	0	1	2	0	1	1
1999	0	1	2	0	1	1
2000	0	1	3	0	1	1
2001	0	1	3	0	1	2
2002	1	1	2	0	2	2
2003	1	2	2	1	1	2

The entries denote the number of times, out of a maximum of 20, that the unit root null can be rejected at the various significance levels. The critical values are from MacKinnon 1996 , adjusted for the exact number of observations

More specifically, in an attempt to interpret what the numbers in the previous table we must point out that as the span of data is augmented the evidence of PPP is strengthened, but the support is very small. For the first five years of data, the univariate tests provide no evidence of PPP for the examined countries. For the first time, during the period 1973-1995, the unit root null is rejected for Portugal and Ireland at 10% significance level. By adding another year of data UK also shows evidence of stationarity in real exchange rates at 10% significance level, while Portugal rejects the unit root null at 5% significance level. These results indicate that support of PPP hypothesis is increased but again the number of countries that show supportive evidence is very small.

Finally, the greatest evidence of PPP is shown for the whole period of flexible exchange rates, which however is not again very satisfactory. For thirty years of data only Italy, Spain, Portugal, Ireland and UK reject the unit root hypothesis. Results are even less supportive for quarterly data. This inability to reject the unit root null with univariate tests has stimulated interest in panel unit root tests, which exploit cross-section variation in order to increase power with short time spans of data. Three of the most important panel tests are conducted in the next section of this research.

5.2.2. PANEL UNIT ROOT TESTS

A variety of procedures are employed for the analysis of unit roots in panel in order to evaluate real exchange rate stationarity over the floating period. As we have already mentioned above, univariate unit root tests often lack sufficient power in relatively small samples like the one we are testing (1973-2003). To overcome this problem we consider pooling cross-section time series data for testing the unit root hypothesis. So, we will employ panel data methods by adopting the three most important panel unit root tests developed by: Levin and Lin (1993), Im Pesaran and Shin (1997) and Pesaran (2003). These tests differ in their assumptions, asymptotic and finite sample properties as well as their implications for PPP and possess important differences concerning the speeds of adjustment and testing for real exchange rate stationarity.

In this section, as a first step we will give the basic technical characteristics of the three panel tests that we use in our research. Secondly, using both monthly and quarterly data we will construct panels for the 21 countries classified by the IMF as industrialized and employ the panel tests to them. Finally, we will give a comparison of the employed panel tests indicating their common features and their differences both in the theoretical framework and the empirical results.

I) Levin and Lin test

In finite samples, univariate unit root test procedures are known to have limited power against alternative hypotheses with highly persistent deviations from equilibrium. Simulation exercises also indicate that this problem is particularly severe for small samples (see Campbell and Perron, 1991).

The work of Levin and Lin aimed at pooling cross-section time series data as a means of generating more powerful unit root tests. The test procedures are designed to evaluate the null hypothesis that each individual in the panel has integrated time series versus the alternative hypothesis that all individuals time series are stationary. Which means that this approach jointly tests if all series in the panel follow a unit root process. The pooling approach yields higher test power than performing a separate unit root test for each individual. The panel-based unit root test proposed allows for individual-specific intercepts and time trends. Moreover, the error variance and the pattern of higher-order serial correlation are also permitted to vary freely across individuals.

A. Model Specifications

Starts with studying the stochastic process {yit} of a panel where i=1,….., N and t=1,……T . the aim is to determine whether {yit} is integrated for each individual in the panel. As in the case of a single time series, the individual regression may include an intercept and time trend. The only assumption maid is that all individuals in the panel have identical 1rst-order partial autocorrelation, but all other parameters in the error process are permitted to vary freely across individuals.

Assumption 1:

a) Assume that {yit} is generated by one of the following models:

$$\text{Model 1}: \Delta y_{it} = dy_{t-1} + z_{it} \tag{24}$$

$$\text{Model 2}: \Delta y_{it} = a_{0i} + dy_{t-1} + z_{it} \tag{25}$$

$$\text{Model 3}: \Delta y_{it} = a_{0i} + a_{1i}t + dy_{t-1} + z_{it} \quad \text{where } -2 < \delta \leq 0 \text{ for all } i = 1, \ldots\ldots, N \tag{26}$$

b) The error process ζit is distributed independently across individuals and follows a stationary invertible ARMA process (individual time series may exhibit serial autocorrelation)

$$z_{it} = \sum_{j=1}^{\infty} q_{ij} z_{it-j} + e_{it} \tag{27}$$

c) For all i=1,….., N and t=1,……T,

$$\mathrm{E}(z_{it}^4) < \infty \ ; \ \mathrm{E}(e_{it}^2) \geq \mathrm{B}_e > 0 \ ; \text{and } \mathrm{E}(z_{it}^2 + 2\sum_{j=1}^{\infty} \mathrm{E}(z_{it} z_{it-j}) < \mathrm{B}_z < \infty$$

Assumption 1(a) includes three data generating processes. In Model 1, the panel unit root test procedure evaluates the null hypothesis H_0: δ=0 against the alternative H_1: δ<0. The series {yit} has an individual-specific mean in Model 2, but does not contain a time trend. In this case, the panel test procedure evaluates the null hypothesis that H_0: δ=0 and α_{0i} =0, for all i, against H1: δ<0 and $\alpha_{0i} \in$ R . Finally, under Model 3, the series {yit} has an individual-specific mean and time trend. In this case, the panel test procedure evaluates the null hypothesis that H_0: δ=0 and α_{1i} =0, for all i, against the alternative H_1: δ<0. and $\alpha_{1i} \in$ R.

As in the a single time series, if a deterministic element (e.g. an intercept or time trend) is present but not included in the regression procedure, the unit root test will be inconsistent. On the other hand, if a deterministic element is included in the regression procedure but is not present in the observed data, the statistical power of the unit root test will be reduced. See Johansen (1992) discussion on the interactions of the unit root test and various deterministic specifications.

Assumption 1(b) is standard; individual time series may exhibit serial correlations. Finally, the boundedness conditions in Assumption 1(c) ensure this ratio remains finite for every individual in the panel as the cross-section N becomes arbitrarily large.

B. Test Procedures

Our main hypothesis is:

$$\Delta y_{it} = dy_{it-1} + \sum_{L=1}^{p_i} q_{iL} \Delta y_{it-L} + a_{mi} d_{mt} + e_{it}, \qquad \text{m} = 1,2,3. \tag{28}$$

And the different models that are being tested are:

- **Model$_1$ (m=1) - d1t={0}**

No deterministic terms included
Ho: δ=0
H1: δ<0

- **Model$_2$ (m=2) - d2t={1}**

The series {yit} has an individual – specific constant but does not contain a time trend
Ho: δ=0 and α0i=0
H1: δ<0 and α0i.R for all i

- **Model$_3$ (m=3) - d3t={1, t}**

The series {yit} has an individual – specific mean and trend
Ho: $\delta=0$ and $\alpha 1i=0$
H1: $\delta<0$ and $\alpha 1i_{\in} R$ for all i

However, since p_i (number of lags) is unknown, we therefore a three-step procedure is suggest to implement this test. In step 1 separate ADF regressions is carried out for each individual in the panel, and orthogonalized residuals are generated. Step 2 requires estimating the ratio of long run to short run innovation standard deviation for each individual. In the final step the pooled t-statistics are computed.

- *Step 1: Perform ADF regressions and generate orthogonalized residuals*

For each individual i, we implement the ADF regression

$$\Delta y_{it} = d_i y_{it-1} + \sum_{L=1}^{p_i} q_{iL} \Delta y_{it-L} + a_{mi} d_{mt} + e_{it}, \qquad m = 1,2,3. \tag{28'}$$

The lag order pi is permitted to vary across individuals. Campbell and Perron (1991) recommend the method proposed by Hall (1990) for selecting the appropriate lag order: for a given sample length T, choose a maximum lag order p_{max}, and then use the t-statistics of $\hat{q}_{iL}$ to determine if a smaller lag order is preferred. (These t-statistics have a standard normal distribution under the null hypothesis (q_{iL} =0), both when $\delta i=0$ and when $\delta_i<0$.

Having determined autoregression order p_i in (1’), we run two auxiliary regressions to generate orthogonalized residuals. Regress Δy_{it} and y_{it-1} against Δy_{it-L} (L=1,…, p_i) and the appropriate deterministic variables, d_{mt}, then save the residuals $\hat{e}_{it}$ and $\hat{u}_{it-1}$ from these regressions. Specifically,

$$\hat{e}_{it} = \Delta y_{it} - \sum_{L=1}^{P_i} \hat{p}_{iL} \Delta y_{it-L} - \hat{a}_{mi} d_{mt} \tag{29}$$

and

$$\hat{u}_{it-1} = y_{it-1} - \sum_{L=1}^{p_i} \hat{p}_{iL} \Delta y_{it-L} \hat{a}_{mi} d_{mt} \tag{30}$$

To control for heterogeneity across individuals, we further normalize $\hat{e}_{it}$ and $\hat{u}_{it-1}$ by the regression standard error form Eq. (28').

$$\tilde{e}_{it} = \frac{\hat{e}_{it}}{\hat{s}_{ei}}, \qquad \tilde{u}_{it-1} = \frac{\hat{u}_{it-1}}{\hat{s}_{ei}} \tag{31}$$

where $\hat{s}_{ei}$ is the regression standard error in (28'). Equivalently, it can also be calculated from the regression of $\hat{e}_{it}$ against $\hat{u}_{it-1}$

$$\hat{s}_{ei}^{2} = \frac{1}{T - P_i - 1} \sum_{t=p_{i+2}}^{T} (\hat{e}_{it} - \hat{d}_i \hat{u}_{it-1})^2 \tag{32}$$

- *Step 2: Estimate the ratio of long-run to short-run standard deviations*

Under the null hypothesis of a unit root, the long-run variance for Model 1 can be estimated as follows:

$$s_{yi}^{2} = \frac{1}{T-1} \sum_{t=2}^{T} \Delta y_{it}^{2} + 2 \sum_{L=1}^{\bar{k}} w_{\bar{K}L} \left[\frac{1}{T-1} \sum_{t=2+L}^{T} \Delta y_{it} \Delta y_{it-L} \right] \tag{33}$$

For Model, we replace Δy_{it} in (5) with $\Delta y_{it} - \overline{\Delta y}_{it}$, where $\overline{\Delta y}_{it}$ is the average value of Δy_{it} for individual i. If the data include a time trend (Model 3), then the trend should be removed before estimating the long-run variance The truncation lag parameter $\bar{K}$ can be data dependent. Andrews (1991) suggests a procedure to determine $\bar{K}$ to ensure the consistency of The sample covariance weights $w_{\bar{K}L}$ depend on the choice of kernel.

For each individual i, we define the ratio of the long-run standard deviation to the innovation standard deviation,

$$s_i = s_{yi} / s_{ei}$$

Denote its estimate by $\hat{s}_i = \hat{\sigma}_{yi} / \hat{\sigma}_{ei}$. Let the average standard deviation ratio be $S_N = (1/N)\sum_{i=1}^{N} s_i$ and its estimator $\hat{S}_N = (1/N)\sum_{i=1}^{N} \hat{s}_i$. This important statistic will be used to adjust the mean of the t-statistic later in step 3.

- *Step 3: Compute the panel test statistics*

Pool all cross sectional and time series observations to estimate:

$$\tilde{e}_{it} = \delta \tilde{u}_{it-1} + e_{it} \quad , \tag{34}$$

based on a total of $N\tilde{T}$ observations, where $\tilde{T} = T - \tilde{p} - 1$ is the average number of observations per individual in the panel, and $\tilde{p} \equiv \frac{1}{N}\sum_{i=1}^{N} p_i$ is the average lag order for the individual ADF regressions. The conventional regression t-statistic for testing δ=0 is given by

$$t_\delta = \frac{\hat{\delta}}{STD(\hat{\delta})} \tag{35}$$

Where

$$\hat{\delta} = \frac{\sum_{i=1}^{N}\sum_{t=2+p_i}^{T} \tilde{u}_{it-1}\tilde{e}_{it}}{\sum_{i=1}^{N}\sum_{t=2+p_i}^{T} \tilde{u}_{it-1}^{2}}, \tag{36}$$

$$STD(\hat{\delta}) = \hat{\sigma}_{\tilde{e}} \left[\sum_{i=1}^{N} \sum_{t=2+p_i}^{T} \tilde{u}_{it-1}^{2} \right]^{-1/2} \tag{37}$$

C. Empirical Results of LL test

In our empirical research, we use the LL test in order to test for evidence of PPP in the floating exchange rate period (1973-2003). As we have already mentioned this method provides a more general testing framework. From the three models that the test suggests we have chosen the second as the data used for conducting the test show intercept but no trend.

Table 7

Panel Unit Root tests: Results from LL test[1]

	Monthly				Quarterly			
Panels	**estimated δ**	**t_δ**	**modified t-stat**	**result**	**estimated δ**	**t_δ**	**modified t-stat**	**result**
All 18	-0.004	-6.580	-8.498***	Reject unit root null- Accept PPP				
All 20					-0.015	-8.116	-10.434***	Reject unit root null - Accept PPP
EC (15)	-0.004	-5.932	-7.663***	Reject unit root null- Accept PPP	-0.013	-6.053	-7.787***	Reject unit root null - Accept PPP
EMU (12)	-0.003	-5.280	-6.823***	Reject unit root null- Accept PPP	-0.013	-6.009	-7.728***	Reject unit root null - Accept PPP
G6	-0.006	-4.450	-5.748***	Reject unit root null- Accept PPP	-0.025	-5.001	-6.417***	Reject unit root null - Accept PPP
G10	-0.007	-5.737	-7.403***	Reject unit root null- Accept PPP	-0.027	-6.251	-8.001***	Reject unit root null - Accept PPP
OECD (13)	-0.009	-7.228	-9.324***	Reject unit root null- Accept PPP	-0.031	-7.401	-9.463***	Reject unit root null - Accept PPP

***,**,* denote rejection of the unit root null with statistical significance at 1%, 5%, 10% respectively.

The critical values derive from standard normal distribution

$$\Delta y_{it} = dy_{it-1} + \sum_{L=1}^{p_i} q_{iL}\Delta y_{it-L} + a_{mi}d_{mt} + e_{it}, \qquad m = 1,2,3.$$

Ho: δ=0 and α0i=0

H1: δ<0 and α0i.R for all i

[1] The countries that are included in each panel are presented analytically in section 6

By employing the Levin and Lin test for examining the null hypothesis (unit root hypothesis) for seven different panels of countries, we reject at 1% significance level the null hypothesis of non stationarity of the real exchange rates of the countries that consist the panels, for all the groups under consideration. As it is shown in table 7 the PPP hypothesis is accepted in all the cases. In contrast to the univariate unit root tests conducted in the previous section this method provides absolute support to PPP.

This outcome indicates the great difference of the conclusions derived from univariate and panel unit root methods and shows that panels provide considerably more evidence against unit root than is found by univariate tests. These results that show absolute rejection of the unit root null at 1% significance level are probably due to the restrictive structure of the test, that by testing such a strict hypothesis (β=0 for all the countries), rejects almost in every case the null hypothesis.

II) IM, Pesaran and Shin (IPS) test

The work of Im, Pesaran and Shin (IPS) proposes an alternative testing procedure based on averaging individual unit root test statistics for panels which relaxes the assumption of identical first order autoregressive coefficients of Levin and Lin approach.

In particular, they propose a test based on the average of (augmented) Dickey–Fuller (Dickey and Fuller, 1979) statistics computed for each group in the panel, which they refer to as the t-bar test. Like the LL procedure, their proposed test allows for residual serial correlation and heterogeneity of the dynamics and error variances across groups. Under very general settings this statistic is shown to converge in probability to a standard normal variate sequentially with $T \rightarrow \infty$, followed by $N \rightarrow \infty$. A diagonal convergence result with T and $N \rightarrow \infty$ while $N/T \rightarrow k$, k being a finite non-negative constant, is also conjectured.

In the special case where errors in individual Dickey.Fuller (DF) regressions are serially uncorrelated a modified version of the (standardized) t-bar statistic, denoted by $Z_{t\text{-bar}}$, is shown to be distributed as standard normal as $N \rightarrow \infty$ for a fixed T, so long as $T > 5$ in the case of DF regressions with intercepts, and $T > 6$ in the case of DF regressions with intercepts and linear time trends. An exact fixed N and T test is also developed using the simple average of the DF statistics. Based on stochastic simulations it is shown that the standardized t-bar statistic provides an excellent approximation to the exact test even for relatively small values of N.

A. The basic framework

Consider a sample of N cross sections (industries, regions or countries) observed over T time periods. We suppose that the stochastic process, {y_{it}}, is generated by the first-order autoregressive process:

$$y_{it} = (1 - j_i)m_i + j_i y_{i,t-1} + e_{it} \qquad \text{i} = 1, \mathbf{K}, \text{N}, \ \text{t} = 1, \mathbf{K} \ \text{T} \tag{38}$$

Where initial values, y_{i0}, are given. We are interested in testing the null hypothesis of unit roots $j_i = 1$ for all i (38) can be expressed as:

$$\Delta y_{it} = a_i + b_i y_{i,t-1} + e_{it} \tag{39}$$

where $a_i = (1 - j_i)m_i$, $b_i = -(1 - j_i)$ and $\Delta y_{it} = y_{it} - y_{i,t-1}$. The null hypothesis of unit roots then becomes:

Ho: βi=0 for all i

against the alternatives,

H_1: βi < 0 ,i=1,2,….,N1 , βi=0 , i=N1+1, N1+2,….,N

This formulation of the alternative hypothesis allows for β_i to differ across groups, and is more general than the homogeneous alternative hypothesis, namely $\beta_i=\beta<0$ for all i, which is implicit in the testing approach of Levin and Lin (LL). It also allows for some (but not all) of the individual series to have unit roots under the alternative hypothesis. Formally we assume that under the alternative hypothesis the fraction of the individual processes that are stationary is non-zero, namely if $\lim_{N\to\infty}(N_1 / N) = d, \quad 0 < d \leq 1$. This condition is necessary for the consistency of the panel unit root tests.

B. Unit root tests for heterogeneous panels with serially correlated errors

In this section we consider the more general case where the errors in (40) may be serially correlated, possibly with diEerent serial correlation patterns across groups, but with T and N sufficiently large.

Suppose that $\{y_{it}\}$ are generated according to the following finite-order AR(p_i+ 1) processes:

$$y_{it} = m_i j_i(1) + \sum_{j=1}^{p_1+1} j_{ij} y_{i,t-j} + e_{it}, \qquad \text{i=1,…,N, t=1,…,T} \tag{40}$$

which can be written equivalently as the ADF(p_i) regressions:

$$\Delta y_{it} = a_i + b_i y_{i,t-1} + \sum_{j=1}^{p_i} r_{ij} \Delta y_{i,t-j} + e_{it} \qquad \text{i=1,…,N, t=1,…,T} \tag{41}$$

where $j_i(1) = 1 - \sum_{j=1}^{p_i+1} j_{ij}$, $a_i = m_i j_i(1)$, $b_i = -j_i(1)$ and $r_{ij} = \sum_{h=j+1}^{p_i+1} j_{ih}$.

Writing the ADF regressions for each i in matrix notations we have:

$$\Delta y_i = b_i y_{i,-1} + Q_i g_i + e_i, \tag{42}$$

where $Q_i = t_T, \Delta y_{i,-1}, \Delta y_{i,-2}, \mathbf{K}, \Delta y_{i,-p_i}$ and $g_i = (a_i, r_{i1}, r_{i2}, \mathbf{K}, r_{ip_i})'$.

We also make the following assumptions:

Assumption 2.1. All the roots of $j_i(Z) = \sum_{j=1}^{p_i+1} j_{ij} Z^j = 0$, i = 1,2,…,N fall on or outside the unit circle, while all the roots of $r_i(Z) = \sum_{j=1}^{p_i} r_{ij} Z^j = 0$, i = 1,2,…,N fall strictly outside the unit circle.

Assumption 2.2. e_{it} , i = 1,2,…,N, t = 1,2,…,T in (41) are independently distributed as normal variates with zero means and finite (possibly) heterogeneous variances, s_i^2 , and the initial values, $y_{i0}, y_{i,-1} \mathbf{K}, y_{i,-p_i}$, are given (either fixed or stochastic).

The t-bar statistic is formed as a simple average of the individual t statistic for testing $\beta_\iota = 0$ in (41), namely

$$t - bar_{NT} = \frac{1}{N}\sum_{i=1}^{N} t_{iT}\left(p_i, r_i\right) \tag{43}$$

Where $t_{iT}(p_i, r_i)$ is given by

$$t_{iT}(p_i, r_i) = \frac{\sqrt{T - p_i - 2}}{(y_{i,-1}' M_{Q_t} y_{i,-1})^{1/2}} \frac{(y_{i,-1}' \mathrm{M}_{Q_t} y_i)}{(\Delta y_i' M x_t \Delta y_i)^{1/2}}, \tag{44}$$

$r_i = (r_{i1}, r_{i2}, \mathbf{K}, r_{ip_i})'$, $\mathrm{M}_{Q_t} = I_T - Q_i (Q_i' Q_i)^{-1} Q_i'$,

$\mathrm{M}_{X_t} = I_T - Q \backslash X_i (X_i' X_i)^{-1} X_i'$ and $X_i = (y_{i,-1}, Q_i)$

When T is fixed, the individual ADF statistics, $t_{iT}(p_i, r_i)$ will depend on the nuisance parameters, ρ_i, i = 1,…, p_i even under $\beta_i = 0$. Therefore, the standardization using $E[t_{iT}(p_i, r_i)]$ and $Var[t_{iT}(p_i, r_i)]$ will not be practical. But when T and N are sufficiently large it is possible to develop asymptotically valid t-bar type panel unit root tests that are free from the nuisance parameters.

Final step of practical value would be to carry out the standardization of the t-bar statistic using the means and variances of $t_{iT}(p_i, 0)$ evaluated under $\beta_i = 0$. This is likely to yield better approximations, since $E[t_{iT}(p_i, 0) \mid \beta_i = 0]$, for example, makes use of the information contained in p_i while $E[t_{iT}(0,0) \mid \beta_i = 0]$ does not. In view of this, the limiting distribution for standardized t-statistic is given as:

$$W_{tbar}(p, r) = \frac{\sqrt{\mathrm{N}}\{t\text{ - bar}_{\mathrm{NT}} - \frac{1}{N}\sum_{i=1}^{N} E[t_{iT}(p_i, 0) \mid b_i = 0]\}}{\sqrt{\frac{1}{N}\sum_{i=1}^{N} Var[t_{iT}(p_i, 0) \mid b_i = 0]}} \Rightarrow N(0,1) \tag{45}$$

The preceding analysis can be readily extended to unbalanced panels and/or to dynamic panels with intercepts and linear time trends. For this reason they computed $E[t_{iT}(p_i,0) \; |b_i = 0]$ and $Var[t_{iT}(p_i,0) \; |b_i = 0]$ for different values of T and p (lag length) via stochastic simulations with 50,000 replications when the underlying regression is estimated with and without a linear time trend.

C. Empirical Results of IPS test

As we move forward this analysis, we illustrate the results of the Im, Pesaran and Shin test that allows for a not so strict zero hypothesis. The results of the IPS panel unit root test support the rejection of the unit root hypothesis for the majority of the panels for both monthly and quarterly data. As it is demonstrated in table 8 the null hypothesis is rejected for the panel that includes all the countries, the countries of European Union, the countries that have adopted the Euro as their national currency unit (EMU), the 10 most industrialised countries (G10) and finally the panel of 13 of the OECD countries. On the contrary, the real exchange rate stationarity it is not able to be ascertained for the panel of the 6 most industrialised countries (G6). In this case the null hypothesis for the existence of unit root in this panel of countries is not able to be rejected at any significance level.

The fact that the test is not able to support the PPP hypothesis for panel that includes the smaller number of countries leads us to come to the conclusion that as the size of the panel augments the power of the test becomes bigger as well. The strength of our evidence against the unit root for real exchange rates appears to be correlated with the size of panels.

With monthly as well as quarterly data the results are in favour of PPP for the the more numerous panels. So, the results of the IPS panel test that in their majority strongly support the evidence of PPP indicate, as the previous one, that panel methods show greater support in the rejection of the unit root hypothesis than the univariate unit root tests.

Table 8
Panel Unit Root tests: Results from IPS test[2]

	Monthly		Quarterly	
Panels	**w-stat**	**result**	**w-stat**	**result**
All 18	-3.157***	Reject unit root null-Accept PPP		
All 20			-3.382***	Reject unit root null-Accept PPP
EC (15)	-3.237***	Reject unit root null-Accept PPP	-2.872***	Reject unit root null-Accept PPP
EMU (12)	-2.914***	Reject unit root null-Accept PPP	-2.621***	Reject unit root null-Accept PPP
G6	-1.632	Accept unit root null-Reject PPP	-1.619	Accept unit root null-Reject PPP
G10	-2.003**	Reject unit root null-Accept PPP	-1.922*	Reject unit root null-Accept PPP
OECD (13)	-2.641***	Reject unit root null-Accept PPP	-2.425**	Reject unit root null-Accept PPP

***,**,* denote rejection of the unit root null with statistical significance at 1%, 5%, 10% respectively.

The critical values derive from standard normal distribution

$$\Delta y_{it} = a_i + b_i y_{i,t-1} + \sum_{j=1}^{p_i} r_{ij} \Delta y_{i,t-j} + e_{it}$$

$Ho: \beta i=0$ for all i

$H_1: \beta i < 0, i=1,2,\ldots,N1, \ \beta i=0, \ i=N1+1, N1+2,\ldots,N$

III) Pesaran (CIPS) test

A new way to deal with the problem of cross sectional dependence is given by Pesaran (2003) who proposes a simple alternative panel unit root test, CIPS test, which focuses on a generalization of the t-bar test proposed by IPS and considers a cross-sectionally augmented version of the IPS. More specifically, instead of basing the unit root tests on deviations from the estimated factors, we augment the standard DF (or ADF) regressions with the cross section averages of lagged levels and first-differences of the individual series. Standard panel unit root tests can now be based on the simple averages of the individual cross sectionally augmented ADF statistics (denoted by CADF), or suitable transformations of the associated rejection

[2] The countries that are included in each panel are presented analytically in section 6

probabilities. The individual CADF statistics or the rejection probabilities can then be used to develop modified versions of the t-bar test proposed by Im, Pesaran and Shin (IPS), the inverse chi-squared test (or the P test) proposed by Maddala and Wu (1999), and the inverse normal test (or the Z test) suggested by Choi (2001).

A truncated version of the test is also considered where the individual CADF statistics are suitably truncated to avoid undue influences of extreme outcomes that could arise when T is small (in the region of 10-20). New asymptotic results are obtained both for the individual CADF statistics, and their simple averages, referred to as the cross-sectionally augmented IPS (CIPS) test.

A. A Simple Dynamic Panel with Cross-Section Dependence

Let y_{it} be the observation on the i^{th} cross-section unit at time t and suppose that it is generated according to the following simple dynamic linear heterogeneous panel data model

$$y_{it} = (1 - j_i)m_i + j_i y_{i,t-1} + u_{it}, \quad i = 1, \ldots, \mathrm{N};\ t = 1, \ldots, \mathrm{T} \qquad (46)$$

where initial value, y_{i0}, is given, and the error term, u_{it}, has the one-factor structure

$$u_{it} = g_i f_t + e_{it} \qquad (47)$$

in which f_t is the unobserved common effect, and e_{it} is the individual-specific (idiosyncratic) error.

It is convenient to write (46) and (47)

$$\Delta y_{it} = a_i + b_i y_{i,t-1} + g_i f_t + e_{it} \qquad (48)$$

where $a_i = (1 - j_i)m_i$, $b_\mathrm{i} = -(1 - j_i)$ and $\Delta \mathrm{y}_{\mathrm{it}} = y_{it} - y_{i,t-1}$. The unit root hypothesis of interest, $j_i = 1$, can now be expressed as

$$\mathrm{H}_0 : b_\mathrm{i} = 0 \text{ for all } i, \qquad (49)$$

against the possibly heterogeneous alternatives

$$\mathrm{H}_1 : b_i < 0,\ i = 1, 2, ..., \mathrm{N}_1, \quad b_\mathrm{i} = 0, \quad i = \mathrm{N}_1 + 1, \mathrm{N}_1 + 2, ..., \mathrm{N} \tag{50}$$

A necessary assumption is that N_1/N, the fraction of the individual processes that are stationary, is non-zero and tends to the fixed value δ such that $0 < \delta \leq 1$ as $N \to \infty$. As noted in Im, Pesaran and Shin (2003) this condition is necessary for the consistency of the panel unit root tests.

The following assumptions are also made:

Assumption 1: The idiosyncratic shocks, e_{it}, i = 1, 2, ...,N, t = 1, 2, ..., T, are independently distributed both across i and t, have mean zero, variance s_i^2, and finite fourth-order moment.

Assumption 2: The common factor, f_t, is serially uncorrelated with mean zero and a constant variance, s_f^2, and finite fourth-order moment. Without loss of generality s_f^2 will be set equal to unity.

Assumption 3: e_{it}, f_t and g_i are independently distributed for all i. The cross-section independence of e_{it} (across i) is standard in one factor models, although its validity in more general settings may require specification of more than one common factor in (47). Assumptions 1 and 2 together imply that the composite error, u_{it}, is serially uncorrelated. This restriction can be relaxed by considering stationary error processes of the type

$$u_{it} = \sum_{j=1}^{p} r_{ij} u_{i,t-j} + g_i f_t + e_{it}$$

B. CIPS Panel Unit Root Test

Given that the null distribution of the individual CADF statistics are asymptotically independent of the nuisance parameters, the various panel unit root tests developed in the literature for the case of the cross-sectionally independent errors can also be applied to the present more general case. Here we focus on a generalization of the t-bar test proposed by IPS and consider a cross-sectionally augmented version of the IPS test based on

$$CIPS(N,T) = N^{-1}\sum_{i=1}^{N} t_i(N,T) \tag{51}$$

where $t_i\ (N,T)$ is the cross-sectionally augmented Dickey-Fuller statistic for the i^{th} cross section unit given by the *t*-ratio of the coefficient of $y_{i,t-1}$ in the CADF regression.

The mean deviations are considered as:

$$D(N,T) = N^{-1}\sum_{i=1}^{N}[t_i(N,T) - CADF_{if}\,],$$

where $CADF_{if}$ is the stochastic limit of $t_i(N, T)$ as N and T tend to infinity such that $N/T \rightarrow k\ (0 < k < \infty)$. It seems reasonable to expect that $D(N, T) = o_p(1)$ for N and T sufficiently large. This conjecture would clearly hold in the case where $\bar{t}_i\,(N,T)$ have finite moments for all N and T above some given threshold values, say N_0, and T_0. However, such moment conditions are difficult to establish even under cross-section independence.

One possible method of dealing with these technical difficulties would be to base the t-bar test on a suitably truncated version of the CADF statistics. After conducting simulations it is suggested that the standardized version of these statistic are very close to being standard Normal with finite first and second order moments. Therefore, for the purpose of the panel unit root test it would be equally valid to base the test on an average of the truncated versions of $t_i(N, T)$, say $t_i^*(N,T)$, where

$$\left\{\begin{array}{l} t_i^*(N,T) = t_i\ (N,T), \text{if } -\mathrm{K}_1 < t_i\ (N,T) < \mathrm{K}_2 \\ t_i^*(N,T) = -\mathrm{K}_1, \text{if } t_i\ (N,T) < -\mathrm{K}_1 \\ t_i^*(N,T) = \mathrm{K}_2, \text{if } t_i(N,T) > \mathrm{K}_2 \end{array}\right\} \tag{52}$$

where K_1 and K_2 are positive constants that are sufficiently large so that $P_r\ [-K_1 < t_i(N, T) < K_2]$ is sufficiently large, say in excess of 0.9999. Using the normal approximation of t_i(N, T) as a crude benchmark we would have

$$K_1 = -E(CADF_{if}) - \Phi^{-1}(e/2)\sqrt{Var(CADF_{if})},$$

and

$$K_2 = E(CADF_{if}) + \Phi^{-1}(1-e/2)\sqrt{Var(CADF_{if})},$$

where ε is a sufficiently small positive constant. For example, setting $e = 1 \text{x} 10^{-6}$ in the case of models without an intercept or trend we would have K_1=0.98+4.8917(1.05) = 6.12, and K_2 = -0.98 +4.8917(1.05) = 4.16. Similarly, for models with an intercept we have K_1 = 6. 19, and K_2 = 2.61, and finally for models with a linear trend we obtain K_1 = 6.42, and K_2 = 1.70

The associated truncated panel unit root test is now given by

$$CIPS^*(N,T) = N^{-1}\sum_{i=1}^{N} t_i^*(N,T) \tag{53}$$

Since, by construction all moments of $t_i\ (N,T)$ exist it then follows that

$$CIPS^*(N,T) = N^{-1}\sum_{i=1}^{N} CADF_{if}^* + o_p(1) \tag{54}$$

where $CADF_{if}^*$ is given by:

$$\left\{ \begin{array}{l} CADF_{if}^* = CADF_{if}, \text{ if } -K_1 < CADF_{if} < K_2 \\ CADF_{if}^* = -K_1, \text{ if } CADF_{if} < -K_1 \\ CADF_{if}^* = K_2, \text{ if } CADF_{if} > K_2 \end{array} \right\} \tag{55}$$

The distributions of the average CADF statistic or its truncated counterpart, $\overline{CADF^*} = N^{-1}\sum_{i=1}^{N} CADF_{if}^*$, are non-standard even for sufficiently large N. This is due to the dependence of the individual $CADF_{if}$ variates on the common process W_f which invalidates the application of the standard central limit theorems to $\overline{CADF}$ or $\overline{CADF^*}$, and is in contrast to the results obtained by IPS under cross-section independence where a standardized version of $\overline{CADF} = N^{-1}\sum_{i=1}^{N} CADF_{if}$, was shown to be normally distributed for N sufficiently large. Nevertheless, it is possible to show that. $\overline{CADF}*$ converges in distribution as N →∞ without any need for further normalization.

The above results establish that the $\overline{CADF}*$ converges almost surely to a distribution which depends on K_1,K_2 and W_f . This distribution does not seem analytically tractable, but can be readily simulated. By simulating the distribution of CIPS* setting N = 100, T = 500, and using 50,000 replications under following cases:

1. Models without intercepts or trends (I), with K_1 = 6.12, and K_2 = 4.16,

2. Models with intercept only (II), with K_1 = 6. 19, and K_2 = 2.61,

3. Models with a linear trend (III), with K_1 = 6.42, and K_2 = 1.70.

The 1%, 5% and 10% critical values of CADF. and CADF are given in Tables 3a-3c, in the appendix, for models I-III, respectively. In most cases the critical values for the two versions of the CIPS test are identical and only one value is reported. In cases where the two critical values differ the truncated version is included in brackets.

C. Empirical Results of CIPS test

Our empirical analysis ends by conducting a third panel test for the quarterly and monthly data of the 20 industrialized countries employed in this research. The model chosen to represent the nature of data is that of intercept and no trend. It is a new test proposed by Pesaran (September 2003) in order to overcome the problem of cross section dependence. The results for the panels constructed for the purpose of this analysis are unanimous in the rejection of the unit root null. The test rejects at 1% significance level the null hypothesis of non stationarity of the real exchange rates for all the groups under consideration.

This improved and more complete version of IPS panel unit root test provides absolute support to the PPP hypothesis. By this outcome we come to the conclusion that by taking into consideration the dependence that exists across the countries of a panel the evidence in favor of PPP in strengthened. So, CIPS is a more powerful panel test that provides more support for PPP not only from univariate tests but even from other panel tests.

Table 9
Panel Unit Root tests: Results from CIPS test[3]

	Monthly		Quarterly	
Panels	**t-stat**	**result**	**t-stat**	**result**
All 18	-2.775***	Reject unit root null-Accept PPP		
All 20			-2.602***	Reject unit root null-Accept PPP
EC (15)	-2.562***	Reject unit root null-Accept PPP	-2.451***	Reject unit root null-Accept PPP
EMU (12)	-2.383**	Reject unit root null-Accept PPP	-2.489**	Reject unit root null-Accept PPP
G6	-2.717***	Reject unit root null-Accept PPP	-2.504**	Reject unit root null-Accept PPP
G10	-2.790***	Reject unit root null-Accept PPP	-2.847***	Reject unit root null-Accept PPP
OECD (13)	-2.801***	Reject unit root null-Accept PPP	-2.653***	Reject unit root null-Accept PPP

***,**,* denote rejection of the unit root null with statistical significance at 1%, 5%, 10% respectively.

The critical values are presented in tables 3a, 3b, 3c of the appendixes

$$\Delta y_{it} = a_i + b_i y_{i,t-1} + g_i f_t + e_{it}$$

$$H_0 : b_i = 0 \text{ for all } i,$$

$$H_1 : b_i < 0, \; i = 1, 2, ..., N_1, \quad b_i = 0, \quad i = N_1 + 1, N_1 + 2, ..., N$$

[3] The countries that are included in each panel are presented analytically in section 6

IV) Comparison of Panel tests

We start the comparison of the three very important panel unit root tests analyzed above by stating their similarities. We would like to underline that there are several common features shared by the previously proposed tests. First, they all require that the number of groups be infinite. Without this condition, asymptotic normality of the tests does not hold. Moreover, for these tests as $N, T \to \infty$, we need to assume either $N/T \to 0$ as in LL and IPS, where N denotes the number of groups and T the number of time series common to all the groups. These assumptions on the number of groups imply that it should be infinite for the asymptotic normality results to hold and, at the same time, small enough relative to the number of time series. A practical implication of this requirement is that the tests may not keep nominal size well either when N is small or when N is large relative to T. Empirical tests of IPS show that their and LL's tests have more size distortions as N becomes large relative to T. But in practice, it is likely either that N is small or that N is large relative to T, which indicates the need for unit root tests that do not require such assumptions on N and T.

Second, the previous tests assume the same number of time series for all the groups. However, depending on circumstances, it is quite likely that time series spans for the groups are different. Though IPS considered such case in passing, the required moment calculations for the IPS tests make it difficult to use them.

Apart from the common features of the tests there are also some basic differences that should be mentioned. Starting with the differences in the basic hypothesis of the tests we must point out that the LL test imposes an identical first order autoregressive coefficient on all series in the panel and rejection of the null hypothesis implies that real exchange rates in all economies adjust at the same rate, while the IPS and CIPS tests allows for heterogeneous first order coefficients so that real exchange rates may adjust at different rates. This is indicated by the fact that Im , Pesaran and Shin in their testing framework allow for β_i to differ across groups but on the other side Levin and Lin test a homogeneous alternative hypothesis, namely $\beta_i = \beta < 0$ for all i.

Second very important difference of the tests I that they have different null and alternative. For LL the alternative is that none of the countries have a unit root. A case where some countries have a unit root and the others do not is not dealt with by this test while on the other hand this is the alternative of IPS and CIPS tests.

As a consequence of the above difference it is obvious that the IPS CIPS tests are less restrictive allowing heterogeneity in the parameters under study. And in the case of serial correlation and heterogeneity in the underlying data generating process, the simulation results of Im, Pesaran and Shin show that if a large lag order is selected for the ADF regressions, then the finite sample performance of the t-bar test is better than that of the Levin and Lin test.

Fourth, the LL test assumes that all the groups have the same type of non stochastic component. This means that for LL test if one group is specified to have a linear time trend, the other groups are automatically assumed to have a linear time trend. On the contrary for IPS and CIPS tests specifying a different type of non stochastic component for each group is allowed.

Finally, only the CIPS test takes into consideration the cross sectional dependence. The dependence that exists between the countries that form a panel is not computed in the cases of LL and IPS test. This is a very important omission of the other tests and makes CIPS test more complete and accurate.

-Comparison of Empirical Results

The three panel tests employed in this research with the similarities and differences in the way they work provide results that should be seen and criticized with respect to one another. Tables 10 and 11 concentrate all the results and give an overall view of what are the conclusions of this research about the stationarity of real exchange rates in the floating exchange rate period for the sample of countries examined.

The findings from the test of Levin and Lin (LL) are different from those obtained from the IPS and CIPS tests. In Particular, when we used the LL test we rejected the unit root null in all of the group of countries at 1% significance level providing absolute support to the PPP hypothesis. While for IPS test, the unit root null is not rejected for the panel consisted of the 6 most industrialized countries (G6) giving support to the PPP hypothesis in six out of the seven panels. This difference between the results of the tests be may be due to the less restrictive approach of Im, Pesaran and Sin which allows for heterogeneity in the tested parameters while the restrictive homogeneous hypothesis of LL test advocates in favor or the PPP hypothesis. Finally, the CIPS test supports the rejection of the unit root hypothesis for all the panels examined for both monthly and quarterly data. These results are similar to those derived from the LL test but for different reason. In this test the inclusion of cross sectional dependence in the calculation of the results lead to an outcome more supportive for the PPP hypothesis in all panels.

Table 10
Comparison of Panel Unit root tests: Monthly data for US Dollar real exchange rates

	LL test		IPS test	CIPS test
Panels	**t_δ**	**modified t-stat**	**w-stat**	**t-stat**
All 18	-6.580	-8.498***	-3.157***	-2.775***
All 20				
EC (15)	-5.932	-7.663***	-3.237***	-2.562***
EMU (12)	-5.2807	-6.8236***	-2.914***	-2.383**
G6	-4.450	-5.748***	-1.632	-2.717***
G10	-5.737	-7.403***	-2.003**	-2.790***
OECD (13)	-7.228	-9.324***	-2.641***	-2.801***

***,**,* denote rejection of the unit root null with statistical significance at 1%, 5%, 10% respectively.

Table 11

Comparison of Panel Unit root tests: Quarterly data for US Dollar real exchange rates

Panels	**LL test** t_δ	**LL test** modified t-stat	**IPS test** w-stat	**CIPS test** t-stat
All 18				
All 20	-8.116	-10.434***	-3.38237***	-2.602***
EC (15)	-6.053	-7.787***	-2.872***	-2.451***
EMU (12)	-6.009	-7.728***	-2.621***	-2.489**
G6	-5.001	-6.417***	-1.619	-2.504**
G10	-6.251	-8.001***	-1.922*	-2.847***
OECD (13)	-7.401	-9.463***	-2.425**	-2.653***

***,**,* denote rejection of the unit root null with statistical significance at 1%, 5%, 10% respectively.

6. THE SAMPLE AND DATA DESCRIPTION

The ‘current’ period of flexible exchange rates amongst the major currencies, since the abandonment of Bretton-Woods Systems in 1971, is now almost 30 years’ old. The purpose of this paper is to evaluate the evidence of long-run purchasing power parity during this period. For this reason will use the newest econometric techniques in testing PPP, as presented above.

The data needed for this research are the quarterly and monthly observations of nominal exchange rates and aggregate Consumer Price Index ratio (CPI) for the 21 countries which the IMF classify as being industrialized. The validity of PPP hypothesis will be tested for the period 1973-2003 with U.S. Dollar as the base currency. For the countries that have switched their domestic currencies to the Euro, we will collect the nominal exchange rate currency by the U.S. Dollar from 1973 to 1998, then the Euro by U.S. Dollar exchange rate from 1999 to 2003 and convert in the currency by U.S. Dollar using the prefixed exchange rates. All the data will be taken from the IMF’s *International Financial Statistics* (IFS) and other databases such as DataStream.

The data for Belgium and Luxembourg are considered together due to their previous union, also there no monthly data available for Australia and New Zealand so these two countries are included only in the quarterly results.

The data is grouped under several panels for conducting the panel unit root tests. The panels that we used in this study are:

- ***All 20 industrialised countries.***

This panel includes:

Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, Netherlands, Norway, New Zealand, Portugal, Spain, Sweden, Switzerland, and the U.K.

- ***the European Community (EC),***

EC includes the 15 countries that are members or the European Union:
Austria, Belgium + Luxemburg, Denmark, Finland, France, Germany, Greece, Ireland, Italy, and the Netherlands, Portugal, Spain, Sweden, UK .

- ***the European Monetary Union (EMU),***

EMU includes the 12 of the 15 state members of the EU that have adopted the Euro are their national currency unit[4]:
Austria, Belgium + Luxemburg, Finland, France, Germany, Greece, Ireland, Italy, and the Netherlands, Portugal, Spain.

- ***the 6 and 10 most industrialized countries (G6, G10),***

G6 includes:
Canada, France, Germany, Italy, Japan, and the U.K.

For G10:
Belgium, the Netherlands, Sweden, and Switzerland are added.

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[4] The EMU is an area with a single currency into the single market in the European Union. In the European Council of Brussels in 1998 it was decided that 11 of the 15 states members of the European Union held the necessary requirements for he adoption of the euro, deciding Denmark, the United Kingdom, Greece and Sweden to stay out of that area voluntarily , though Greece joined the eurozone on the 1st of January 2001. Sweden's situation will be revised each 2 years or even more frequently if Sweden asks it so. Denmark and the United Kingdom will not have this systematic revisions but they will be able to reconsider in any moment their decision of no participating in the euro zone.

- ***and the OECD countries (13),***

OECD (13) includes 13 older members of the 30 member countries of the Organisation for Economic Co-operation and Development :
Australia, Belgium, Canada, Denmark, Finland, France, Germany, Italy, the Netherlands, Norway, Sweden, and the U.K.

7. CONCLUSIONS

The 'current' period of flexible exchange rates among the world's major currencies is now almost 30 years old. The purpose of this paper is to examine the evidence of PPP during this period and to determine what can be learned about real exchange rates with the United States dollar as the numeraire currency by using both univariate and panel unit root tests for PPP. We proceed in two stages. First, using econometric techniques, we document the evidence of PPP with univariate tests as the span of the data is increased from 1973-1988 to 1973-2003. Second, we perform three panel unit root tests for seven different panels during the post Bretton-Woods period of flexible exchange rates. Specifically, we conduct LL test, IPS test and a new panel test proposed by Pesaran (September 2003), CIPS test. This new panel test is the innovation of this research since it is the first time that it is used for testing the validity of the PPP hypothesis.

From our empirical results we come to the conclusion that the panel framework of unit root tests provides greater power than do single time series unit root tests. Evidence from univariate unit root test show little support to the PPP hypothesis even when the sample of data doubles from 1988 to 2003. On the other hand, results from the panel unit root tests generally support PPP for the seven panel of countries constructed .

These results indicate that the post Bretton-Woods period of flexible exchange rates, 1973 to 2003, provides an almost ideal setting for panel unit root tests of purchasing power parity for industrialized countries. While the span of the data is too short for univariate tests to have good power, it is long enough for panel methods. The development of panel unit root tests presents both a challenge and an opportunity for researchers attempting to find strong evidence of long-run purchasing power parity using data from the current float. The opportunity occurs because, in contrast with univariate methods, panel unit root tests have sufficient power to reject the unit root null. The challenge arises because, again in contrast with univariate methods, failure to reject the unit root null in real exchange rates can no longer be ascribed to low power of the tests.

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Appendix

Table 3a

Critical Values of Average of Individual Cross-Sectionally Augmented Dickey-Fuller Distribution (Case I: No Intercept and No Trend)[5],[6],[7]

-

1% ***(CADF)***

T/N	10	15	20	30	50	70	100	200
10	-2.16	-2.02	-1.93	-1.85	-1.78	-1.74	-1.71	-1.70
	(-2.14)	(-2.00)	(-1.91)	(-1.84)	(-1. 77)	(-1.73)		(-1.69)
15	-2.03	-1.91	-1.84	-1.77	-1.71	-1.68	-1.66	-1.63
20	-2.00	-1.89	-1.83	-1.76	-1. 70	-1.67	-1.65	-1.62
30	-1.98	-1.87	-1.80	-1.74	-1.69	-1.67	-1.64	-1.61
50	-1.97	-1.86	-1.80	-1.74	-1.69	-1.66	-1.63	-1.61
70	-1.95	-1.86	-1.80	-1.74	-1.68	-1.66	-1.63	-1.61
100	-1.94	-1.85	-1.79	-1.74	-1.68	-1.65	-1.63	-1.61
200	-1.95	-1.85	-1.79	-1.73	-1.68	-1.65	-1.63	-1.61

5% ***(CADF)***

T/N	10	15	20	30	50	70	100	200
10	-1.80	-1.71	-1.67	-1.61	-1.58	-1.56	-1.54	-1.53
	(-1.79)		(-1.66)		(-1.57)	(-1.55)	(-1.53)	(-1.52)
15	-1.74	-1.67	-1.63	-1.58	-1.55	-1.53	-1.52	-1.51
20	-1.72	-1.65	-1.62	-1.58	-1.54	-1.53	-1.52	-1.50
30	-1.72	-1.65	-1.61	-1.57	-1.55	-1.54	-1.52	-1.50
50	-1.72	-1.64	-1.61	-1.57	-1.54	-1.53	-1.52	-1.51
70	-1.71	-1.65	-1.61	-1.57	-1.54	-1.53	-1.52	-1.51
100	-1.71	-1.64	-1.61	-1.57	-1.54	-1.53	-1.52	-1.51
200	-1.71	-1.65	-1.61	-1.57	-1.54	-1.53	-1.52	-1.51

10% ***(CADF)***

T/N	10	15	20	30	50	70	100	200
10	-1.61	-1.56	-1.52	-1.49	-1.46	-1.45	-1.44	-1.43
		(-1.55)		(-1.48)			(-1.43)	
15	-1.58	-1.53	-1.50	-1.48	-1.45	-1.44	-1.44	-1.43
20	-1.58	-1.52	-1.50	-1.47	-1.45	-1.45	-1.44	-1.43
30	-1.57	-1.53	-1.50	-1.47	-1. 46	-1.45	-1.44	-1.43
50	-1.58	-1.52	-1.50	-1.47	-1.45	-1.45	-1.44	-1.43
70	-1.57	-1.52	-1.50	-1.47	-1.46	-1.45	-1.44	-1.43
100	-1.56	-1.52	-1.50	-1.48	-1.46	-1.45	-1.44	-1.43
200	-1.57	-1.53	-1.50	-1.47	-1.45	-1.45	-1.44	-1.43

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5 $\overline{CADF}$ statistic is computed as the simple average of the individual-specific $CADF_i$ statistics.

6 The calculations are carried out for 50,000 replications based on the OLS regression of Δy_{it} on $y_{i,t-1}$, and Δy_t. The CADFi refers to the OLS t-ratio of the coefficient of $y_{i,t-1}$.

7 The critical values for the truncated version of the test statistics are indicated in brackets if they differ from the non-truncated ones.

Table 3b
Critical Values of Average of Individual Cross-Sectionally Augmented Dickey-Fuller Distribution (Case II: Intercept only)[8],[9],[10]

-

1% *(CADF)*

TIN	**10**	**15**	**20**	**30**	**50**	**70**	**100**	**200**
10	-2.97	-2.76	-2.64	-2.51	-2.41	-2.37	-2.33	-2.28
	(-2.85)	(-2.66)	(-2.56)	(-2.44)	(-2.36)	(-2.32)	(-2.29)	(-2.25)
15	-2.66	-2.52	-2.45	-2.34	-2.26	-2.23	-2.19	-2.16
20	-2.60	-2.47	-2.40	-2.32	-2.25	-2.20	-2.18	-2.14
30	-2.57	-2.45	-2.38	-2.30	-2.23	-2.19	-2.17	-2.14
50	-2.55	-2.44	-2.36	-2.30	-2.23	-2.20	-2.17	-2.14
70	-2.54	-2.43	-2.36	-2.30	-2.23	-2.20	-2.17	-2.14
100	-2.53	-2.42	-2.36	-2.30	-2.23	-2.20	-2.18	-2.15
200	-2.53	-2.43	-2.36	-2.30	-2.23	-2.21	-2.18	-2.15

5% *(CADF)*

TIN	**10**	**15**	**20**	**30**	**50**	**70**	**100**	**200**
10	-2.52	-2.40	-2.33	-2.25	-2.19	-2.16	-2.14	-2.1 0
	(-2.47)	(-2.35)	(-2.29)	(-2.22)	(-2.16)	(-2.13)	(-2.11)	(-2.08)
15	-2.37	-2.28	-2.22	-2.17	-2.11	-2.09	-2.07	-2.04
20	-2.34	-2.26	-2.21	-2.15	-2.11	-2.08	-2.07	-2.04
30	-2.33	-2.25	-2.20	-2.15	-2.11	-2.08	-2.07	-2.05
50	-2.33	-2.25	-2.20	-2.16	-2.11	-2.10	-2.08	-2.06
70	-2.33	-2.25	-2.20	-2.15	-2.12	-2.1 0	-2.08	-2.06
100	-2.32	-2.25	-2.20	-2.16	-2.12	-2.1 0	-2.08	-2.07
200	-2.32	-2.25	-2.20	-2.16	-2.12	-2.10	-2.08	-2.07

-

10% *(CADF)*

TIN	**10**	**15**	**20**	**30**	**50**	**70**	**100**	**200**
10	-2.31	-2.22	-2.18	-2.12	-2.07	-2.05	-2.03	-2.01
	(-2.28)	(-2.20)	(-2.15)	(-2.1 0)	(-2.05)	(-2.03)	(-2.01)	(-1.99)
15	-2.22	-2.16	-2.11	-2.07	-2.03	-2.01	-2.00	-1.98
20	-2.21	-2.14	-2.10	-2.07	-2.03	-2.01	-2.00	-1.99
30	-2.21	-2.14	-2.11	-2.07	-2.04	-2.02	-2.01	-2.00
50	-2.21	-2.14	-2.11	-2.08	-2.05	-2.03	-2.02	-2.01
70	-2.21	-2.15	-2.11	-2.08	-2.05	-2.03	-2.02	-2.01
100	-2.21	-2.15	-2.11	-2.08	-2.05	-2.03	-2.03	-2.02
200	-2.21	-2.15	-2.11	-2.08	-2.05	-2.04	-2.03	-2.02

[8] $\overline{CADF}$ statistic is computed as the simple average of the individual-specific $CADF_i$ statistics.

[9] The calculations are carried out for 50,000 replications based on the OLS regression of Δy_{it} on an intercept $y_{i,t-1}$, and Δy_t . The CADFi refers to the OLS t-ratio of the coefficient of $y_{i,t-1}$.

[10] The critical values for the truncated version of the test statistics are indicated in brackets if they differ from the non-truncated ones

Table 3c
Critical Values of Average of Individual Cross-Sectionally Augmented Dickey-Fuller Distribution (Case III: Intercept and Trend)[11] [12][13]

1% *(CADF)*

TIN	10	15	20	30	50	70	100	200
10	-3.88	-3.61	-3.46	-3.30	-3.15	-3.10	-3.05	-2.98
	(-3.51)	(-3.31)	(-3.20)	(-3.10)	(-3.00)	(-2.96)	(-2.93)	(-2.88)
15	-3.24	-3.09	-3.00	-2.89	-2.81	-2.77	-2.74	-2.71
	(-3.21)	(-3.07)	(-2.98)	(-2.88)	(-2.80)	(-2.76)		(-2.70)
20	-3.15	-3.01	-2.92	-2.83	-2.76	-2.72	-2.70	-2.65
30	-3.1 0	-2.96	-2.88	-2.81	-2.73	-2.69	-2.66	-2.63
50	-3.06	-2.93	-2.85	-2.78	-2.72	-2.68	-2.65	-2.62
70	-3.04	-2.93	-2.85	-2.78	-2.71	-2.68	-2.65	-2.62
100	-3.03	-2.92	-2.85	-2.77	-2.71	-2.68	-2.65	-2.62
200	-3.03	-2.91	-2.85	-2.77	-2.71	-2.67	-2.65	-2.62

5% *(CADF)*

TIN	10	15	20	30	50	70	100	200
10	-3.27	-3.11	-3.02	-2.94	-2.86	-2.82	-2.79	-2.75
	(-3.1 0)	(-2.97)	(-2.89)	(-2.82)	(-2.75)	(-2.73)	(-2.70)	(-2.67)
15	-2.93	-2.83	-2.77	-2.70	-2.64	-2.62	-2.60	-2.57
	(-2.92)	(-2.82)	(-2.76)	(-2.69)			(-2.59)	
20	-2.88	-2.78	-2.73	-2.67	-2.62	-2.59	-2.57	-2.55
30	-2.86	-2.76	-2.72	-2.66	-2.61	-2.58	-2.56	-2.54
50	-2.84	-2.76	-2.71	-2.65	-2.60	-2.58	-2.56	-2.54
70	-2.83	-2.76	-2.70	-2.65	-2.61	-2.58	-2.57	-2.54
100	-2.83	-2.75	-2.70	-2.65	-2.61	-2.59	-2.56	-2.55
200	-2.83	-2.75	-2.70	-2.65	-2.61	-2.59	-2.57	-2.55

10% *(CADF)*

TIN	10	15	20	30	50	70	100	200
10	-2.98	-2.89	-2.82	-2.76	-2.71	-2.68	-2.66	-2.63
	(-2.87)	(-2.78)	(-2.73)	(-2.67)	(-2.63)	(-2.60)	(-2.58)	(-2.56)
15	-2.76	-2.69	-2.65	-2.60	-2.56	-2.54	-2.52	-2.50
		(-2.68)	(-2.64)	(-2.59)	(-2.55)	(-2.53)	(-2.51)	
20	-2.74	-2.67	-2.63	-2.58	-2.54	-2.53	-2.51	-2.49
30	-2.73	-2.66	-2.63	-2.58	-2.54	-2.52	-2.51	-2.49
50	-2.73	-2.66	-2.63	-2.58	-2.55	-2.53	-2.51	-2.50
70	-2.72	-2.66	-2.62	-2.58	-2.55	-2.53	-2.52	-2.50
100	-2.72	-2.66	-2.63	-2.59	-2.55	-2.53	-2.52	-2.50
200	-2.73	-2.66	-2.63	-2.59	-2.55	-2.54	-2.52	-2.51

[11] $\overline{CADF}$ statistic is computed as the simple average of the individual-specific $CADF_i$ statistics.

[12] The calculations are carried out for 50,000 replications based on the OLS regression of Δy_{it} on an intercept, trend $y_{i,t-1}$, and Δy_t . The CADFi refers to the OLS t-ratio of the coefficient of $y_{i,t-1}$.

[13] The critical values for the truncated version of the test statistics are indicated in brackets if they differ from the non-truncated ones.