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BEFORE THE FALL WAS THE TURKISH LIRA OVERVALUED?

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Abstract

This paper examines validity of purchasing power parity to evaluate whether the Turkish Lira was overvalued on the eve of 2001 crises. Univariate and multivariate time series techniques are used to test whether the real exchange rate is mean reverting. Half-life of deviation from purchasing power parity for various definitions is derived. The Johansen cointegration test procedure is applied to bilateral exchange rates using CPI and WPI price indexes. Finally, different measures of misalignments are calculated. Evidence provided that calculated half lives are, in general, short compared to low inflation countries. Further, data supports the long-run relationships among exchange rates, domestic and foreign prices. Calculated misalignments give mixed results on bilateral exchange rates based on CPI and WPI. While WPI based bilateral real exchange rate gives under valuation, the CPI based bilateral real exchange rate and trade weighted real exchange rates based on WPI reveals that TL was overvalued before the eve of 2001 crises.

1. Introduction

The real exchange rate is often considered to be indicative of international competitiveness and is used as a guide to monetary and exchange rate policies. However, in a highly competitive world the structure and directions of trade adapt to exchange rate changes in a complex way which requires detailed understanding of the behaviour of exchange rates. There is also a growing agreement that prolonged and substantial real exchange rate misalignment can create severe macroeconomic disequilibria and that the correction of external balance will, in general, require both real exchange rate devaluation and demand management policies. Developments in 1990s and 2000s show that the cost associated with real exchange rate misalignment is very high. Turkish and Mexican currency crises in 1994, Asian crisis in 1997, Brazilian crises in 1999, Turkish crises in 2000 and 2001 and Argentinean crises in 2002 have served as a reminder of the macroeconomic disruption that can be caused by real exchange rate misalignment.

In addition, a number of papers have pointed to exchange rate misalignment as a robust empirical determinant of currency crises (Frankel and Rose, (1996); Sachs et al., (1996); Kaminsky, Lizondo and Reinhart (1997); Kaminsky and Reinhart (1999); and Goldfajn and Valdes (1999)). Hence, the presence of misalignment is potentially important for policy purposes because of its role as a component of an early warning system (see Berg et al., (2000)).

However, it is not easy to set nominal and real exchange rate in their intended path. There are number of issues to be confronted. There is a conceptual discrepancy as to what exactly is meant by the long-run equilibrium exchange rate. There is also an empirical issue regarding what the value of long-run equilibrium rate is for a given country at any moment in time. In the literature, there are at least three broad definitions of misalignment (see Williamson (1994), Miles-Feretti and Razin (1996), and Hinkle and Monteil (1999)). The first one is price based criteria, such as purchasing power parity and its variants; the second one is model based criteria, based on a formal model of nominal exchange rates; and the third one is solvency and sustainability based criteria, which make reference to trends in the current account and the external debt to GDP ratio. Implementing model based and sustainability-based criteria require more detailed analysis. Price based criterion is relatively easy to implement and has strong operational advantages, but does not address the economically interesting question of whether a particular exchange rate is at an optimal level. On the other hand, the sustainability measures can make reference to an optimal level, but are very difficult to calculate as they require a fully-fleshed out macroeconomic model. PPP based analysis can be used to make initial diagnoses and for identifying hypotheses for analysing more detailed models. Therefore, in this paper a more modest goal of implementing the price based criteria is set forth.

The main objective of this study is to evaluate whether PPP is a valid criteria for calculating misalignment in Turkey for 1987:1-2000:12. Given this objective we try to provide answers to the following questions: are the various definitions of real exchange rates mean reverting or unit root non-stationary? Can we draw conclusions about the relevance of PPP from calculated half-life deviation from parity? Since the PPP implies cointegration between the nominal exchange rate, domestic price level and foreign price level do the cointegration results provide further evidence on the validity of our chosen approach? Finally, if the approach is valid was the TL overvalued before the 2001 crises?

The paper proceeds in the following manner. In Section 2, the price-based measures are described. In section 3, data, and their time series properties are evaluated and half-lives of various definitions of real exchange rates are calculated. In section 4 the tests for purchasing power parity is undertaken by using the Johansen (1995) technique. Section 5 discusses various estimates of the equilibrium real exchange rate and calculates misalignments. Finally, section 6 concludes.

2. Price Based Measures of Equilibrium Real Exchange Rates

Purchasing Power Parity (PPP) theory of exchange rate determination asserts that the exchange rate between two currencies over any period of time is determined by the change in the two countries price levels. This theory singles out changes in price levels as the overriding determinant in the determination of exchange rate.

According to this theory, exchange rate in the short run would diverge from PPP. There can be many reasons why deviations from PPP occur. Firstly, there may be restrictions on trade and capital movements or transfer pricing in a country, which will distort the relationship between home and foreign prices. Secondly, speculative activities and official intervention may create a PPP disparity. Lastly, the productivity bias when there is a relatively faster growing productivity growth in the tradable sector than the non-tradable sector will result in systematic divergence of internal prices (Balassa (1964) and Chinn (2000)). The basic concept underlying PPP is that arbitrage forces will equalise prices of goods internationally if they are measured in the same currency.

What can be tested is how well purchasing power parity PPP holds up to a constant, α,

$$s_t = \delta_1 p_t - \delta_2 p_t^* + \alpha + \varepsilon_t \tag{1}$$

where s is the log nominal exchange rates, p and p^* are domestic and foreign prices respectively and ε is stationary random variable. If we assume that δ_l=-δ_2=1, this specification implies that real exchange rate rs is given by

$$rs_t = s_t - p_t + p_t^* = \alpha \tag{2}$$

and equilibrium real exchange rate is a constant and equals to α. In the above specifications there are no fundamental variables. Fundamentals are subsumed into constant term and their random components into the error term.

Long-run PPP posits a stable long-run relationship between nominal exchange rates and relative price levels. A number of early studies on industrial nations find little empirical support for such a relationship using data from the modern float Froot and Rogoff (1995). However, recent studies using long span of data and/or panel data find support for long-run PPP for the post-Bretton Woods era Frankel and Rose (1996), Papell (1997), Taylor and Sarno (1998), Glen (1992), Lothian and Taylor (1996, 2000), and Taylor (2002). Even in relatively short span of data mean reversion is sometimes identified, especially in countries with experience of high inflation see Breuer (1994).

A number of other studies on the validity of the PPP using Turkish data at best give mixed results. Akinci and Demir (1993) find no support for PPP by using monthly data for 1982:1-1992:12 periods and Engle-Granger two-step cointegration approach. Metin (1994) uses annual data for the period 1948-1988, which can be characterized as a fixed exchange rate system and cointegration method and finds no support for PPP. Taskin and Metin (1994) find no support for PPP by using cointegration technique and monthly data for the period 1981-1993. Further, a study by Telatar and Kazdaglı use Engle-Granger two-step approach and monthly data for the period 1980:10-1993:10. They also find no support for PPP. However, Sarno (2000) finds support for the PPP by using the exponential smooth transition autoregressive model and the same data set as Telatar and Kazdaglı (1998). The most recent study by Erlat (2001) uses sequential unit root test with single and double shift in constant and trend. Erlat also used fractional integration techniques and finds that both CPI and WPI based real exchange rate series and provide empirical support for PPP using Turkish monthly data for 1984:1-2000:9 periods. Studies using cointegration techniques and data before 1994 find no support for PPP. Our study also used cointegration techniques but used extended data for the 1987:1-2000:12 period and finds evidence of cointegration, and thus PPP.

3. Data, Unit Roots and Half-Life of Real Exchange Rate

3.1. Data

Time series used in this paper are from the *Monthly Bulletin* of Central Bank of Turkey and IMF's *International Financial Statistic*s, and spans the 1987:1-2000:12 period. The exchange rate is average-of-month data, expressed in TL per US dollar unit. For the broad deflator, the CPI *IFS* line 64 is used. The 'tradable' price deflator is proxied by the WPI data reported in *IFS* line 63.

Price levels should represent tradable goods and services bundles, and should be similar across countries. Since consumer bundles might be more similar across countries than producer or wholesale bundles, consumer price indices (CPIs) may provide a more consistent measure of price levels and thus of real exchange rates. However, we are cynical on this issue, and use wholesale price index (WPI) as well.

In principle, one might like to use trade weighted measures of the real exchange rate. The problem that one encounters is that the patterns of trade flows change substantially over the sample period and hence so too do the appropriate trade-weights. Nonetheless, the trade weighted real exchange rate (TWRER) calculated by the Central Bank is also used in calculating misalignment[1].

3.2. Unit Root Tests

In this section the long-run mean reversion properties of the real exchange rate series and their components are investigated. Tests on nominal exchange rates, domestic and foreign price levels individually reveal the presence of a unit root, so their combination in the real exchange rate should be stationary. However, studies that used a short span of data found it difficult to prove that there is any mean reversion in the real exchange rate series, which means that there is no convergence to purchasing power parity even in the long-run. However, studies that used a long span of data have found satisfactory evidence that the real exchange rate converge to PPP in the very long-run.

Even in relatively short span of data, mean reversion is sometimes identified especially in countries with experience of high inflation (Breuer, 1994).

However, when one interprets the price index as one pertaining to a broad set of goods and services PPP does not hold even over long periods (Froot and Rogoff, 1995; and Chinn, 2000). Since some of the items in a typical consumption or production bundle are not tradable and subject to international price pressures from international trade, this result is not completely unexpected. On the other hand, since consumer bundles might be more similar across countries than producer or wholesale bundles, consumer price indices (CPIs) may provide a more consistent measure of price levels and thus of real exchange rates. However, we adopted an agnostic view on the issue, and used calculations based on wholesale price index (WPI) and WPI based trade weighted real exchange rate are also presented.

Given the contradictory evidence between short and long span of data and various deflators, it is appropriate to test for the presence of unit roots in the real exchange rate series and their component.

[1]Trade weighted RER is calculated by using the following identity TWRER= P_{Turkey}*[(0.75\$/TL+0.25DM/TL)*(\$/DM)]/[0.75P_{US}+ 0.25$P_{Germany}$)*(\$/DM)] where P is WPI.

Two tests are applied to see whether variables exhibit random walk or stationary behaviour, namely augmented Dickey-Fuller (ADF) and Phillips-Perron tests. In the absence of serially correlated errors, Augmented Dickey-Fuller tests reduce to the Dickey-Fuller test, which is based on following regression

$$\Delta rs_t = \gamma_o + \gamma_1 Trend + \beta rs_{t-1} + \varepsilon_t \qquad (3)$$

to account for serially correlated errors augmented Dickey-Fuller test is used, which is based on the following regression

$$\Delta rs_t = \gamma_o + \gamma_1 Trend + \beta rs_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta rs_{t-i} + \varepsilon_t \qquad (4)$$

where Δ denotes the first difference operator, rs_t is the real exchange rate measured in natural logarithm, and ε_t is a random disturbance term which is assumed to be white noise. Under the null hypothesis, time series has a unit root, $\beta = 0$. The alternative hypothesis is that series mean reverting is, $\beta<0$. For the real exchange rate series rejection of null hypothesis implies that PPP holds. If, however, the estimated value of β is statistically not different from zero, then rs_t contains unit root, which indicates lack of evidence of PPP. Phillips-Perron test accommodates serial correlation and heteroscedasticity in the residuals. In each case the optimal lag structure is chosen using Akaike Information Criteria.

Results reported in columns A of Table 1 indicate that all the real exchange rate series and their components are non-stationary. Further, the coefficient of trend is restricted to zero and both tests are applied. Results are presented in columns B. Results show that all variables exhibit unit roots. In Table 1 we also present unit root tests on the first difference of the time series, which reveals that all the variables are stationary after first differencing.

3.3. Calculation of Half-Lives

Conventional unit root tests are rather uninformative as to the speed of parity reversion. Alternatively, we can concentrate on measuring the duration of shocks to the RER and characterize the extent of parity reversion in terms of point estimates of the "half-life" of deviation from PPP, where the half-life is typically defined as the duration of time required for half the magnitude of a unit shocks to the level of series to dissipate. This provides information for drawing conclusions about the relevance of PPP by using the univariate technique (Cashin and McDermott, 2001).

To estimate the speed of convergence to PPP researchers generally used the Dickey-Fuller type regression as in equation (3) (Froot and Rogoff, 1995; Lothian and Taylor, 1996; Edison et.al., 1997). Time trend is usually not included in the DF regression (Enders, 1995). However, inclusion of time trend controls for the Balassa-Samuelson effect, where the failure of PPP to hold can be due to differential rates of productivity growth in tradable and nontradable sectors (see Froot and Rogoff (1995) and Chinn (2000). Further, Goldfajn and Valdes (1996, 1998) and Goldfajn and Gupta (2001) calculate equilibrium exchange rates from a regression of the logarithm of the reel exchange rate on time trend. For comparison purposes, we present results with and without trend in the regressions. The half-life is calculated from (1-β). However, presence of serial correlation in DF regression in equation (3) will bias estimate of β. Therefore, ADF regression in equation (4) will be more appropriate. Further, if there exists heteroscedastic error in the ADF regression we need another technique. Fortunately, the Phillips-Perron (1987) semi-nonparametric technique can deal with more general error process. The PP technique estimates equation (3) and accounts for serial correlation and heteroscedasticity using the non-parametric method. For AR(p) models half-life gives the length of time until the impulse response of a unit shock is half its original magnitude and is calculated as H=-Ln(2) / Ln(1-β)[2]. If the mean half-life of RER is finite this will be taken as an indication of existence of the PPP relationship.

We proceed by measuring the duration of shocks to the RER and characterize the extent of parity reversion in terms of point estimates of the half-life of deviation from PPP. Table 2 presents the results for the half-life of duration of shocks to the various real exchange rates. The half-lives are calculated by using the least square estimates of (1-β) from ADF and Phillips-Perron regressions. Second and third columns in Table 2 give the estimates of (1-β) from regressions with a constant and associated half-life respectively. Fourth and fifth columns show the results from the regression with a constant and a trend. As can be seen from Table 2a, half-lives calculated from ADF regression with a constant and a trend give on average 12 months.

The ADF regressions presented on the left side of the tables do not attempt to account for the presence of the heteroscedasticity[3]. Accordingly, to account for heteroscedasticity the results of Phillips-Perron regression, which are valid in the presence of heteroscedasticity are presented on the right hand side of the Table. In Table 2b, calculated half-life for CPI based real exchange rate from PP regression is about 16 months. Including time trend does not make much difference.

Table 2c and 2d gives calculated half lives for WPI based real exchange rate from ADF and PP regressions. Calculated half lives from the regressions without

[2] This formula assumes that shocks decay monotonically.

[3] Test for heteroscedasticity carried out on the error terms from the least square regression of equation (3) indicates that in all the models heteroscedasticity is a problem. These test results are available from the author upon request.

trend gives us higher values compared to CPI based real exchange rates. However, regressions with constant and trend reveals half-life values, which are close to CPI based real exchange rates in both regressions.

The last panels in Table 2, give the least square estimates of (1-β) and calculated half-lives for WPI based trade weighted real exchange rate. Least square estimates of β from ADF and PP regressions are about 0.91 and 0.93 respectively. The time it takes for half of the shocks to the TWRER to dissipate is about 8-9 months. Accordingly shocks to the RER of Turkey do not appear to be persistent.

Broadly, least square results indicate that across all definitions of RER in Turkey mean half-life is finite with relatively short mean lengths compared to industrial nations (root and Rogoff, 1995).

4. Cointegration Tests Results

Univariate unit root tests impose undue restrictions on several variables. In estimating an ADF on the RER, one forces the short run dynamics for the nominal exchange rates and both price levels to be the same. In principle, there is no reason to believe that this condition should hold. Hence previous attempts to find mean reversion in the real exchange rate using univariate techniques, have usually failed (Kremers et al., 1992).

A general specification implied by cointegration in the form of vector error correction model (VECM) can be written as following:

$$\Delta s_t = \lambda_1 \left[\beta_1 s + \beta_2 P + \beta_3 P^*\right]_{t-1} + \sum_{i=1}^{k} \alpha_{1i} \Delta s_{t-i} + \sum_{i=1}^{k} \phi_{1i} \Delta P_{t-i} + \sum_{i=1}^{k} \gamma_{1i} \Delta P^*_{t-i} + \varepsilon_{1i}$$

$$\Delta P_t = \lambda_2 \left[\beta_1 s + \beta_2 P + \beta_3 P^*\right]_{t-1} + \sum_{i=1}^{k} \alpha_{2i} \Delta s_{t-i} + \sum_{i=1}^{k} \phi_{2i} \Delta P_{t-i} + \sum_{i=1}^{k} \gamma_{2i} \Delta P^*_{t-i} + \varepsilon_{2i} \qquad (5)$$

$$\Delta P^*_t = \lambda_3 \left[\beta_1 s + \beta_2 P + \beta_3 P^*\right]_{t-1} + \sum_{i=1}^{k} \alpha_{3i} \Delta s_{t-i} + \sum_{i=1}^{k} \phi_{3i} \Delta P_{t-i} + \sum_{i=1}^{k} \gamma_{3i} \Delta P^*_{t-i} + \varepsilon_{3i}$$

These equations show that when $\lambda<0$ any deviation from PPP in the previous period would reduce the growth rate of the exchange rate in the current period, that is, there is a tendency for the exchange rate to return to the equilibrium rate over time. On the other hand, when $\lambda = 0$, there is no correction mechanism in the system that indicates a tendency towards PPP. The VECM specification reveals at least two conceptual problems presented in unit root test for PPP (Steigerwald, 1996). First, coefficients of P_{t-1} and P_{t-1}^* equals unity by construction. Second, the coefficients of Δs_{t-i}, Δp_{t-i} and Δp^*_{t-i} are restricted to be the same. The VECM representation relaxes these assumptions, and also takes care of endogeneity of prices.

Johansen (1988), Johansen and Juselius (1990) and Johansen (1995) describe the maximum likelihood method of estimating this vector error correction model (VECM) and cointegration tests. After finding cointegration among the exchange rate and domestic and foreign prices both symmetry restriction ($\beta_2 = \beta_3$) and homogeneity restrictions (β_2=1), which are required by the strong form of PPP, can be tested by a likelihood ratio test (or ($(\beta_1 \beta_2 \beta_3)$ equals (1 –1 1)). This form of PPP test is first applied by Cheung and Lai (1993).

Johansen procedure is used to determine the rank r and to identify PPP amongst the cointegrating vectors. The number of lags used in the VAR is based on the evidence provided by both a likelihood ratio test and AIC, however, in the case of serial correlation sufficient number of lags introduced to eliminate the serial correlation of the residuals. The cointegration tests amongst s, p and p^* include six lags in the VAR. We introduced a set of monthly centered seasonal dummy variables, restricted trend, unrestricted constant term[4] and further, the estimates of unrestricted VAR include also three impulse dummy variables: *D91* is included to capture the Gulf War in 1991, *D94* is included to capture currency crises in 1994, and *D00* is included to capture effects of 2000 stabilization program.

The specification including trend in the cointegration space imply that the effects of the other variables are not assumed away. The effects of other variables on the real exchange rate are incorporated in the behaviour of the trend. This can be justified on the proposition that while these variables are individually dominated by price changes, they may exert some effect on the exchange rate collectively. Therefore, the effects of other variables are not assumed away but will be tested by the significance of the trend.

Table 3a reports CPI based PPP estimates of Johansen procedure and standard statistics. In determining the number of cointegrating vectors we used the degrees of freedom adjusted version of the maximum eigenvalue and trace statistics, since in the existence of small samples with too many variables or lags, Johansen procedure tends to over estimate the number of cointegrating vectors (Cheung and Lai, 1993); and Gonzalo and Pitarakis, 1994). CPI based results show evidence of at least one cointegration relationship.

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[4] Doornik et.al. (1998) statistically analyze over-specified trend in the cointegration space and suggest that adopting a model that includes a trend in the cointegration space have low cost even when DGP does not display trend. They found that including an unrestricted trend was problematic. However, a restricted trend in the cointegration space with an unrestricted constant produced a good power and reasonable size (for further details see Doornik et al. (1998). Franses (1999) also suggests that exclusion of deterministic trend from cointegration space is not safe. Hjelm and Johansson (2002) reach the same conclusion.

Table 3a also reports standardised eigenvectors, β', and adjustment coefficients, α. The first row of β' is the estimated cointegration vector, can be written as:

$$s_t = 1.691p_t - 1.893\, p_t^* - 0.032\text{Trend}$$
$$(\text{std.err.}) \quad (0.238) \quad (1.480) \quad (0.014)$$

All the coefficients are correctly signed and statistically significant. The unitary coefficient on domestic prices is rejected at a 5 percent significance level, the likelihood ratio statistics and the associated asymptotic p-values are $\chi^2(1) =$ 5.5862 and [0.0181] respectively. However, the symmetry constraint is not rejected (likelihood ratio statistics with $\chi^2(1)$ and p-value are 0.01328 and [0.9083] respectively). The adjustment coefficients (-0.12) suggest relatively quick reversion to PPP over the period under investigation.

Table 3b reports WPI based PPP test results. All the coefficients are statistically significant and correctly signed. The estimated cointegrating vector can be written as

$$s_t = 1.378p_t - 1.432\, p_t^* - 0.015\text{Trend}$$
$$(\text{std.err.}) \quad (0.120) \quad (0.594) \quad (0.005)$$

The unit coefficient on domestic prices restriction is marginally rejected at a 5 percent significance level, the associated likelihood ratio statistic and asymptotic p-values are $\chi^2(1) = 4.1939$ and [0.0406] respectively. Furthermore, the symmetry restrictions of coefficients of domestic and foreign prices are approximately the equal magnitude and opposite sign has been tested and we cannot reject this hypothesis $\chi^2(1) = 0.003$ [0.9554]. The adjustment coefficients -0.085 suggest relatively quick reversion to PPP over the period under investigation.

Enders (1995) argues that restricted trend in the cointegration space is not consistent with the absolute version of the PPP. Therefore, we investigate existence of PPP relationships by including a constant term into the cointegration space instead of time trend. CPI based PPP estimates of the Johansen procedure with restricted constants is presented in Table 3c. CPI based results show evidence of at least one cointegration relationship.

The first row of β' is the estimated cointegration vector is written below

$$s_t = 1.185p_t - 4.985\, p_t^* + 27.545\text{Const.}$$
$$(\text{std.err.}) \quad (0.033) \quad (0.565) \quad (2.425)$$

The coefficients are correctly signed and statistically significant. The test result for the unitary coefficients on domestic price is valid where likelihood ratio statistics with $\chi^2(1)$ and p-value are 2.5687 and [0.1090] respectively. However, the symmetry constraint is rejected (likelihood ratio statistics with $\chi^2(1)$ and p-value are 14.536 and [0.0002] respectively). This may be because of the trend in relative prices of traded and non-traded goods (for other possible reasons (Froot and Rogoff, 1995). The adjustment coefficient found in this specification is -0.112, which is very close to our previous specification and suggests a relatively quick reversion to PPP.

Table 3d reports WPI based PPP test results with restricted constant. The estimated cointegrating vector can be written as

$$s_t = 1.075p_t - 4.209\, p_t^* + 22.824\text{Const}$$
$$(\text{std.err.}) \quad (0.083) \quad (2.450) \quad (10.826)$$

All the coefficients are statistically significant and correctly signed. In this specification the unit coefficient on domestic prices restriction is not rejected at a 5 percent significance level, (the associated likelihood ratio statistic and asymptotic p-values are $\chi^2(1) = 0.2454$ and [0.6203] respectively). Further, the symmetry restriction of coefficients of domestic and foreign prices approximately equal in magnitude and opposite in sign is tested, and we cannot reject this hypothesis either. The adjustment coefficient 0.038 suggests overshooting of real exchange rate of its equilibrium value and indicates slow reversion to PPP over the period under investigation.

Cointegration results presented here show that relaxing the implicit restriction in standard tests for unit roots appears to lead non-rejection of PPP. Finally our test results show that the unitary coefficient on domestic prices implied by PPP is rejected for the specification with restricted trend in the cointegration space while the coefficients on domestic and foreign prices opposite in sign and equal in magnitude is not rejected for this specification. On the other hand, the specification with constant accepts unitary coefficient on domestic prices but rejects the symmetry restrictions. Therefore, we conclude that the Turkish data does only support the weak form of PPP, for 1987:1-2000:12 periods.

5. Estimated Equilibrium Rates and Misalignment

When the real exchange rate is mean reverting, or there exists a long-run relationship among the exchange rate, domestic and foreign prices, the long-run equilibrium exchange rate may be estimated on the basis of PPP.

The equilibrium real exchange rate involves removing the effects of non-systematic transitory shocks. In practice, these are eliminated by identifying a base period in which such shocks are negligible. This ensures that actual real exchange rate coincides with its equilibrium value in the base period. Thus the actual real exchange rate in the base period represents the estimate of the equilibrium rate, and the nominal exchange rate is consistent with the equilibrium real exchange rate from that moment calculated by simply adjusting the nominal exchange rate for the cumulative difference between domestic and foreign inflation.

The alternative case is that the equilibrium real exchange rate is interpreted as subject to change in response to changes in underlying fundamentals. In this case the equilibrium real exchange rate can still be measured using a base year value, but identifying a suitable base year involves analyzing fundamentals that determine the equilibrium real exchange rate and actual real exchange rate were at a sustainable level. If the fundamentals do not change after the base year or return to their level in that year then the misalignment can be calculated as a difference between actual RER in the current year and its equilibrium value in the base year.

The base year approach assumes that all the fundamentals are close to their sustainable levels. In practice, one usually focuses on the external balance criterion. This involves choosing a year with a reasonable current account deficit. For assessing the sustainability of the variable, one looks for terms of trade that are reasonable, close to their long-run trend level and capital flows that are consistent with the likely long-term availability of capital and with the country's debt servicing capacity. For assessing the sustainability of objective variables one looks at growth, investment, employment, inflation performance and compares these to the country's long-run policy targets (Bayoumi et.al., 1994; and Monteil, 1999).

It is also desirable to select a base year as recent as possible to minimize the changes in the economy's structure taking place between the base and the current year. If a country has a market determined exchange rate that fluctuates significantly, selecting a short time period might be more representative of the equilibrium values of the RER than a single base year estimate.

One way of dealing with fluctuations in the fundamentals during the sample period is to estimate their sustainable values on the basis of their sample mean, or in the case of trend stationarity, on the basis of trend values within the sample. This procedure amounts to estimating the equilibrium real exchange rate at the sample mean or the trend values of the real exchange rate within the sample rather than as the particular value in a specified base year. Hence, instead of trying to identify a particular year or short span of years in which the real exchange rate is believed to be at its equilibrium value, one tries to identify the long-term trend value toward which the actual real exchange rate tends. Thus, the equilibrium real exchange rate could be estimated as being the mean value of the real exchange rate over a long period of time or evolving along a deterministic or stochastic trend and misalignment is then measured as the deviation from this trend or mean value.

Following the above discussion, we have used three different measures of misalignment. First, we obtained the mean real exchange rates using a regression of the real exchange rate on the constant and the fitted values from the regression is taken as the estimated equilibrium exchange rate. Then misalignment is calculated as the difference between the actual and equilibrium rate.

In the second measure of misalignment the real exchange rate is allowed to move with linear deterministic trend over the period. This amounts to estimating following regression:

$$rs_t = \alpha + \beta \ Trend + \varepsilon_t$$

In the third alternative, the Hodrick-Prescott- filtered real exchange rate series is taken as an estimate of equilibrium real exchange rate, which captures the permanent changes in relative prices between foreign and domestic currency. The difference between the actual and the filtered series represents the cyclical components of RER movements (Goldfajn and Gupta, 2001). Then misalignment is calculated as before.

Table 4a reports these three alternative measures of misalignment for the CPI based real exchange rates. The second column of Table 4a shows misalignment where equilibrium real exchange rate is obtained from running regression of the real exchange rate on the constant over 1987:1-2000:12 periods. Results indicate that in December 2000 overvaluation of the TL is about 8 percent. Given our uncertainty regarding all types of PPP calculations, it is important to undertake some robustness check against the use of different sample periods. We recalculated equilibrium value using 1995:1-2000:12 periods instead of the whole sample. If the real exchange rate series were really mean reverting, changing the sample period should not matter very much. This result, presented in the third column, is very close to the misalignment in the whole sample.

Misalignment based on alternative estimates of the equilibrium real exchange rate, allowing for a trend in the real exchange rate for the whole sample, is given in the fourth column of Table 4a. This measure of misalignment gives us about a 4 percent overvaluation in December 2000. Again to check the robustness of this result, estimates of equilibrium rate are obtained for 1995:1-2000:12 periods. The result, which is presented in the fifth column of the table, shows a 4.2 percent overvaluation in December 2000.

Finally, the sixth column in the table gives misalignment based on the Hodrick-Prescott filtered series. This result gives about a 4 percent overvaluation in December 2000.

In Figure 1, the equilibrium rates and actual levels of real exchange rates corresponding to columns 2-6 in Table 4a are plotted. The implied overvaluation derived from the CPI measures is consistent with historical accounts.

Table 4b reports the alternative measures of misalignment for the WPI based real exchange rates. As can be seen from the table in all accounts TL appears to be undervalued against the US dollar, however, the magnitude of undervaluation varies depending on the estimation of equilibrium rate. Only in the last column where misalignment is based on H-P filtered series it shows very marginal overvaluation. The implied undervaluation derived from the WPI measures are plotted in Figure 2.

Results obtained by using CPI and WPI are not consistent. There may be several reasons for this. One of them is related to developments in the terms of trade (relative price of imports and exports). When the terms of trade change, movements in the production and expenditure price indexes differ significantly. Deterioration in the terms of trade leads to faster rise in the CPI compared to the WPI. Therefore, calculated real exchange rates based on these prices will also diverge. Further, we wanted to see whether misalignment, using trade weighted real effective real exchange rate based on WPI, shows any overvaluation. Results are presented in Table 4c. All of the results show overvaluation. Overvaluation, where the equilibrium rates obtained from the whole sample, is less than the one from sub sample of 1995:1-2000:12 and CPI based overvaluations. Figure 3 confirms these findings.

6. Summary and Conclusions

This paper has documented the findings of mean reversion for exchange rates over the 1987-2000 period. These results can be interpreted as detection of purchasing power parity. We implemented two univariate time series techniques, unit root tests and calculated the “half-life” of various real exchange rates. Augmented Dickey-Fuller and Phillips-Perron unit root test results showed that real exchange rates are non-stationary. However, these tests impose ex-ante undue restrictions on the parameters and short-run dynamics and, do not allow for endogeneity of prices. Therefore, these results are expected. Further univariate analysis based on point estimates of autoregressive parameters from ADF and PP regressions were used to calculate half-life deviation from the parity. These results showed that deviation from PPP is very short, implying that real exchange rate means revert to PPP level relatively quickly. The only exception is that WPI based bilateral real exchange rate without trend gives a relatively longer half-life in ADF regression. Given that the estimated (1-β) is biased downwards in ADF regression we should be cautious about the exact values of the calculated half-lives. Further evidence on the validity of PPP is provided by cointegration analysis. Finally three different estimates of equilibrium real exchange rates have been used to calculate the misalignment to see whether TL was overvalued before the 2001 crises in Turkey. Evidence provided in this paper shows that there is a significant overvaluation on the CPI based bilateral real exchange rate and the WPI based trade weighted real exchange rate. However, the WPI based bilateral real exchange rate shows undervaluation of TL against US dollar. These results suggest that CPI based bilateral real exchange or trade weighted real exchange rate should be used as a leading indicators of a financial crisis.

References

Akinci, D.A. and Demir, A., 1993. "Purchasing Power Parity and Exchange Rates: A Cointegration Analysis". Turkish Development Banks, Discussion Papers 13-1APM-30.

Balassa, B., 1964. "The Purchasing Power Parity Doctrine: A Reappraisal". *Journal of Political Economy*. 72, pp.584-596.

Bayoumi, T., Clark, P., Symansky, S., Taylor, M., 1994. "The Robustness of Equilibrium Exchange Rate Calculations to Alternative Assumptions and Methodologies". In J. Williamson (ed.), *Estimating Equilibrium Exchange Rates*., pp.19-60. Institute for International Economics, Washington, DC.

Berg, A., Borensztein, E., Milesi-Ferretti, G.M., Pattillo, C., 2000. "Anticipating Balance of Payment Crises, The Role of Early Warning Systems". IMF Occasional Paper 186. IMF, Washington, DC.

Breuer, J., 1994. "Purchasing Power Parity: A Survey of and Challenge to Recent Literature." In J. Williamson (ed.) *Estimating Equilibrium Exchange Rates*, , pp. 245-277. Institute for International Economics, Washington, DC.

Cashin, P. and McDermott, C.J., 2001. "An Unbiased Appraisal of Purchasing Power Parity". IMF Working Paper, WP/01/196.

Cheung, Y.-W., Lai, K.S., 1993a. "Finite-Sample Sizes of Johansen's Likelihood Ratio Tests for Cointegration". *Oxford Bulletin of Economics and Statistics*, 55, pp.313-328.

Cheung, Y. W., Lai, K.S., 1993b. "Long-Run Purchasing Power Parity During the Recent Float". *Journal of International Economics*. 34, pp.181-192.

Chinn, M.D., 1999. "Measuring Misalignment: PPP and East Asian Currencies in the 1990s". IMF Working Paper WP/99/120. International Monetary Fund, Washington, DC.

Doornik, J. A., Hendry, D.F. and Nielsen, B., 1998. "Inference in Cointegrating Models: UK M1 Revisited". *Journal of Economic Surveys*, 12(5), pp. 533-572.

Dornbusch, R., 1976. "Expectations and Exchange Rate Dynamics". *Journal of Political Economy* 84, pp.1161-1176.

Edison, H., Gagnon, J., Melick, W., 1997. "Understanding the Empirical Literature on Purchasing Power Parity: The Post-Bretton Woods Era". *Journal of International Money Finance* 16, pp.1-17.

Enders, W., 1995. *Applied Econometric Time Series*, John Wiley & Sons, New York.

Erlat, H., 2001. "The Nature of Persistence in Turkish Real Exchange rates". METU Economic Research Centre Working Paper.

Frankel, J.A. and Rose, A.K., 1996. "A Panel Project on Purchasing Power Parity: Mean Reversion within and Between Countries". *Journal of International Economics*, 40, pp.209-224.

Franses, P.H., 1999. "How to Deal with Intercept and Trend in Practical Cointegration Analysis" Econometric Institute Research Report, 9904/A.

Froot, K.A., Rogoff, K., 1995. "Perspectives on PPP and Long-Run Real Exchange Rates". In: Grossman, G.,Rogoff, K. Eds. , *Handbook of International Economics*, vol. 3. Elsevier Press, Amsterdam, pp.1648-1684.

Glen, J.D., 1992. "Real Exchange Rates in the Short, Medium, and Long Run". *Journal of International Economics* 33, pp.147-166.

Goldfajn, I., Valdes, R.O., 1996, "The Aftermath of Appreciations". *NBER, Working Paper* No. 5650. Washington MA.

Goldfajn, I. and Valdes, R.O., 1998. "Current Account Sustainability: Are Currency Crises Predictable?" *European Economic Review*, 42, pp.873-885.

Goldfajn, I., Valdes, R., 1999. "The Aftermath of Appreciations". *Quarterly Journal of Economics* 114, pp.229-262.

Goldfajn, I., Gupta, P., 2001. "Overshootings and Reversals: The Role of Monetary Policy". Central Bank of Chile Working Papers, No: 126

Gonzalo, J. and Pitarakis, J. Y., 1994. "Cointegration Analysis in Large Systems", Technical Report, Department of Economics, Boston University.

Hjelm, G. and Johansson, M.W., 2002. "A Monte Carlo Study on the Pitfalls in Determining Deterministic Components in Cointegrating Models", mimeo, Lund University.

Hinkle, L.E., Montiel, P.J., 1999. *Exchange Rate Misalignment*. Oxford University Press, World Bank, New York.

Isard, P., Symansky, S., 1996. "Long Run Movements in Real Exchange Rates". In Ito, T., Isard, P., S. Symansky, and T. Bayoumi (eds.) *Exchange rate movements and their impact on trade and investment in the APEC region* . Occasional Paper 145. International Monetary Fund, Washington, DC.

Johansen, S., 1988. "Statistical Analysis of Cointegrating Vectors". *Journal of Economics and Dynamic Control*, 12, pp.231-254.

Johansen, S., 1995. *Likelihood-Based Inference in Cointegrated Vector Autoregressive Models*, Oxford, Oxford University Press.

Johansen, S., Juselius, K., 1990. "Maximum Likelihood Estimation and Inference on Cointegration with Applications to the Demand for Money". *Oxford Bulletin of Economics and Statistics*. 52, pp.169-210.

Kaminsky, G.L., Reinhart, C.M., 1999. "The Twin Crises: The Causes of Banking and Balance-of-Payment Problems". *American Economic Review* 89, pp.473-500.

Kamisky, G. Lizondo, S. and Reinhart, C.M., 1998. "Leading Indicators of Currency Crises". *IMF Staff Papers*, March 1998, pp.1-48

Kremers, J., Ericsson, N., Dolado, J., 1992. "The Power of Cointegration Tests". *Oxford Bulletin of Economics and Statistics* 54, pp.325-348.

Lee, D.Y., 1999. "Purchasing Power Parity and Dynamic Error Correction: Evidence from Asia Pacific Economies". *International Review of Economics and Finance* 8, pp.199-212

Lothian, J.R. and Taylor, M.P., 1996. "Real Exchange Rate Behaviour: The Recent Float from the Perspective of Two Centuries". *Journal of Political Economy*, 104, pp.488-509

Lothian, J.R. and Taylor, M.P., 2000. "Purchasing Power Parity over Two Centuries: Strengthening the Case for Real Exchange Rate Stability. A Reply to Cuddington and Liang", *Journal of International Money and Finance* 19, pp.759-764.

Metin, K., 1994. "A Test for Long-Run Purchasing Power Parity and Uncovered Interest Rate Parity: Turkish case", Discussion paper no 94-2. Department of Economics, Bilkent University

Milesi-Ferretti, G.M., Razin, A., 1996. "Current Account Sustainability: Selected East Asian and Latin American Experiences". NBER Working Paper 5791.

Papell, D.H., 1997. "Searching for Stationarity: Purchasing Power Parity Under the Current Float". *Journal of International Economics* 43, pp.313-332.

Phylaktis, K., Kassimatis, Y., 1994. "Does the Real Exchange Rate Follow a Random Walk? The Pacific Basin perspective". *Journal of International Money Finance* 13, pp.476-495.

Phillips, P.C.B, and Perron, P., 1988. "Testing for Unit Root in Time Series Regression". *Biometrica*, 75, pp.335-346

Sachs, J., Tornell, A., Velasco, A., 1996. "Financial Crises in Emerging Markets: the Lessons from 1995". *Brookings Papers on Economics Activity*, 1996, 1 , pp.147-215.

Samuelson, P., 1964. "Theoretical Notes on Trade Problems". *Review of Economics and Statistics*. 46, pp.145-154.

Sarno, L., 2000. "Real Exchange Rate Behaviour in High Inflation Countries: Empirical Evidence from Turkey, 1980-1997". *Applied Economic Letters*, 7, pp.285-291.

Steigerwald, D.G., 1996. "Purchasing Power Parity, Unit Roots, and Dynamic Structure". *Journal of Empirical Finance* 2, pp.343-357.

Taskın, F. and Metin, K., 1994. "Does Purchasing Power Parity Hold in the Long Run? Evidence for Developing Countries", Working paper No: 94-19. Department of Economics, Bilkent University.

Taylor, A. M., 2002. "A Century Of Purchasing-Power Parity" *The Review of Economics and Statistics*, 84, pp. 139-150.

Taylor, M. P. and Sarno, L., 1998. "The Behavior of Real Exchange Rates During the Post-Bretton Woods Period". *Journal of International Economics* 46 (2), pp.281–312.

Telatar, E. and Kazdağlı, H., 1998. "Re-examine the Long Run Purchasing Power Parity Hypothesis for a High Inflation Country: The case of Turkey". *Applied Economics Letters*, 5, pp.51-53.

Williamson, J. Ed. , 1994. *Estimating Equilibrium Exchange Rates*. Institute for International Economics, Washington, DC.

Figure 1: CPI Based Actual and Equilibrium RERs

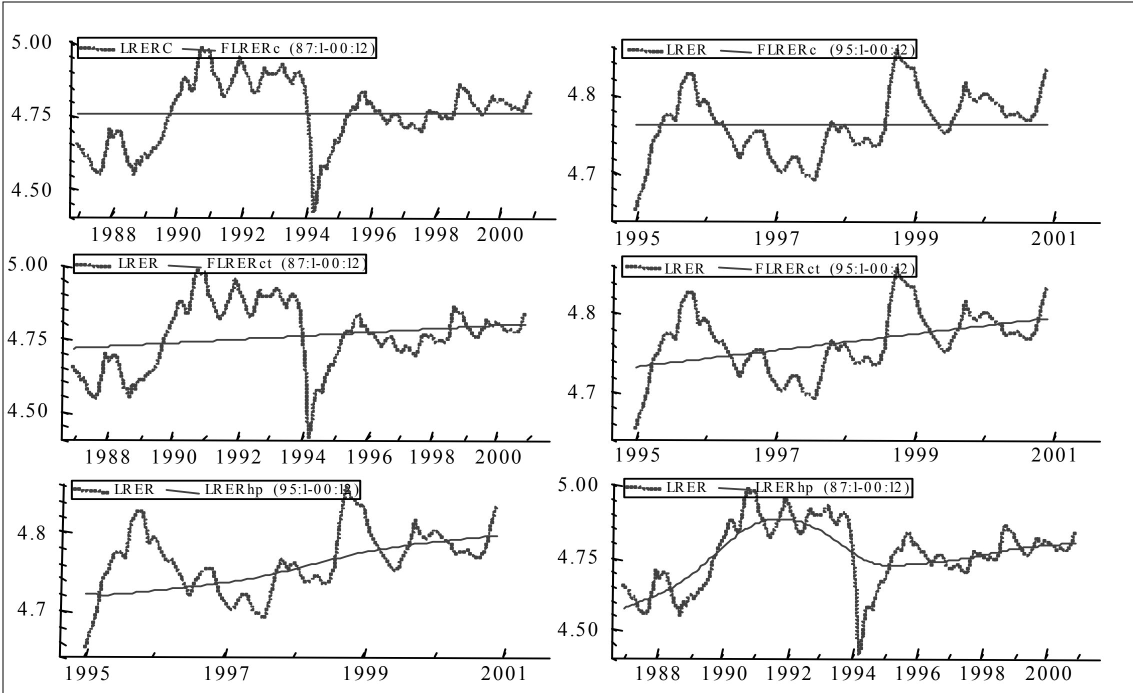

Figure 2: WPI Based Actual and Equilibrium RER

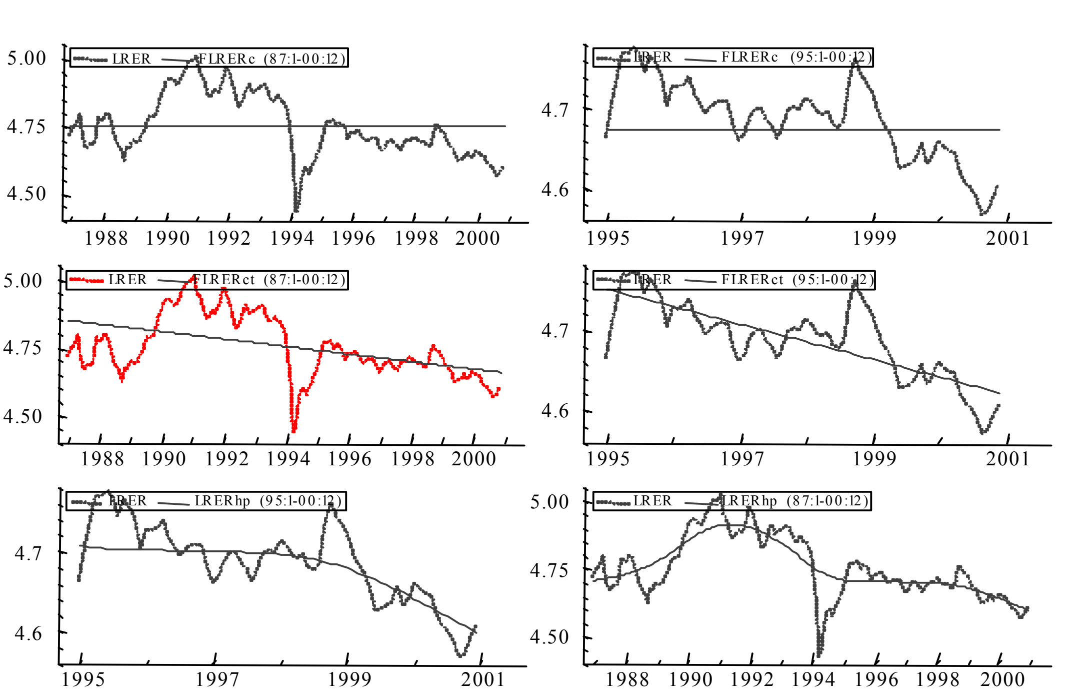

Figure 3: WPI Based Actual and Equilibrium TWRER

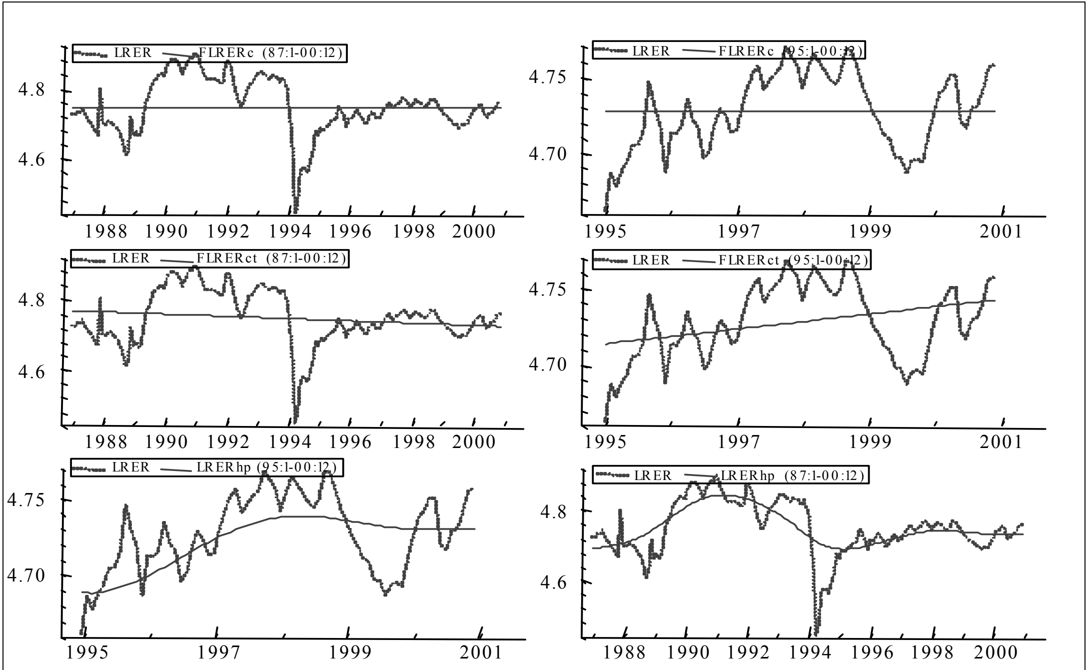

Table 1a: ADF(k) Unit Root Test Results

(Levels)				(First Differences)			
Variables	**k**	**A**	**B**	**Variables**	**k**	**A**	**B**
LRERc	12	-2.399	-2.390	ΔLRERc	12	-3.679*	-3.719**
LRERw	12	-2.371	-1.523	ΔLRERw	12	-3.801*	-3.745**
LRERwt	12	-3.001	-2.815	ΔLRERwt	12	-4.238**	-4.229**
LEXU	12	-2.050	0.151	ΔLEXU	7	-3.859*	-3.879**
LCPI	12	-0.817	-2.139	ΔLCPI	7	-4.533**	-4.547**
LWPI	12	-0.671	-1.738	ΔLWPI	7	-4.140**	-4.154**
LCPIU	12	-2.652	-2.999	ΔLCPIU	7	-3.766*	-3.491**
LWPIU	12	-3.291	-1.450	ΔLWPIU	7	-4.196**	-3.354*

Table 1b: Phillips-Perron (k) Unit Root Test Results

(Levels)				(First Differences)			
LRERc	12	-2.358	-2.345	ΔLRERc	12	-8.013**	-8.049**
LRERw	12	-2.474	-1.863	ΔLRERw	12	-8.197**	-8.219**
LRERwt	12	-2.697	-2.663	ΔLRERwt	12	-11.401**	-11.441**
LEXU	12	-2.226	0.388	ΔLEXU	7	-8.198**	-8.218**
LCPI	12	-2.389	0.428	ΔLCPI	7	-8.471**	-8.498**
LWPI	12	-1.722	-0.101	ΔLWPI	7	-7.928**	-7.961**
LCPIU	12	-1.970	-4.504	ΔLCPIU	7	-10.259**	-9.582**
LWPIU	12	-2.687	-1.813	ΔLWPIU	7	-9.129**	-9.119**
1% Crt.Val*		-4.026	-3.478	1% Crt.Val*		-4.026	-3.478
5% Crt. Val		-3.443	-2.882	5% Crt. Val		-3.443	-2.882

Notes: 1. LRERc, LERw and LRERwt are log of CPI , WPI and WPI based trade weighted real exchange rates respectively; LEXU is log TL per US dollar nominal exchange rate; LCPI and LCPIU are log of domestic and US CPI (1994=100) respectively; LWPI and LWPIU are log of domestic and US WPI (1994=100) respectively; 2. Sample period is 1987:1-2000:12. k is the number of lagged dependent variables in the ADF regression. 3. Column A and B give the t-statistics from ADF regression including constant and trend and, constant respectively. 4. The critical values are from MacKinnon (1991). The superscripts * and ** denotes rejection at 5% and 1% critical values.

Table 2: Half-Life of Parity Deviations

Table 2a: Half-Lives of CPI based RER Deviations in ADF Regression

Lag	$(1-\beta)^1$	Half-life$^{1/2}$	$(1-\beta)^3$	Half-Life$^{3/2}$
12	0.931	10	0.935	10
11	0.932	10	0.937	11
10	0.934	10	0.939	11
9	0.940	11	0.944	12
8	0.940	11	0.944	12
7	0.942	12	0.946	13

Table 2b: Half-Lives of CPI based RER Deviations in Phillips-Perron Regression

Lag	$(1-\beta)^4$	Half-life$^{4/2}$	$(1-\beta)^5$	Half-Life$^{5/2}$
12	0.957	16	0.960	17
11	0.957	16	0.960	17
10	0.957	16	0.960	17
9	0.957	16	0.960	17
8	0.957	16	0.960	17
7	0.957	16	0.960	17

Table 2c: Half-Life of WPI based RER Deviations in ADF Regression

Lag	$(1-\beta)^1$	Half-life$^{1/2}$	$(1-\beta)^3$	Half-Life$^{3/2}$
12	0.962	18	0.943	12
11	0.957	16	0.938	11
10	0.963	18	0.944	12
9	0.960	17	0.941	11
8	0.961	17	0.942	12
7	0.959	17	0.940	11

Table 2d:Half-Life of WPI based RER Deviations in Phillips-Perron Regression

Lag	$(1-\beta)^4$	Half-life$^{4/2}$	$(1-\beta)^5$	Half-Life$^{5/2}$
12	0.970	15	0.951	14
11	0.970	15	0.951	14
10	0.970	15	0.951	14
9	0.970	15	0.951	14
8	0.970	15	0.951	14
7	0.970	15	0.951	14

Table 2e:Half-Life of WPI based TWRER Deviations in ADF Regression

Lag	(1-β)[1]	Half-life[1/2]	(1-β)[3]	Half-Life[3/2]
12	0.902	7	0.887	6
11	0.914	8	0.901	7
10	0.915	8	0.902	7
9	0.904	7	0.892	6
8	0.914	8	0.903	7
7	0.910	7	0.899	7

Table 2f:Half-Life of WPI based TWRER Deviations in Phillips-Perron Regression

Lag	(1-β)[4]	Half-life[4/2]	(1-β)[5]	Half-Life[5/2]
12	0.926	9	0.919	8
11	0.926	9	0.919	8
10	0.926	9	0.919	8
9	0.926	9	0.919	8
8	0.926	9	0.919	8
7	0.926	9	0.919	8

Notes: 1.Results are based on the least squares estimates of the Augmented Dickey-Fuller regression with constant. 2. Half-life is the length of time it takes for a unit impulse to dissipate by half. It is derived using the formula H=ln(2)/ln(1-β), where (1-β) is the autoregressive parameter. 3. Results are based on the least squares estimates of the Augmented Dickey-Fuller regression with constant and trend. 4. Results are based on the least squares estimates of the Phillips-Perron regression with constant. 5. Results are based on the least squares estimates of the Phillips-Perron regression with constant and trend

Table 3a: Cointegration Analysis of PPP (CPI)

Eigenvalues	0.161	0.085	0.045
Hypotheses	r = 0	r <= 1	r <= 2
λ-max	29.58*	14.840	7.713
λ-max (d.f. adjusted)	26.41*	13.250	6.887
95% critical values	25.500	19.000	6.887
λ-trace	52.14**	22.560	7.713
λ-trace (d.f. adjusted)	46.55*	20.14	6.887
95% critical values	42.4	25.3	12.3
Standardized eigenvectors (β')			
LEXU	LCPI	LCPIU	Trend
1	-1.6909	1.8925	0.03224
-1.2724	1	-6.04	0.027663
-0.032325	0.257	1	-0.013236
Standardized adjustment coefficients (α)			
LEXU	-0.120	0.051	-0.380
LCPI	0.060	0.007	-0.205
LCPIU	-0.005	-0.002	-0.012

Notes: 1) The estimation period is 1987:1-2000:12. VAR includes 6 lags on each variable, a constant term, centred seasonal monthly dummy variables, D91, D94 and D00 dummy variables. Trend variables are restricted to the cointegration space. 2) The λ-max and λ-trace are maximum eigenvalue and trace test statistics. The critical values are taken from Osterwald-Lenum (1992).

3) The values in [.] are p-values. The * and ** indicate rejection of likelihood ratio tests at 5% and 1% significance levels, respectively

Table 3b: Cointegration Analysis of PPP (WPI)

Eigenvalues	0.127	0.095	0.019
Hypotheses	r = 0	r <= 1	r <= 2
λ-max	22.760	16.680	3.245
λ-max (d.f. adjusted)	19.920	14.600	2.839
95% critical values	25.500	19.000	2.839
λ-trace	42.69*	19.930	3.245
λ-trace (d.f. adjusted)	42.69*	17.44	2.839
95% critical values	42.4	25.3	12.3
Standardized eigenvectors (β')			
LEXU	LWPI	LWPIU	Trend
1	-1.3778	1.4317	0.014812
-0.83219	1	-4.3394	-0.0032115
-0.45404	1.1388	1	-0.032759
Standardized adjustment coefficients (α)			
LEXU	-0.085	-0.049	-0.065
LWPI	0.032	-0.057	-0.023
LWPIU	0.021	0.008	-0.005

Notes:1) The estimation period is 1981:1-2000:12. VAR includes 7 lags on each variable, a constant term, centred seasonal monthly dummy variables, D91, D94 and D00 dummy variables. Trend variables are restricted to the cointegration space. 2) The λ-max and λ-trace are maximum eigenvalue and trace test statistics. The critical values are taken from Osterwald-Lenum (1992).
3) The values in [.] are p-values. The * and ** indicate rejection of likelihood ratio tests at 5% and 1% significance levels, respectively.

Table 3c: Cointegration Analysis of PPP (CPI)

Eigenvalues	0.189	0.097	0.043
Hypotheses	r = 0	r <= 1	r <= 2
λ-max	28.66**	17.1*	7.383
λ-max (d.f. adjusted)	28.84**	14.050	6.065
95% critical values	22.000	15.700	6.065
λ-trace	59.59**	24.48*	7.383
λ-trace (d.f. adjusted)	48.95**	20.11*	6.065
95% critical values	34.9	20	9.2
Standardized eigenvectors (β')			
LEXU	LCPI	LCPIU	Constant
1	-1.1852	4.9856	-27.545
-0.78705	1	-5.3449	26.514
0.58954	-0.63621	1	-7.4952
Standardized adjustment coefficients (α)			
LEXU	-0.112	-0.040	-0.049
LCPI	0.015	-0.028	-0.005
LCPIU	-0.005	0.000	0.004

Notes: 1) The estimation period is 1981:1-2000:12. VAR includes 10 lags on each variable, a constant term, centred seasonal monthly dummy variables, D91, D94 and D00 dummy variables. Trend variables are restricted to the cointegration space. 2) The λ-max and λ-trace are maximum eigenvalue and trace test statistics. The critical values are taken from Osterwald-Lenum (1992). 3) The values in [.] are p-values. The * and ** indicate rejection of likelihood ratio tests at 5% and 1% significance levels, respectively

Table 3d: Cointegration Analysis of PPP (WPI)

Eigenvalues	0.159	0.082	0.053
Hypotheses	r = 0	r <= 1	r <= 2
λ-max	29.15**	14.340	9.234
λ-max (d.f. adjusted)	22.91*	11.270	7.255
95% critical values	22.000	15.700	7.255
λ-trace	52.73**	23.58*	9.234
λ-trace (d.f. adjusted)	41.43**	18.52	7.255
95% critical values	34.9	20	9.2
Standardized eigenvectors (β')			
LEXU	LWPI	LWPIU	Constant
1	-1.0751	4.2086	-22.824
-0.91558	1	-2.3737	15.54
0.15463	-0.18048	1	-5.2926
Standardized adjustment coefficients (α)			
LEXU	0.038	0.100	0.030
LWPI	0.026	-0.006	0.114
LWPIU	0.002	-0.009	-0.054

Notes: 1) The estimation period is 1981:1-2000:12. VAR includes 12 lags on each variable, a constant term, centred seasonal monthly dummy variables, D91, D94 and D00 dummy variables. Trend variables are restricted to the cointegration space. 2) The λ-max and λ-trace are maximum eigenvalue and trace test statistics. The critical values are taken from Osterwald-Lenum (1992). 3) The values in [.] are p-values. The * and ** indicate rejection of likelihood ratio tests at 5% and 1% significance levels, respectively

Table 4a: CPI based Real Exchange Rate Misalignment ($rs_t - r\hat{s}_t$)(1)

Date	MisRER(2)	MisRER(3)	MisRER(4)	MisRER(5)	MisRER(6)
1999-12	0.034	0.028	-0.002	0.009	0.004
2000-1	0.045	0.039	0.008	0.019	0.014
2000-2	0.042	0.036	0.004	0.015	0.010
2000-3	0.033	0.027	-0.005	0.005	0.000
2000-4	0.029	0.024	-0.009	0.001	-0.004
2000-5	0.014	0.009	-0.025	-0.015	-0.020
2000-6	0.017	0.012	-0.022	-0.012	-0.018
2000-7	0.019	0.013	-0.021	-0.012	-0.017
2000-8	0.012	0.007	-0.028	-0.019	-0.025
2000-9	0.009	0.003	-0.032	-0.024	-0.029
2000-10	0.019	0.013	-0.023	-0.015	-0.020
2000-11	0.048	0.042	0.006	0.014	0.008
2000-12	0.077	0.071	0.034	0.042	0.036

Table 4b: WPI based Real Exchange Rate Misalignment ($rs_t - r\hat{s}_t$ (1)

	MisRER(2)	MisRER(3)	MisRER(4)	MisRER(5)	MisRER(6)
1999-12	-0.112	-0.033	-0.032	-0.003	-0.003
2000-1	-0.093	-0.014	-0.012	0.018	0.020
2000-2	-0.099	-0.019	-0.017	0.014	0.018
2000-3	-0.106	-0.026	-0.022	0.009	0.015
2000-4	-0.107	-0.027	-0.022	0.010	0.018
2000-5	-0.133	-0.053	-0.047	-0.014	-0.004
2000-6	-0.145	-0.065	-0.058	-0.024	-0.013
2000-7	-0.152	-0.073	-0.064	-0.030	-0.016
2000-8	-0.165	-0.085	-0.076	-0.040	-0.025
2000-9	-0.184	-0.105	-0.094	-0.058	-0.040
2000-10	-0.181	-0.101	-0.090	-0.053	-0.033
2000-11	-0.160	-0.081	-0.068	-0.031	-0.009
2000-12	-0.147	-0.067	-0.054	-0.015	0.009

Table 4c: Trade Weighted Effective Real Exchange Rate Misalignment ($rs_t - r\hat{s}_t$) (1)

	MisRER(2)	MisRER(3)	MisRER(4)	MisRER(5)	MisRER(6)
1999-12	-0.037	-0.017	-0.019	-0.027	-0.021
2000-1	-0.021	-0.002	-0.003	-0.012	-0.005
2000-2	-0.009	0.011	0.010	0.001	0.008
2000-3	-0.005	0.015	0.014	0.004	0.012
2000-4	0.004	0.023	0.023	0.012	0.021
2000-5	0.004	0.023	0.023	0.012	0.021
2000-6	-0.027	-0.007	-0.007	-0.019	-0.010
2000-7	-0.030	-0.011	-0.011	-0.023	-0.014
2000-8	-0.018	0.002	0.003	-0.011	-0.001
2000-9	-0.016	0.003	0.004	-0.010	0.000
2000-10	-0.006	0.013	0.014	0.000	0.011
2000-11	0.009	0.028	0.030	0.015	0.026
2000-12	0.010	0.030	0.031	0.015	0.027

Notes: (1) Positive (negative) number indicates overvaluation (undervaluation) of TL against US $. (2) Equilibrium RER is estimated by running regression LRER on constant for 1987:1-2000:12 period. (3) Equilibrium RER is estimated by running regression LRER on constant for 1995:1-2000:12 period. (4) Equilibrium RER is estimated by running regression LRER on constant and trend for 1987:1-2000:12 period. (5) Equilibrium RER is estimated by running regression LRER on constant and trend for 1995:1-2000:12 period. (6) Equilibrium RER is Hodrick-Prescott filtered RER series 1987:1-2000:12 period