# PERSISTENCE IN TURKISH REAL EXCHANGE RATES: PANEL APPROACHES

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**ABSTRACT**

Testing whether real exchange rates are stationary and, thereby, obtaining evidence of whether the absolute version of the purchasing power parity (PPP) hypothesis holds, have, initially, be done by using the ADF statistic to test for a unit root. Subsequently, to mitigate the low power of the ADF test, several alternatives have been used for the same purpose. Panel unit root testing is one of these alternatives.

In Erlat (2003), I had previously considered two other alternatives; namely, introducing multiple structural shifts in the deterministic terms and fractional integration, in the context of the two primary bilateral Turkish real exchange rates; the $US and the German DM based rates. This investigation did indicate that these two rates may, in fact, be taken to be stationary with significant long-memory components. In the present paper, I utilise panel procedures to see if they, also, give corroborating evidence.

I used monthly data for the period 1984.01-2001.06 and constructed a panel of 17 bilateral CPI-based real exchange rates corresponding to Turkey's main trading partners for which complete data were available. I implemented seven panel procedures. The first two, Levin, Lin and Chu (LLC) (2002) and Im, Pesaran and Shin (IPS) (2003) are the most commonly used procedures. LLC assumes a common coefficient for the lagged dependent variable in the autoregressions while IPS recognises the full heterogeneity of the coefficients. The third procedure utilised, Hadri (2000), also assumes full heterogeneity but has stationarity as its null hypothesis. These three procedures take account of the dependence between the series that make up the panel by subtracting the means obtained for each time period across cross sections, from the observations. On the other hand, the remaining four procedures, due to Taylor and Sarno (TS) (1998), Breuer, McNown and Wallace (BMW) (2001), Pesaran (P) (2007) and Bai and Ng (BN) (2004a) handle the problem of dependence in a somewhat more elaborate manner. TS and BMW do this by considering the autoregressions corresponding to each series as set of seemingly unrelated regressions. TS consider a joint test of a unit root while BMW consider individual tests, thereby complementing each other. P and BN, on the other hand, assume that there is a common factor in the panel of series. P adds this common factor, proxied by the time-wise mean, as a regressor to the autoregressions and performs the ADF test while BN decompose the series into this common factor and the idiosyncratic components and test for a unit root in both components, thereby enabling us to determine the source of the persistence if it exists.

Of these seven procedures, LLC and IPS lead to the rejection of the null hypothesis of a unit root, while Hadri, TS and BMW do not. The LLC result has the, rather sharp, implication that all 17 series are stationary which, obviously, is not realistic. The IPS result, on the other hand, implies that, at least one series is stationary. This is corroborated by individual ADF tests for, say, the UK, Italy, France, the Netherlands and Belgium based series. The same corroboration is, however, lacking from the other panel approaches, implying that the evidence about the stationarity of the Turkish real exchange rate is mixed and not very strong if panel procedures are used alone as an alternative to univariate ADF tests. Structural shifts in the deterministic terms may need to be introduced into these procedures to obtain stronger evidence of stationarity but this is the subject of further research.

## 1. Introduction

Testing whether real exchange rates are stationary and, thereby, obtaining evidence on the absolute version of the purchasing power parity (PPP) hypothesis has, initially, been done by using the Augmented Dickey-Fuller (ADF) statistic to test for a unit root. Subsequently, to mitigate the low power of the ADF test, several alternatives have been used for the same purpose. [See, e.g., Sarno and Taylor (2002) for a recent survey.]. Panel unit root testing is one of these alternatives.

The logic behind the use of a panel unit root test is to combine the information from time series with the information from cross-sectional units. The addition of cross-sectional variation to time series variation improves estimation efficiency, leading to smaller standard errors and, consequently, to higher t-ratios. Levin, Lin and Chu (LLC) (2002) show that, in situations where there is not enough time-series variation to produce good power in the ADF test, a relatively small amount of cross-section variation can result in substantial improvement.

Unit root tests have been applied to Turkish real exchange rates (RER) to test the absolute version of the PPP hypothesis. Erlat (2003) contains a survey of all (both unit root and cointegration based) evidence regarding the PPP hypothesis for Turkey. The results, usually, do not favour the PPP hypothesis, except when nonlinear time series methods are used as in Sarno (2000a and b). Erlat (2004) further checks out Sarno's findings using tests for unit roots where the alternative is nonlinear stationarity and concludes that nonlinear modeling of the Turkish real exchange rate depends upon the foreign currency used as a base and that linear models with multiple shifts in the deterministic terms, and fractional integration techniques with structural shifts, as implemented in Erlat (2003), may provide an alternative account of Sarno's findings. Erlat (2003)'s application of these models to the two primary bilateral Turkish real exchange rates; the $US and the German DM based rates, indicate that these two rates may, in fact, be taken to be stationary with significant long-memory components. These findings may not provide evidence in favour of the absolute PPP hypothesis in its purest form (where there is no trend term or structural shifts) but they do indicate that the absolute version of the "quasi" PPP hypothesis cannot be rejected for Turkey.

In this paper, I utilize panel procedures to see if they provide evidence in favour of the PPP hypothesis, not its "quasi" version; hence, structural shifts in the deterministic terms have not been taken into account in the present application. Panel procedures were first used on Turkish data by Özdemir (2002), which is her M.S. thesis written under my supervision. As I shall discuss below, the first generation of panel procedures, LLC (2002), Im, Pesaran and Shin (2003), Maddala and Wu (MW) (1999), Choi (2001) and Hadri (2000), among others, are, in general, based on the assumption that the series that make up the panel are independent of each other, which, of course, is hardly a realistic assumption to make where exchange rates are concerned. A common way to deal with this problem has been to subtract the means obtained for each time point across cross-sections, from the observations. An alternative, due to Taylor and Sarno (1998) and Breuer, McNown and Wallace (2001, 2002), handles the problem of dependence by considering the autoregressions corresponding to each series as a set of seemingly unrelated regressions. Taylor and Sarno consider a joint test of a unit root while Breuer et al. consider individual tests, thereby complementing each other.

Özdemir (2002) contains the results of applying most of these procedures to a panel of seventeen monthly Turkish real exchange rates that cover the period 1984.01-2001.06. In this paper I, in addition, implement two new procedures to account for the dependence between the series due to Bai and Ng (2004a and b) and Pesaran (2005). These procedures are based on the notion that the time series that make up the panel have a common component but differ as to how this common component is treated. Bai and Ng decompose the actual time series into their common and idiosyncratic components and apply tests of unit roots to these components separately. One can also apply the panel unit root tests mentioned above to the idiosyncratic components since they will now be asymptotically independent. Pesaran, on the other hand, does the same decomposition for the disturbance terms of the autoregressions used to test for a unit root. He estimates the common component by the average over the cross-sectional units and adds its lag, its first difference together with its lags as additional regressors to the autoregression mentioned above.

Thus, the plan of the paper will be as follows: In the next section I shall give an account of the panel procedures utilized. Subsequently, in Section 3 I shall describe our data and, in Section 4, present the empirical results. The final section will contain the conclusions.

## 2. Panel Unit Root Tests

### a. The First Generation Procedures

I shall be interested in testing the presence of a unit root in a panel of real exchange rates, the natural log of which I shall denote by $q_{it}$ and define as

$$(1) \qquad q_{it} = e_{it} + p_{it}^{*} - p_{it}$$

where $e_{it}$ denotes the logarithm of the nominal exchange rate of Turkey with its $i^{th}$ trading partner (expressed as TL/Foreign Currency), $p_{it}^{*}$, the logarithm of the $i^{th}$ trading partner's price level and $p_t$, the log of the domestic price level. I shall discuss the LLC, IPS, MW, Choi and Hadri approaches to this problem.

For the LLC, IPS, MW and Choi approaches, I shall start by considering the autoregressions used to obtain the ADF test for each time series in the panel. Let there be N such series. Then,

$$(2) \qquad \Delta q_{it} = \beta_{ir}' d_{tr} + \alpha_i q_{i,t-1} + \sum_{j=1}^{p_i} \gamma_{ij} \Delta q_{i,t-j} + \varepsilon_{it}, \quad i = 1,\dots,N;\ r = 0,1,2$$

where $d_{t0} = 0$ or $d_{t1} = 1$ or $d_{t2} = (1, t)'$. Note that I allow for different configurations of the deterministic term and different lag lengths for each series. The choice of each $p_i$ may be done by using a general-to-specific procedure based on either information criteria, such as AIC or the Schwartz criterion, or on sequentially testing the last coefficient of the $\Delta q_{i,t-j}$.

After deciding upon the $p_i$ and the $d_{tr}$, the *first* step in the ***LLC approach*** is to control for the differences in the variances of the $\varepsilon_{it}$. For this purpose, the equations in (2) are estimated in two steps. First, $\Delta q_{i,t}$ and $q_{i,t-1}$ are regressed on the $\Delta q_{i,t-j}$ and $d_{tr}$ to yield the residuals $\hat{e}_{it} = \Delta q_{it} - \hat{\beta}_{ir}' d_{tr} - \sum_{j=1}^{p_i} \hat{\pi}_{ij} \Delta q_{i,t-j}$ and $\hat{v}_{i,t-1} = q_{i,t-1} - \hat{\beta}_{ir}' d_{tr} - \sum_{j=1}^{p_i} \hat{\pi}_{ij} \Delta q_{i,t-j}$. These residuals are used to estimate the $\alpha_i$ from $\hat{e}_{it} = \alpha_i \hat{v}_{i,t-1} + \eta_{it}$ and are corrected for heterogeneity as $\tilde{e}_{it} = \hat{e}_{it} / \hat{\sigma}_{\varepsilon i}$ and $\tilde{v}_{i,t-1} = \hat{v}_{i,t-1} / \hat{\sigma}_{\varepsilon i}$ where

(3) $$\hat{\sigma}_{\varepsilon i}^2 = \frac{1}{T - p_i - 1} \sum_{t=p_i+2}^{T} (\hat{e}_{it} - \hat{\alpha}_i \hat{v}_{i,t-1})^2$$

The *second* step in this approach is to compute the panel test statistic. This statistic assumes that, as opposed to the formulation in (2), all the $\alpha_i$ have a common value, $\alpha$, so that the null hypothesis to be tested is

$$H_0: \alpha = 0 \qquad \text{vs.} \qquad H_1: \alpha < 0.$$

Thus, I need an estimate of this common coefficient $\alpha$ and its t-ratio. The estimator of $\alpha$ will be obtained from the pooled regression

$$\tilde{e}_{it} = \alpha \tilde{v}_{i,t-1} + \varepsilon_{it}, \quad i = 1,...,N;\ t = p_i + 2,...,T$$

where the total number of observations is $N\tilde{T}$ and $\tilde{T} = T - \bar{p} - 1$ with $\bar{p} = \sum_{i=1}^{N} p_i / N$. Thus, I obtain

$$\hat{\alpha} = \frac{\sum_{i=1}^{N} \sum_{t=p_i+2}^{T} \tilde{v}_{i,t-1} \tilde{e}_{it}}{\sum_{i=1}^{N} \sum_{t=p_i+2}^{T} \tilde{v}_{i,t-1}^2} \quad \text{and} \quad \hat{\sigma}_{\hat{\alpha}} = \frac{\hat{\sigma}_{\tilde{\varepsilon}}}{\left[ \sum_{i=1}^{N} \sum_{t=p_i+2}^{T} \tilde{v}_{i,t-1}^2 \right]^{1/2}} \quad \text{where} \quad \hat{\sigma}_{\tilde{\varepsilon}}^2 = \frac{1}{N\tilde{T}} \sum_{i=1}^{N} \sum_{t=p_i+2}^{T} (\tilde{e}_{it} - \hat{\alpha} \tilde{v}_{i,t-1})^2, \quad \text{and}$$

calculate the t-ratio as $t_\alpha = \hat{\alpha} / \hat{\sigma}_{\hat{\alpha}}$.

The *final* step in this approach is to adjust $t_\alpha$ so that, asymptotically, it has the standard normal distribution under the null hypothesis. But, one of the components needed for this adjustment is the ratio of the long-run and short-run standard deviations of each of the series. The short-run standard deviation will be given by the $\hat{\sigma}_{\varepsilon i}^2$ of (3) above. The long-run variances, on the other hand, are estimated as,

(4) $$\hat{\sigma}_{q_i}^2 = \frac{1}{T-1} \sum_{t=2}^{T} (\Delta q_{it} - \overline{\Delta q_i})^2 + 2 \sum_{j=1}^{\bar{k}} w_{\bar{k}j} \left( \frac{1}{T-1} \sum_{t=j+2}^{T} (\Delta q_{it} - \overline{\Delta q_i})(\Delta q_{i,t-j} - \overline{\Delta q_i}) \right)$$

for $d_{t1}$ and $d_{t2}$, but, in the case of $d_{t2}$, the series are first detrended. The $w_{\bar{k}j}$ are weights used to ensure that the $\hat{\sigma}^2_{q_i}$ are always positive. In these applications, I follow LLC in using the Bartlett weights, which may be expressed as $w_{\bar{k}j} = 1 - (j/(\bar{k}+1))$. Having obtained the $\hat{\sigma}^2_{q_i}$, I now form the ratios, $\hat{s}_i = \hat{\sigma}_{q_i} / \hat{\sigma}_{ei}$ and calculate their average, $\hat{S}_N = \sum_{i=1}^{N} \hat{s}_i / N$.

The adjustment to $t_\alpha$ may now be done as

$$t^*_\alpha = \frac{t_\alpha - N\tilde{T}\hat{S}_N(\hat{\sigma}_{\hat{\alpha}} / \hat{\sigma}^2_{\tilde{\varepsilon}})\mu^*_{rT}}{\sigma^*_{rT}} \qquad (5)$$

where the mean-adjustment, $\mu_{rT}{}^*$, and the standard-deviation adjustment, $\sigma_{rT}{}^*$, have been calculated by LLC and are given in Table 2 of that paper. They are presented for each deterministic configuration (r = 0, 1, 2) and for $\tilde{T}$. They also suggest, for each $\tilde{T}$, lag truncation parameters, $\bar{k}$, to be used in obtaining the $\hat{\sigma}^2_{q_i}$. I chose $\bar{k}$ from that Table. As I mentioned above, $t_\alpha{}^*$ will, asymptotically, be N(0,1) under $H_0$.

The starting point of the ***IPS approach*** is also the ADF regressions given in (2). But, the null and alternative hypotheses are different from that of the LLC approach, where the rejection of the null hypothesis implies that *all* the series are stationary. I now have

$H_0$: $\alpha_1 = \alpha_2 = \ldots = \alpha_N = 0$ vs. $H_1$: Some but not necessarily all $\alpha_i < 0$

The test statistic itself is rather simple to compute. Again, after deciding upon $d_{tr}$ and the $p_i$, I obtain the t-ratios for the $\alpha_i$, $t_{\alpha_i}$, and calculate their arithmetic average, $\bar{t}_{NT} = \sum_{i=1}^{N} t_{\alpha_i} / N$. IPS show that $\bar{t}_{NT}$ may be adjusted to yield an asymptotic N(0, 1) statistic under the null hypothesis;

$$\bar{t}^*_{NT} = \frac{N^{1/2}\left(\bar{t}_{NT} - N^{-1}\sum_{i=1}^{N} E(t_{\alpha_i})\right)}{\left[N^{-1}\sum_{i=1}^{N} Var(t_{\alpha_i})\right]^{1/2}} \qquad (6)$$

The $E(t_{\alpha_i})$ and $Var(t_{\alpha_i})$ have been obtained by simulation and are given in Table 1 of IPS.

The ***MW and Choi approaches*** have the same framework. The hypothesis they test is the same as in the IPS case. But, instead of averaging the individual ADF statistics, they aggregate their p-values. Denoting these p-values by $\pi_i$, the statistic proposed may be expressed as

(7) $$P = -2\sum_{i=1}^{N} \ln \pi_i$$

Under the null hypothesis $P$ is distributed asymptotically as $\chi^2$ with $2N$ degrees of freedom. This result is obtained as $T \rightarrow \infty$ while $N$ is taken to be fixed. When N also tends to infinity, Choi (2001) shows that P may be standardized as

(8) $$P_m = \frac{1}{2\sqrt{N}}\sum_{i=1}^{N}(-2\ln \pi_i - 2) = -\frac{1}{\sqrt{N}}\sum_{i=1}^{N}(\ln \pi_i + 1)$$

to have an asymptotic N(0,1) distribution.[1]

Choi (2001) also suggests an alternative test for the case where N is finite:

(9) $$Z = \frac{1}{\sqrt{N}}\sum_{i=1}^{N}\Phi^{-1}(\pi_i)$$

where $\Phi$ is the standard normal cumulative distribution function. $Z$ is asymptotically N(0,1) when $T \rightarrow \infty$. He also shows that $Z$ has the same asymptotic distribution when $N$ also tends to infinity.

I described this approach in terms of the p-values for the ADF statistic. It can, however, be used for any test of a unit root as long as its p-value is obtained, but that is where the difficulty lies. These p-values need to be simulated and this can be a tall order for most cases, particularly when N is large. Such p-values are readily available for the ADF tests (MacKinnon, 1996) and our applications will be based on them.

[1] This follows from the fact that $-2\ln \pi_i \sim \chi_2^2$ so that $E(-2\ln \pi_i) = 2$ and $Var(-2\ln \pi_i) = 4$.

Finally, in the case of the ***Hadri approach***, the null hypothesis is the stationarity of the series instead of nonstationarity. The framework is the one dealt with in Kwiatowski et al. (KPSS) (1992) for a single series. The models may now be expressed as,

$$q_{it} = \beta_{irt}' d_{rt} + \varepsilon_{it}, \qquad i = 1,\dots,N; \quad r = 1,2 \tag{10}$$

where $\beta_{irt} = \beta_{i1t}$ when r = 1 and $\beta_{irt} = (\beta_{i1t}, \beta_i)'$ when r = 2. I assume that the intercept, $\beta_{i1t}$, is generated by a random walk, $\beta_{i1t} = \beta_{i1,t-1} + u_{it}$, where $E(u_{it}) = 0$ and $E(u_{it}^2) = \sigma_u^2 \geq 0$. In other words, I assume that the variances of the $u_{it}$ are the same for every series. Thus, the hypothesis to be tested becomes,

$$H_0: \sigma_u^2 = 0 \qquad \text{vs.} \qquad H_1: \sigma_u^2 > 0$$

This hypothesis may be tested under two different assumptions concerning the variances of the $\varepsilon_{it}$. If they are assumed to be the same for every time series in the panel, then the statistic may be obtained as

$$LM_s = \frac{N^{-1}\sum_{i=1}^{N} T^{-2} \sum_{t=1}^{T} S_{it}^2}{\hat{\sigma}_\varepsilon^2}, \qquad s = \mu, \tau \tag{11}$$

where $S_{it} = \sum_{j=1}^{t} \hat{\varepsilon}_{ij}$ and $\hat{\sigma}_\varepsilon^2 = N^{-1}\sum_{i=1}^{N}\left(T^{-1}\sum_{t=1}^{T}\hat{\varepsilon}_{it}^2 + 2\sum_{j=1}^{\bar{k}} w_{\bar{k}j}\left(T^{-1}\sum_{t=j+1}^{T}\hat{\varepsilon}_{it}\hat{\varepsilon}_{i,t-j}\right)\right)$. Note that the long-run variance of the $\varepsilon_{it}$ is estimated using the Bartlett weights (as in equation (4) above) in order to take into account the autocorrelation in the $\varepsilon_{it}$. The subscript s=μ refers to r = 1 and s = τ refers to r = 2, from equation (10). The $\hat{\varepsilon}_{it}$, of course, come from the OLS estimation of equation (10). I shall refer to this version of the test as *Hadri 1* when presenting the empirical results.

Note that $\hat{\sigma}_\varepsilon^2$ is, in fact, the average of the estimated long-run disturbance variances for each $\varepsilon_{it}$, the $\hat{\sigma}_{\varepsilon_i}^2$. Since the KPSS statistic for each series in the panel may be obtained as,

(12) $$LM_{si} = \frac{T^{-2}\sum_{t=1}^{T} S_{it}^2}{\hat{\sigma}_{\varepsilon_i}^2}, \qquad s = \mu, \tau; \quad i = 1, \ldots N$$

the statistic in (11) may be regarded as being obtained by averaging the numerators and the denominators of the $LM_{si}$ separately. If I assume that the $\sigma_{\varepsilon_i}^2$ are, in fact, different for each time series, then the test statistic may be obtained as the arithmetic average of the $LM_{si}$; $LM_s = \sum_{i=1}^{N} LM_{si} / N$ and referred to as *Hadri 2*. Hadri (2000) shows that, for both cases,

(13) $$Z_s = \frac{N^{1/2}(LM_s - \mu_s)}{\sigma_s} \overset{A}{\sim} N(0,1), \qquad s = \mu, \tau$$

where $\mu_\mu = 1/6$, $\sigma_\mu^2 = 1/45$ and $\mu_\tau = 1/15$, $\sigma_\tau^2 = 11/6300$.

**b. Dealing with the Problem of Dependence**

The problem of dependence between the series that make-up the panel has several implications: (i) As O'Connell (1998) showed, panel unit root tests will over reject the null hypothesis of a unit root; there will be an upward bias in the size of the tests, giving the impression of high power. Such distortions in size will come about, particularly, if the dependence is due to cross-unit cointegration (Banerjee, Marcellino and Osbat, 2005). (ii) If the unit root null were not rejected, this would imply that there exist N independent unit roots. But, if these series have common stochastic trends, the number of unit roots would be less than N (Bai and Ng, 2004b). The procedures I am going to discuss in this subsection are designed to remove this dependence so that most, if not all, of these implications no longer hold.

The ***first*** solution to deal with the problem of dependence was implemented by LLC and IPS. They assume that, in addition to a series specific intercept and/or trend term as given in (2), there is a time specific intercept that may be estimated by taking the average across the series at each point in time. In other words, this dependence is accounted for by calculating

$\overline{q}_t = \sum_{i=1}^{N} q_{it} / N, \quad t = 1,\ldots,T$ , and subtracting it from each cross-sectional observation at point *t*; namely, for each *t*, using $q_{it} - \overline{q}_t$ instead of $q_{it}$ in the calculations given above. This correction will not remove the correlation between the series, but, as Luintel (2001) demonstrates, it may reduce it considerably.

The ***second*** solution would be to assume, at the outset, that the $\varepsilon_{it}$ of (2) are contemporaneously correlated so that the *N* equations involved may be treated as a set of seemingly unrelated regressions (SUR). In the case where $T > N$[2], such an approach is taken by Taylor and Sarno (1998), Groen (2000) and Breuer, McNown and Wallace (2001, 2002).[3] The first two consider testing the joint null hypothesis

$$H_0: \alpha_1 = \alpha_2 = \ldots = \alpha_N = 0$$

while Breuer et al. test the individual hypotheses

$$H_{0i}: \alpha_i = 0, \qquad i = 1,\ldots,N$$

Taylor and Sarno (1998) use the two-step Estimated GLS (EGLS) procedure to estimate the system of equations in (2) and test the joint null hypothesis using the Wald statistic. Groen (2000), on the other hand, estimates the system by maximum likelihood and uses the likelihood ratio statistic to test the same hypothesis. I preferred to implement Taylor and Sarno (1998)'s approach since it is also the one taken by Breuer et al. (2001) and the two tests complement each other.[4]

Formally, the $i^{th}$ equation in (2) may be expressed as $\Delta q_i = Z_i \delta_i + \varepsilon_i$ where $Z_i = [\, q_{i,-1}, D_r, \Delta Q_i \,]$, $\delta_i = (\, \alpha_i, \beta_{ir}', \gamma_i' \,)'$ and

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[2] When $T < N$, the estimate of the disturbance covariance matrix cannot be inverted so that the procedures discussed below cannot be implemented. However, one may estimate the SUR system by OLS and use the systems covariance matrix for the coefficient estimators to obtain the standard errors since, now, the disturbance covariance matrix need not be inverted. Breitung and Das (2005) and Jonsson (2005) discuss such a procedure for the case of a common $\alpha_i$.

[3] In earlier work, restricted versions of the SUR system were used where either the $\alpha_i$ were taken to be equal to a common value (Abuaf and Jorion (1990), Jorion and Sweeney (1996), O'Connell (1998)) and/or the lag length, $p_i$, was either set to a common non-zero value for all equations (O'Connell, 1998) or to zero (Flores et al., 1999). Higgins and Zakrajsek (2000) come closest to the models discussed above, with only the $\alpha_i$ restricted to be the same across equations.

[4] Kao and Mikola (2001) have shown that the models in (10) could also be treated as a system of equations, leading to a multivariate generalization of the Hadri approach. I decided not to implement this generalization due, partly, to its heavy computational burden and, also, because the fourth approach we shall discuss below provides us with a simpler way of dealing with the dependence problem in the Hadri or KPSS context.

$$\Delta q_i = \begin{bmatrix} \Delta q_{i,p_{max}+1} \\ \vdots \\ \Delta q_{i,T-1} \end{bmatrix}, \quad q_{i,-1} = \begin{bmatrix} q_{i,p_{max}+1} \\ \vdots \\ q_{i,T-1} \end{bmatrix}, \quad \Delta Q_i = \begin{bmatrix} \Delta q_{i,p_{max}} & \cdots & \Delta q_{i,p_{max}-p_i+1} \\ \vdots & \ddots & \vdots \\ \Delta q_{i,T-2} & \cdots & \Delta q_{i,T-p_i-1} \end{bmatrix},$$

$$\varepsilon_i = \begin{bmatrix} \varepsilon_{i,p_{max}+2} \\ \vdots \\ \varepsilon_{i,T} \end{bmatrix}, \quad D_r = \begin{bmatrix} d_{p_{max}+2,r}' \\ \vdots \\ d_{T,r}' \end{bmatrix}, \quad r = 0,1,2$$

where $p_{max}$ is the largest $p_i$ in the system of equations. Stacking these vectors and matrices, we may express the $N$-equation system as

(13) $$\Delta q = Z\delta + \varepsilon$$

where $\Delta q = [\, \Delta q_1', \ldots, \Delta q_N' \,]'$, $Z = diag[\, Z_1, \ldots, Z_N \,]$, $\delta = (\, \delta_1', \ldots, \delta_N' \,)'$ and $\varepsilon = (\, \varepsilon_1', \ldots, \varepsilon_N' \,)'$ with $\varepsilon \sim N(\, 0, \Omega \otimes I_{T^*} \,)$, $\Omega$ being the $N$x$N$ contemporaneous variance-covariance matrix of the disturbances, $T^* = T\text{-}p_{max}\text{-}1$[5] and "⊗" denoting the Kroenecker product of two matrices.

The null hypothesis for the joint test may now be expressed as $H_0$: $R\delta = 0$ where $R = diag(\, r_1', \ldots, r_N' \,)$ and $r_i' = (\, 1,0,\ldots,0 \,)_{1x(\, p_i+r+1\,)}$ and the Wald statistic, as formulated by Taylor and Sarno (1998) and called the Multivariate ADF (MADF) statistic, may be stated as[6]

(14) $$MADF = \frac{(\, NT^* \,)\hat{\delta}' R' [\, R(\, Z'(\, \hat{\Omega}^{-1} \otimes I_{T^*} \,)Z \,)^{-1} R' \,]^{-1} R\hat{\delta}}{(\, \Delta q - Z\hat{\delta} \,)'((\, \hat{\Omega}^{-1} \otimes I_{T^*} \,)(\, \Delta q - Z\hat{\delta} \,)}$$

Now, Breuer et al. (2001) also estimate the same equations as in (13) but use the individual significance tests for the $\alpha_i$. They call the corresponding t-ratios, the SURADF statistics. These may be regarded as complements to the MADF test as they would indicate which series are stationary when a MADF test rejects the joint null hypothesis.

[5] Taylor and Sarno (1998) take the $p_i$ to be equal to the same value, but there is no need to do that if the sample is set equal to $T^*$, as we have done above. See also Groen (2000) and Breuer et al. (2001).

[6] This is not the only formulation of the Wald statistic one may use for a SUR system. Strictly speaking, the more appropriate formulation, as used by Ho (2002) to test the same null hypothesis above, would be $W = \hat{\delta}' R' [\, R(\, Z'(\, \hat{\Omega}^{-1} \otimes I_{T^*} \,)Z \,)^{-1} R' \,]^{-1} R\hat{\delta}$. The expression in (14) is based on the F-statistic formulation discussed by Zellner (1962).

For the MADF and SURADF tests, theoretically derived asymptotic null distributions are not available. The desired critical values are generated using Monte Carlo methods and are, therefore, case specific.

The ***third*** solution to the dependence problem is provided by Pesaran (2007). In order to see what is involved, consider a simple version of (2) with only a constant term:

(15) $$\Delta q_{it} = \beta_{i1} + \alpha_i q_{i,t-1} + \varepsilon_{it}, \quad i = 1,...,N$$

where it is assumed that

(16) $$\varepsilon_{it} = \eta_i f_t + u_{it}$$

in which $f_t$ is the unobserved common effect and $u_{it}$ is the individual-specific or idiosyncratic disturbance. Combining (15) and (16) yields

(17) $$\Delta q_{it} = \beta_{i1} + \alpha_i q_{i,t-1} + \eta_i f_t + u_{it}$$

In order to estimate (17) by OLS, an observable counterpart for $f_t$ is required. To simplify things, assume that in (17), $\beta_{i1} = \beta_1$, $\alpha_i = \alpha$ and $\eta_i = \eta$. Then, take the average of (17) over $i$ to yield $\Delta\bar{q}_{it} = \beta_1 + \alpha\bar{q}_{t-1} + \eta f_t + \bar{u}_t$. Now, it is shown by Pesaran (2006), that if $\bar{u}_t$ converges in quadratic mean to 0, then $f_t$ will converge in probability to $\Delta\bar{q}_{it} - \beta_1 - \alpha\bar{q}_{t-1}$ as $N \to \infty$. Substituting it in (17) gives

(18) $$\Delta q_{it} = c_{i1} + \alpha_i q_{i,t-1} + \eta_i \Delta\bar{q}_{it} + \varphi_i \bar{q}_{t-1} + u_{it}$$

where $c_{i1} = \beta_{i1} - \eta_i \beta_1$ and $\varphi_i = -\eta_i \alpha$. Hence, the t-ratio of $\alpha_i$ obtained from the OLS estimation of (18) may be used to test for a unit root. The resultant statistic is called the *Cross-Sectionally Augmented ADF (CADF)* statistic. If there is autocorrelation in the $\varepsilon_{it}$, then (18) may be generalized to parallel (2) as,

(19) $$\Delta q_{it} = c_{ir}' d_{tr} + \alpha_i q_{i,t-1} + \sum_{j=1}^{p_i} \gamma_{ij} \Delta q_{i,t-j} + \varphi_i \bar{q}_{t-1} + \sum_{j=0}^{p_i} \eta_{ij} \Delta \bar{q}_{t-1} + u_{it}$$

The critical values of CADF have been generated by Monte Carlo and are tabulated in Pesaran (2007).

The panel version of this test, CIPS, is simply the arithmetic average of the individual CADF statistics:

(20) $$CIPS = \frac{\sum_{i=1}^{N} CADF_i}{N}$$

As opposed to the IPS statistic given in (6) above, this average is not standardized to provide a statistic which is N(0,1). Instead, as in the case of the individuals CADFs, critical values have been generated by Monte Carlo and are presented in Pesaran (2007).

The procedure above assumes that there is a single common factor and that it rests in the disturbance term. The ***fourth***[7] solution to the dependence problem, as provided by Bai and Ng (2004a), assumes that there is more than one common factor involved and that a time series may be decomposed into these common factors and its idiosyncratic component. Formally, the model is specified as

(21) $$\begin{aligned} q_{it} &= \beta_{ir}' d_{tr} + \varphi_i' F_t + e_{it}, & t &= 1,\dots,T \\ F_{jt} &= \sum_{s=1}^{m_j} \alpha_{js} F_{j,t-s} + u_{jt}, & j &= 1,\dots,n \\ e_{it} &= \sum_{s=1}^{p_i} \rho_i e_{i,t-s} + \varepsilon_{it}, & i &= 1,\dots,N \end{aligned}$$

[7] There are other approaches based on a common factor specification. These are due to Choi (2002), Phillips and Sul (2003), and Moon and Perron (2004). A textbook account of these procedures may be found in Baltagi (2008) and a detailed survey is given by Hurlin and Mignon (2004). These tests are all based on removing the common component(s) from the data before performing unit root tests. None of them do it as simply as the Pesaran approach and none of them test the common component(s) for a unit root. This is why I chose to apply only the Pesaran and Bai-Ng approaches.

where $F_t$ is an $n$x$1$ vector of common factors, each element of which has an AR($m_j$) structure and $e_{it}$ is the idiosyncratic component exhibiting an AR($p_i$) structure. The $n$x$1$ vector $\varphi_i$ contains the factor loadings. The setup is roughly similar to the first solution to the dependence problem where the $\overline{q}_t$ were subtracted from each observation in a series and the panel tests were applied to the adjusted series which were expected to be less dependent. In the present case, one obtains estimates of $F_t$ and the $e_{it}$ and tests for unit roots in $F_t$ and the $e_{it}$ separately so that the source of the presence or absence of a unit root in $q_{it}$ may be determined. Since the estimated $e_{it}$'s are expected to be asymptotically independent, the panel procedures described in Section 2.a may be applied to these series.

Bai and Ng (2004a) describe a procedure, based on principal components, for the case of $d_{1t}$ and $d_{2t}$, separately. Even though I shall consider both cases in the applications, I shall only describe the procedure for the $d_{2t}$ case. Hence, the model to be considered is the first difference of the model in (21) for r = 2.[8] $\Delta q_{it} = \beta_{1i} + \varphi_i{}'\Delta F_t + \Delta e_{it}, \quad t = 2,\ldots,T$. It is put in mean-deviation form to yield

(22) $$\Delta q_{it} - \overline{\Delta q_i} = \varphi_i{}'(\Delta F_t - \overline{\Delta F}) + (\Delta e_{it} - \overline{\Delta e_i}), \quad t = 2,\ldots,T$$

where, e.g., $\overline{\Delta q_i} = \sum_{t=2}^{T} \Delta q_{it} /(T-1)$. The steps of the procedure may then be stated as follows:

i. Form the matrices

$$Q = \begin{bmatrix} \Delta q_{12} - \overline{\Delta q_1} & \cdots & \Delta q_{N2} - \overline{\Delta q_N} \\ \vdots & \ddots & \vdots \\ \Delta q_{1T} - \overline{\Delta q_1} & \cdots & \Delta q_{NT} - \overline{\Delta q_N} \end{bmatrix} \text{ and } F = \begin{bmatrix} \Delta F_{12} - \overline{\Delta F_1} & \cdots & \Delta F_{n2} - \overline{\Delta F_n} \\ \vdots & \ddots & \vdots \\ \Delta F_{1T} - \overline{\Delta F_1} & \cdots & \Delta F_{nT} - \overline{\Delta F_n} \end{bmatrix},$$

and estimate $F$ by forming the $(T\text{-}1)$x$(T\text{-}1)$ cross-product matrix $QQ'$ and obtaining the $n$ eigenvectors (multiplied by $(T\text{-}1)^{1/2}$) corresponding to the first $n$ largest eigenvalues of $QQ'$. The estimated loading matrix will be obtained as $\hat{\Phi} = Q'\hat{F}/(T-1)$.

[8] In the case of r = 1, the constant only case, first-differencing will yield $\Delta q_{it} = \varphi_i{}'\Delta F_t + \Delta e_{it}$ and the data used in the succeeding steps would be in first-difference form.

ii. Set $f_t = \Delta F_t - \overline{\Delta F}$ and obtain $\hat{F}_{jt} = \sum_{s=2}^{t} \hat{f}_{js}$. How we treat the $\hat{F}_{jt}$ differs according to whether there is only one ($n$ = *1*) or more ($n$ > *1*) common factors. The treatment of the latter case is quite involved and since only one factor was obtained in the applications, I shall describe the former case alone. Hence, when $n$ = *1*, test for a unit root in $\hat{F}_{1t}$ by including an intercept and trend term (or only an intercept if r =1) in the autoregression.

iii. Set $\hat{z}_{it} = (\Delta q_{it} - \overline{\Delta q_i}) - \hat{\varphi}_i' \hat{f}_t$ and obtain $\hat{e}_{it} = \sum_{s=2}^{t} \hat{z}_{is}$, $i = 1, \ldots, N$. Then, test for a unit root in each $\hat{e}_{it}$ without including an intercept and trend term.

One may test for unit roots in $\hat{F}_{1t}$ and the $\hat{e}_{it}$ using the ADF or any other statistic that has the unit root as a null. The distribution of the ADF test when applied to $\hat{F}_{1t}$ remains the same as when it is applied to the $q_{it}$. Its distribution, when applied to the $\hat{e}_{it}$, however, is now given by the distribution of the LM test of a unit root as developed by Schmidt and Lee (1991). But, note that this result is not affected by whether $\hat{F}_{1t}$ is I(1) or I(0). One may also implement the first generation panel procedures using the $\hat{e}_{it}$.

If one wishes to test the null hypothesis of stationarity, one may use the KPSS statistic to test $H_0$ for $\hat{F}_{1t}$ with $d_{1t}$ or $d_{2t}$ as the deterministic specification. If $\hat{F}_{1t}$ is found to be I(0), then one regresses the $\hat{e}_{it}$ on a constant if r = 1, and on a constant and a time trend if r = 2, and applies the KPSS statistic to the residuals, $\hat{e}_{it}^{0}$, from these regressions. If $\hat{F}_{1t}$ is found to be I(1), then the residuals to which the KPSS test will be applied will be obtained from the regression of $\hat{e}_{it}$ on a constant and $\hat{F}_{1t}$ if r =1, and on a constant, a time trend and $\hat{F}_{1t}$ if r = 2. These residuals will be denoted by $\hat{e}_{it}^{1}$. Bai and Ng (2004b) show that the KPSS statistics to test stationarity in $\hat{F}_{1t}$ and the $\hat{e}_{it}^{0}$ have the distributions derived in Kwiatowski et al (1992) but that the KPSS statistic to test stationarity in the $\hat{e}_{it}^{1}$ has the distribution of the statistic developed by Shin (1994) for testing the null of cointegration between I(1) variables. Bai and Ng (2004b) also point out that the $\hat{e}_{it}^{0}$ are asymptotically independent while the $\hat{e}_{it}^{1}$ are not, so that panel procedures can only be applied to the $\hat{e}_{it}^{0}$. Thus, the Hadri approach may only be implemented if we end up obtaining the $\hat{e}_{it}^{0}$ in our applications.

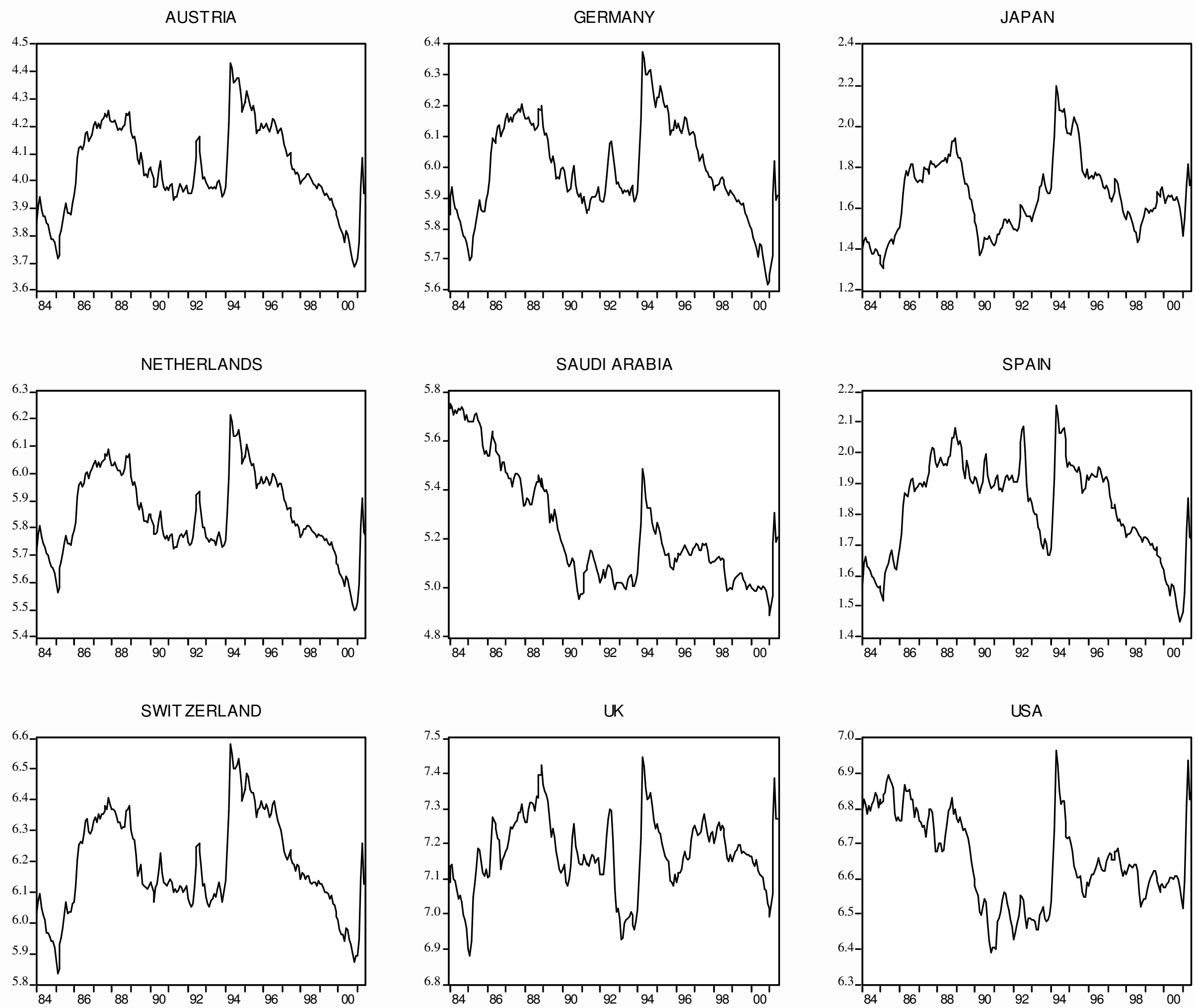
### 3. The Data

A panel of real exchange rates with Turkey's seventeen major trading partners, namely, Austria, Belgium, Denmark, Finland, France, Germany, Greece, Italy, Japan, the Netherlands, Norway, Saudi Arabia, Spain, Sweden, Switzerland, the UK and the USA, was constructed. The choice of trading partners was dictated by (a) the share they had in Turkey's total trade, (b) data availability, and (c) the desire to benefit from the added heterogeneity that a larger panel may provide. It was found that these seventeen countries account, on the average, for 64.5% of Turkey's trade for the period 1989-2001. Important trading partners such as Russia (with an average share of 5%) and Iran (1.8%) had to be left out because price and/or exchange rate data were not available. On the other hand, relatively smaller trading partners, such as Denmark (0.52%), Finland (0.52%) and Greece (0.81%) were included to increase the heterogeneity in the panel.

The series are monthly and cover the period 1984.01-2001.06. The price index used in the construction of the series is the Consumer Price Index (1987=100). The exchange rates and the domestic CPI series were obtained from the Central Bank database. The foreign CPIs were downloaded from the International Financial Statistics database and their base years were shifted to 1987.

To give an idea as to what to expect from the empirical results presented in the next section, in Figure 1 I plot the Turkish real exchange rate with nine trading partners. Austria, Germany, the Netherlands, Spain are European Union (EU) countries and note that

Figure 1
Plots of the Turkish Real Exchange Rate With Selected Trading Partners




the plots of the Turkish RERs are very similar for these countries. The UK, on the other hand, is also an EU country but the plot of its RER with Turkey differs somewhat from the other four. However, the RER with Switzerland, who is a non-EU European country, is quite similar to the series for these four EU countries. This is also true of the RERs with the other EU countries in the panel. I further note that the RERs with non-European countries; namely, Japan, Saudi Arabia and the USA, show very different patterns, but such countries constitute a minority in the panel. Thus, it would not be surprising to find a very strong dependence between the series that make up this particular panel.

## 4. Empirical Results

Let us start by presenting the unit root tests on the individual series. The tests are the ADF and KPSS tests. Since the objective is to test the PPP hypothesis, the equations needed for both tests should only contain intercepts. However, previous work (Erlat, 2003 and 2004) has shown that adding a trend term is warranted if the objective is the broader one of testing the persistence in the Turkish RERs. Thus, the results given below will be based on equations that contain only an intercept and both an intercept and a linear trend term.

In this and future applications of the ADF statistic, the lag length, $p_i$, was chosen using three criteria: AIC, Schwartz Information Criterion (SIC) and the t-ratio for the coefficient of the last lag. A general-to-specific procedure was implemented, starting with an equation for which a large enough lag length, $p_{max}$, was specified. In all applications, $p_{\max}$ was chosen to be 13. Following Erlat (2002), initially agreement was sought between, at least, two of the criteria. If there was no agreement, then the result of the criterion indicating the largest lag was chosen. For this choice of $p_i$, autocorrelation in the residuals was tested using the Ljung-Box statistic and if significant autocorrelation was found, $p_i$ was increased until it was eliminated.[9]

For the KPSS statistic, the number of weights, $\bar{k}$, (see equation (4) above) was decided upon by using a procedure developed by Newey and West (1994). This procedure is rather complex and I shall not attempt a simplified description here.

The results of the ADF and KPSS tests are given in Table 1. Note that, for the *intercept-only* model, the ADF tests reject the unit root null only for eight series; Belgium, Denmark, Greece, Italy, Netherlands, Norway, the UK and the USA. The rejection for five of them is only at the 10% level. The rejection becomes a bit stronger for Greece and the USA at the 5% level and is strongest for the UK series at the 1% level. The KPSS results confirm the ADF results for Belgium, Denmark, Netherlands and the UK and add to these the series for Austria, France, Germany, Japan and Switzerland. Hence, the number of series for which the PPP hypothesis is supported is almost the same for both tests but only four of these series are in common for both tests. The fact that the KPSS statistic indicates stationarity for the series not picked up by the ADF statistic may be viewed as a useful result, given that the power of the ADF is

[9] The Ljung-Box results are not presented but may be obtained from the author upon request.

**Table 1**

**ADF and KPSS Tests Results**

| | Intercept | | | | Intercept and Trend | | | |
|---|---|---|---|---|---|---|---|---|
| | **p** | **ADF** | $\bar{k}$ | **KPSS** | **p** | **ADF** | $\bar{k}$ | **KPSS** |
| Austria | 2 | -2.155 (0.224)[1] | 11 | 0.196 | 2 | -2.189 (0.493) | 11 | 0.191** |
| Belgium | 1 | -2.604 (0.094)* | 11 | 0.227 | 1 | -2.689 (0.243) | 11 | 0.187** |
| Denmark | 1 | -2.675 (0.080)* | 11 | 0.197 | 1 | -2.714 (0.232) | 11 | 0.183** |
| Finland | 1 | -2.094 (0.247) | 11 | 0.874*** | 1 | -2.876 (0.173) | 11 | 0.178** |
| France | 1 | -2.534 (0.109) | 11 | 0.306 | 1 | -2.736 (0.224) | 11 | 0.184** |
| Germany | 1 | -2.518 (0.113) | 11 | 0.208 | 1 | -2.579 (0.291) | 11 | 0.178** |
| Greece | 1 | -2.946 (0.042)** | 11 | 0.350* | 1 | -2.980 (0.140) | 11 | 0.191** |
| Italy | 1 | -2.741 (0.069)* | 11 | 0.637** | 1 | -3.282 (0.072)* | 11 | 0.208** |
| Japan | 1 | -2.542 (0.107) | 11 | 0.178 | 1 | -2.541 (0.308) | 11 | 0.114 |
| Netherlands | 1 | -2.652 (0.084)* | 11 | 0.220 | 2 | -2.356 (0.402) | 11 | 0.158** |
| Norway | 1 | -2.785 (0.062)* | 11 | 0.607** | 1 | -3.196 (0.088)* | 11 | 0.158** |
| S. Arabia | 1 | -2.446 (0.131) | 11 | 1.289*** | 1 | -2.450 (0.353) | 11 | 0.326*** |
| Spain | 2 | -2.335 (0.162) | 11 | 0.370* | 2 | -2.507 (0.325) | 11 | 0.307*** |
| Sweden | 1 | -2.460 (0.127) | 11 | 0.745*** | 1 | -3.217 (0.084)* | 11 | 0.251*** |
| Switzerland | 1 | -2.492 (0.119) | 11 | 0.169 | 1 | -2.491 (0.332) | 11 | 0.169** |
| UK | 1 | -4.302 (0.001)*** | 10 | 0.087 | 1 | -4.302 (0.004)*** | 10 | 0.088 |
| USA | | -2.951 (0.041)** | 11 | 0.624** | 1 | -2.856 (0.179) | 10 | 0.271*** |

Notes:

1. The figures in parentheses are p-values obtained using MacKinnon (1996).
2. The critical values for the KPSS tests have been obtained from Table 1 of Kwiatowski et al (1992).

| | 0.10 | 0.05 | 0.01 |
|---|---|---|---|
| Intercept | 0.347 | 0.463 | 0.739 |
| Intercept and Trend | 0.119 | 0.146 | 0.216 |

3. "*" : significant at the 10% level.
   "**" : significant at the 5% level
   "***" : significant at the 1% level.

low. On the other hand, the fact that the KPSS statistic does not offer collaboration of the ADF results for Greece, Italy, Norway and the USA is not that surprising in view of Caner and Kilian (2001) where they show that the KPSS statistic tends to reject the stationarity null more often than it should.

Turning to the results for the model with *intercept* + *trend*, the number of series found to be stationary by the ADF test is reduced by half to Italy, Norway, Sweden and the UK. The Italy, Norway and UK series had been picked up before; the Sweden series is new. The KPSS results only confirm the stationarity of the UK series and add to it the RER with Japan. In other words, when a linear trend is added, support for the stationarity of the RER series is considerably reduced.

**Table 2**

**LLC, IPS, Maddala-Wu, Choi and Hadri Test Results**

| | **Intercept** | **Intercept and Trend** |
|---|---|---|
| LLC | -4.366 (0.000)*** | -5.360 (0.000)*** |
| IPS | -5.406 (0.000)*** | -3.424 (0.000)*** |
| P | 9.726 (0.000)*** | 12.693 (0.000)*** |
| $P_m$ | 7.239 (0.000)*** | 12.837 (0.000)*** |
| Z | 88.345 (0.000)*** | 60.446 (0.004)*** |
| Hadri 1 | 6.590 (0.000)*** | 3.207 (0.001)*** |
| Hadri 2 | -5.672 (0.000)*** | -3.535 (0.000)*** |

Notes:
1. The figures in parentheses are p-values. For LLC, IPS, Hadri 1 and 2, $P_m$ and Z, they are based on the standard normal distribution, while, for P, it is based on the $\chi^2_{2N}$ distribution.
2. "***" : significant at the 1% level.

Turning next to the results of the first generation panel unit root tests discussed in Section 2.a, namely, LLC, IPS, P, $P_m$, Z and Hadri 1 and 2. The results are given in Table 2. Note that all the tests with a unit root null reject the null hypothesis for both the intercept-only and the intercept + trend cases. The Hadri results do not corroborate this outcome as the stationarity null is strongly rejected for both cases. The Hadri result is consistent with the individual KPSS results for the intercept + trend case in Table 1 but the same cannot be said for the intercept-only case where the stationary series are in the majority. This also holds for the LLC, IPS, P, $P_m$ and Z results, particularly for the intercept + trend case. It now needs to be seen if the latter results, in particular, are due to the dependence between the series.

| Table 3 | | | | |
|---|---|---|---|---|
| **ADF, LLC, IPS, P, $P_m$ and Z Test Results for Demeaned Data** | | | | |
| | **Intercept** | | **Intercept and Trend** | |
| LLC | -2.214 (0.013)** | | -0.602 (0.273) | |
| IPS | -1.787 (0.047)** | | 0.699 (0.758) | |
| P | 41.564 (0.175) | | 24.248 (0.892) | |
| $P_m$ | 0.917 (0.180) | | -1.183 (0.882) | |
| Z | -1.748 (0.040)** | | 0.870 (0.808) | |
| | **p** | **ADF** | **p** | **ADF** |
| Austria | 7 | -2.240 (0193) | 1 | -1.115 (0.923) |
| Belgium | 3 | -2.126 (0.235) | 3 | -1.804 (0.699) |
| Denmark | 1 | -2.578 (0.099)* | 1 | -2.187 (0.494) |
| Finland | 12 | -1.782 (0.389) | 12 | -3.087 (0.112) |
| France | 3 | -1.952 (0.308) | 3 | -1.912 (0.645) |
| Germany | 1 | -1.714 (0.423) | 1 | -1.574 (0.800) |
| Greece | 12 | -0.931 (0.777) | 12 | -1.931 (0.634) |
| Italy | 4 | -1.481 (0.542) | 4 | -2.130 (0.526) |
| Japan | 1 | -2.180 (0.215) | 1 | -2.632 (0.267) |
| Netherlands | 1 | -2.221 (0.200) | 1 | -2.151 (0.514) |
| Norway | 1 | -2.405 (0.142) | 1 | -3.172 (0.093)* |
| S. Arabia | 1 | -2.656 (0.084)* | 1 | -1.429 (0.850) |
| Spain | 1 | -1.821 (0.369) | 1 | -1.594 (0.793) |
| Sweden | 1 | -1.005 (0.752) | 1 | -2.193 (0.490) |
| Switzerland | 3 | -2.140 (0.229) | 3 | -2.238 (0.466) |
| UK | 1 | -1.482 (0.541) | 1 | -2.204 (0.484) |
| USA | 1 | -1.435 (0.565) | 4 | -1.091 (0.928) |

Notes:

1. The figures in parentheses are p-values. The ones associated with the ADF test are obtained using MacKinnon (1996). For LLC, IPS, $P_m$ and Z, they are based on the standard normal distribution, while, for P, it is based on the $\chi^2_{2N}$ distribution.
2. "*" : significant at the 10% level. "**" : significant at the 5% level.

That there is a great deal of dependence between the $q_{it}$ can easily be seen from their correlation matrix. However, instead of presenting this matrix, following Luintel (2001)'s lead, I simply calculated the average of the correlations to be 0.68, which is a considerably high value.

The simplest way to deal with the dependence problem was to demean the data by subtracting $\overline{q}_t$ from each $q_{it}$. The average of the correlations between the demeaned series was now found to be 0.02, which indicates an appreciable reduction in dependence. Thus, I calculated the individual ADF tests, as well as the panel unit root tests (except those due to Hadri) using

$q_{it} - \overline{q}_t$ instead of $q_{it}$. The results are given in Table 3. The LLC, IPS and Z tests are still significant for the intercept-only case but at a lower level, while the P and $P_m$ tests are no longer significant. In the case of intercept + trend, none of the panel unit root tests are significant. As for the individual ADF tests, only the series for Denmark and Saudi Arabia are significant for the intercept-only case, and only the series for Norway in the constant + trend case; all at the 10% level. Only the Norwegian series has remained significant after demeaning.

**Table 4**

**MADF and SURADF Test Results**

| | **MADF** | ***Critical Values*** | | |
|---|---|---|---|---|
| | | *0.10* | *0.05* | *0.01* |
| Intercept | 80.029* | *76.179* | *81.215* | *91.555* |
| Intercept and Trend | 98.578 | *121.102* | *127.226* | *139.417* |

| | **Intercept** | | | | | **Intercept and Trend** | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | **p** | **SURADF** | *0.10* | *0.05* | *0.01* | **p** | **SURADF** | *0.10* | *0.05* | *0.01* |
| Austria | 2 | -5.987 | *-0.340* | *-6.742* | *-7.401* | 2 | -7.229 | *-8.336* | *-8.669* | *-9.243* |
| Belgium | 1 | -7.066** | *-6.661* | *-7.044* | *-7.657* | 1 | -8.275 | *-8.767* | *-9.066* | *-9.642* |
| Denmark | 1 | -6.335 | *-6.549* | *-6.930* | *-7.555* | 1 | -7.664 | *-8.604* | *-8.933* | *-9.560* |
| Finland | 1 | -3.727 | *-5.782* | *-6.188* | *-6.915* | 1 | -5.666 | *-7.419* | *-7.831* | *-8.559* |
| France | 1 | -6.811* | *-6.620* | *-6.976* | *-7.566* | 1 | -8.122 | *-8.671* | *-9.001* | *-9.610* |
| Germany | 1 | -6.631* | *-6.554* | *-6.907* | *-7.566* | 1 | -7.790 | *-8.588* | *-8.900* | *-9.484* |
| Greece | 1 | -2.582 | *-5.168* | *-5.597* | *-6.378* | 1 | -3.551 | *-6.508* | *-6.949* | *-7.713* |
| Italy | 1 | 4.352 | *-5.595* | *-6.013* | *-6.763* | 1 | -5.830 | *-7.144* | *-7.534* | *-8.282* |
| Japan | 1 | -3.736 | *-4.149* | *-4.575* | *-5.275* | 1 | -4.288 | *-5.137* | *-5.551* | *-6.250* |
| Netherlands | 1 | -6.738* | *-6.491* | *-6.856* | *-7.502* | 2 | -7.423 | *-8.443* | *-8.757* | *-9.353* |
| Norway | 1 | -4.654 | *-6.164* | *-6.548* | *-7.303* | 1 | -5.851 | *-7.966* | *-8.335* | *-9.069* |
| S. Arabia | 1 | -3.929 | *-4.448* | *-4.822* | *-5.477* | 1 | -3.566 | *-5.503* | *-5.856* | *-6.534* |
| Spain | 2 | -4.244 | *-5.906* | *-6.319* | *-7.019* | 2 | -5.745 | *-7.617* | *-7.994* | *-8.714* |
| Sweden | 1 | -2.757 | *-5.399* | *-5.822* | *-6.588* | 1 | -4.449 | *-6.873* | *-7.295* | *-8.036* |
| Switzerland | 1 | -5.656 | *-5.685* | *-6.088* | *-6.834* | 1 | -6.847 | *-7.298* | *-7.692* | *-8.368* |
| UK | 1 | -4.361 | *-5.043* | *-5.505* | *-6.242* | 1 | -5.417 | *-6.505* | *-6.742* | *-7.473* |
| USA | 1 | -3.456 | *-4.592* | *-4.956* | *-5.676* | 1 | -3.377 | *-5.731* | *-6.112* | *-6.838* |

Notes: The critical values were generated using Monte Carlo methods based on 10,000 replications, as was done by Breuer et al (2001). The authors are grateful to Myles Wallace for providing them with the necessary RATS code.

When the second solution, the MADF and SURADF tests, are applied to the data, Table 4 indicates that, in the intercept-only case, MADF is significant at the 10% level, while in the intercept + trend case it is not and neither are the individual SURADF tests. In the intercept-only case, on the other hand, the SURADF tests for the series due to Belgium, France, Germany and

Netherlands are significant. The SURADF test results appear to be consistent with the MADF results and the latter results are consistent with the LLC, IPS and P results given in Table 3.

**Table 5**

**The CADF and CIPS Test Results**

| | **Intercept** | | **Intercept and Trend** | |
|---|---|---|---|---|
| | **p** | **CADF** | **p** | **CADF** |
| Austria | 2 | -2.010 | 2 | -1.675 |
| Belgium | 1 | -2.545 | 1 | -2.233 |
| Denmark | 1 | -2.984* | 1 | -3.431* |
| Finland | 1 | -1.404 | 1 | -2.155 |
| France | 1 | -2.376 | 1 | -1.990 |
| Germany | 1 | -2.165 | 1 | -2.371 |
| Greece | 1 | -0.847 | 1 | -2.322 |
| Italy | 1 | -1.612 | 1 | -2.165 |
| Japan | 1 | -2.105 | 1 | -2.435 |
| Netherlands | 1 | -2.556 | 2 | -2.931 |
| Norway | 1 | -2.653 | 1 | -3.131 |
| S. Arabia | 1 | -2.946* | 1 | -1.979 |
| Spain | 2 | -2.066 | 2 | -1.854 |
| Sweden | 1 | -1.046 | 1 | -2.035 |
| Switzerland | 1 | -1.983 | 1 | -2.402 |
| UK | 1 | -1.985 | 1 | -2.585 |
| USA | 1 | -2.357 | 1 | -1.931 |
| **CIPS** | | -2.096 | | -2.331 |

Notes:

1. The critical values for the CADF and CIPS tests have been obtained from Pesaran (2005), Tables 1b, 1c, 3b and 3c.

Critical Values for CADF, p > 0, N = 20, T = 200 (Tables 1b and 1c)

| | 0.10 | 0.05 | 0.01 |
|---|---|---|---|
| Intercept | -2.91 | -3.23 | -3.84 |
| Intercept and Trend | -3.41 | -3.71 | -4.32 |

Critical Values for CIPS, p > 0, N = 20, T = 200 (Tables 3b and 3c)

| | 0.10 | 0.05 | 0.01 |
|---|---|---|---|
| Intercept | -2.11 | -2.20 | -2.36 |
| Intercept and Trend | -2.63 | -2.70 | -2.85 |

2. '*' : significant at the 10% level.

The results of applying the third solution, namely, using the CADF and CIPS tests give similar results as the previous two solutions. These may be seen from Table 5. The pooled test, CIPS, is insignificant for both cases while CADF is significant (at the 10% level) for the Danish and Saudi Arabian series in the intercept-only case (same as in the solution based on demeaned data) and for the Danish series in the intercept + trend case.

The final solution that was implemented to deal with dependence was to partition each series into common factors and idiosyncratic components. The common factors and the idiosyncratic components were separately tested for unit roots and the pooled tests were applied to the idiosyncratic components.

The first question that needed to be solved, however, was to choose the *n* common factors, $F_{tj}$. Bai and Ng (2002) had developed information criteria for this purpose but they yielded good results only when both *N* and *T* were large. Since *N*, in the present case, was rather small, I was not able to use these criteria. I, instead, used a simpler procedure and calculated the percentage of the total variance accounted for by the first n eigenvectors (i.e., the common factors). Since the sum of the eigenvalues is equal to the trace of the matrix $(T-1)^{-1}QQ'$ [see. e.g., Srivastava (2002: 404)], then this percentage may be obtained as $\sum_{i=1}^{n}\lambda_i / \sum_{i=1}^{T-1}\lambda_i$ where $\lambda_i$ denotes the eigenvalues. It was found that the percentage due to the first eigenvector, in both cases, was 86.7 and one gained only 7.3 or 7.4 percentage points when one considered the first three eigenvectors. Thus, I decided to choose *n* = 1; that is, I chose the first eigenvector as the common factor.[10]

The ADF test results for $\hat{F}_t$ and the idiosyncratic components are given in Table 6. Note that the null hypothesis of a unit root is rejected, at the 10% level, for the common factor in the intercept-only case but is not rejected for the intercept + trend case. In the intercept-only case, the null hypothesis is rejected only for the idiosyncratic component of the Netherlands series while it is rejected for the idiosyncratic component for the Japanese series in the case of intercept + trend. It is also noted, from columns (5) and (6) and columns (9) and (10) of Table 6, that the variation in the real exchange rates are dominated by the common factor. If all variations had been idiosyncratic, then the figures in columns (5) and (9) would have been close to unity and those in columns (6) and (10) would have been very small. But the reverse is found to hold in all cases.

| Table 6 | | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| **The ADF Test on the Common Factor and the Idiosyncratic Components** | | | | | | | | |
| | **Intercept** | | | | **Intercept and Trend** | | | |
| | **p** | **ADF** | $\frac{Var(\Delta\hat{e})}{Var(\Delta q)}$ | $\frac{\sigma(\varphi'\hat{F})}{\sigma(\hat{e})}$ | **p** | **ADF** | $\frac{Var(\Delta\hat{e})}{Var(\Delta q)}$ | $\frac{\sigma(\varphi'\hat{F})}{\sigma(\hat{e})}$ |
| $\hat{F}$ | 2 | -2.586<br>(0.098)* | | | 1 | -3.120 | | |
| Austria | 6 | -0.690<br>(0.417) | 0.0487 | 2.7276 | 4 | -0.855 | 0.0492 | 3.2939 |
| Belgium | 3 | -0.558<br>(0.474) | 0.0349 | 3.5934 | 3 | -1.063 | 0.0353 | 4.0873 |
| Denmark | 1 | -0.068<br>(0.659) | 0.0382 | 3.3436 | 2 | -0.983 | 0.0385 | 4.3638 |
| Finland | 12 | -1.587<br>(0.106) | 0.0896 | 1.3235 | 12 | -2.153 | 0.0903 | 1.8220 |
| France | 3 | -0.998<br>(0.285) | 0.0353 | 4.4832 | 3 | -1.026 | 0.0356 | 4.6489 |
| Germany | 1 | -1.326<br>(0.171) | 0.0428 | 3.1708 | 1 | -1.458 | 0.0432 | 3.3930 |
| Greece | 12 | -0.702<br>(0.412) | 0.1473 | 1.5698 | 12 | -1.121 | 0.1475 | 2.0198 |
| Italy | 3 | -1.604<br>(0.102) | 0.1024 | 2.0945 | 3 | -1.569 | 0.1029 | 2.3565 |
| Japan | 1 | -0.874<br>(0.336) | 0.3584 | 1.0044 | 8 | -2.905** | 0.3572 | 0.7126 |
| Netherlands | 1 | -1.955<br>(0.049)** | 0.0456 | 4.2159 | 1 | -2.034 | 0.0460 | 4.4383 |
| Norway | 5 | -0.760<br>(0.386) | 0.0584 | 3.4677 | 1 | -2.118 | 0.0586 | 4.8370 |
| S. Arabia | 1 | 0.244<br>(0.756) | 0.3762 | 0.5071 | 1 | -0.616 | 0.3754 | 0.7126 |
| Spain | 1 | -0.473<br>(0.510) | 0.0756 | 1.8935 | 1 | -0.836 | 0.0765 | 1.9313 |
| Sweden | 1 | -1.054<br>(0.263) | 0.1264 | 1.7591 | 1 | -1.252 | 0.1266 | 2.3136 |
| Switzerland | 3 | -1.584<br>(0.107) | 0.1064 | 2.2580 | 3 | -2.128 | 0.1065 | 2.6512 |
| UK | 1 | -1.425<br>(0.144) | 0.1668 | 1.4842 | 1 | -1.296 | 0.1671 | 1.7602 |
| USA | 1 | -0.809<br>(0.364) | 0.3486 | 0.8392 | 1 | -0.796 | 0.3480 | 0.8683 |

Notes:

1. The ADF statistic for $\hat{F}$ has the usual Dickey-Fuller distribution. Hence, the p-values in parentheses are based on MacKinnon (1996) and refer to autoregressions containing only and intercept and both an intercept and trend term.
2. The ADF statistics for the idiosyncratic components in the intercept-only case also have the usual Dickey-Fuller distribution. Hence, their p-values are also based on MacKinnon (1996) and refer to autoregressions without intercept and trend terms.
3. The critical values regarding the ADF test on the idiosyncratic components for the intercept + trend case are from Table 1 of Schmidt and Lee (1991) and correspond to T = 200.

| 0.10 | 0.05 | 0.01 |
|---|---|---|
| -2.34 | -2.63 | -3.19 |

4. "*" : significant at the 10% level "**" : significant at the 5% level

[10] The average of the correlations between the idiosyncratic components, $\hat{e}_{it}$, was found to be –0.002, indicating an even sharper reduction in correlation than what was obtained through demeaning.

| Table 7 | | | | |
|---|---|---|---|---|
| **KPSS and Hadri Test Results as Applied to the $\hat{e}_{it}^{0}$ and $\hat{e}_{it}^{1}$** | | | | |
| | **Intercept** | | **Intercept and Trend** | |
| | $\bar{k}$ | **KPSS** | $\bar{k}$ | **KPSS** |
| Austria | 11 | 0.858*** | 12 | 0.198*** |
| Belgium | 11 | 0.589** | 11 | 0.201*** |
| Denmark | 11 | 0.936*** | 12 | 0.167** |
| Finland | 11 | 1.174*** | 14 | 0.125** |
| France | 11 | 0.227 | 11 | 0.168** |
| Germany | 11 | 0.552** | 14 | 0.148** |
| Greece | 11 | 1.537*** | 18 | 0.140** |
| Italy | 11 | 0.629** | 23 | 0.157* |
| Japan | 11 | 0.669** | 14 | 0.063 |
| Netherlands | 11 | 0.404* | 12 | 0.100* |
| Norway | 11 | 1.349*** | 12 | 0.119* |
| S. Arabia | 11 | 1.027*** | 37 | 0.159** |
| Spain | 11 | 0.372* | 32 | 0.153** |
| Sweden | 11 | 0.935*** | 11 | 0.230*** |
| Switzerland | 11 | 0.916*** | 11 | 0.120* |
| UK | 11 | 0.602** | 14 | 0.175** |
| USA | 11 | 0.427* | 14 | 0.290*** |
| **Hadri 1** | 19.338 (0.000)*** | | | |
| **Hadri 2** | 16.867 (0.000)*** | | | |

Notes:

1. The critical values for the KPSS statistics in the intercept-only case are from Kwiatowski et al (1992), Table 1.

| 0.10 | 0.05 | 0.01 |
|---|---|---|
| 0.347 | 0.463 | 0.739 |

2. The critical values for the KPSS statistics in the intercept + trend case are from Table 1 of Shin (1994).

| 0.10 | 0.05 | 0.01 |
|---|---|---|
| 0.097 | 0.121 | 0.184 |

3. The p-values for the Hadri tests are based on the standard normal distribution.
4. "*" : significant at the 10% level, "**" : significant at the 5% level, "***" : significant at the 1% level

Finally, I turn to testing the null hypothesis of stationarity. In the intercept-only case, I found the KPSS statistic for $\hat{F}_t$ to be 0.335 and the critical value at the 10% level being 0.347, I do not reject the null hypothesis that $\hat{F}_t$ is stationary. This implies that the $Z_\mu$ test given in (13) above may directly be applied to the idiosyncratic components; in other words, the $\hat{e}_{it}^{0}$ are to be used. On the other hand, in the case of intercept + trend, since the KPSS statistic for $\hat{F}_t$ was

0.126 and that indicated that the stationarity null should be rejected at the 10% level (see the critical value in Table 1), I needed to obtain the $\hat{e}_{it}^{1}$ to test the stationarity in the idiosyncratic components. Also, I was able to apply Hadri's approach to the idiosyncratic components in the intercept-only case, but not to the $\hat{e}_{it}^{1}$ since they are not asymptotically independent. Thus, in Table 7, the KPSS test results are presented as applied to the $\hat{e}_{it}^{0}$ and $\hat{e}_{it}^{1}$ and the Hadri test results as applied to the $\hat{e}_{it}^{1}$. It is found that, in the intercept-only case, there is again (as in the ADF case) only one stationary series but this is now the French series. The pooled Hadri tests also indicate that the panel of series, as a whole, are not stationary. In the intercept + trend case however, the KPSS results agree exactly with the ADF results as applied to the $\hat{e}_{it}$; namely, only the Japanese series appear to be I(0), the rest are all I(1).

## 5. Conclusions

In this paper I investigated the persistence in Turkish real exchange rates using panel procedures. The reason for using panel models was the expected improvement in power over univariate tests due to the added increase in the variability of the data when the cross section dimension is taken into account. In other words, evidence in favour of the absolute version of the PPP hypothesis was expected to be obtained when such procedures were utilized.

I first implemented seven panel procedures, LLC, IPS, P, $P_m$, Z, Hadri 1 and 2, under the unrealistic assumption that the series making up the panel were independent of each other. I then took the dependence between the series into account by demeaning, by applying multivariate procedures based on SUR systems, by decomposing the disturbances in the autoregressions that yield the ADF statistic into their common factors and idiosyncratic components and, finally, by doing the same decomposition for the series themselves. I applied all these procedures to a panel of 17 Turkish bilateral real exchange rates that covered the period 1984.01-2001.06. The conclusions are as follows:

1. The application of the individual ADF and KPSS tests to these 17 series indicated that there was some weak support of the PPP hypothesis for the period in question when the intercept only case is considered. When a trend term is added, it is difficult to claim any support for PPP.

2. On the other hand, when first generation panel unit root tests were applied support for the PPP hypothesis was given by the all the tests with a unit root null while both Hadri tests rejected the stationarity of the series. This result was obtained irrespective of whether a trend term was included or not.
3. When the data was demeaned, LLC, IPS and Z still supported the PPP hypothesis in the intercept-only case, but at a lower level of significance while none of the panel unit root tests rejected the null when a trend term was added. The support for PPP from individual ADF tests were further reduced.
4. There was some weak support from the MADF test for the intercept-only case and only four significant outcomes for the SURADF tests, but there was no support for PPP from these tests when a trend term was added.
5. The results obtained from the CADF and CIPS tests were not any different from the demeaning and multivariate testing solutions for the cross-sectional dependence problem.
6. In decomposing the series into their common factors and idiosyncratic components, it was found that, in both cases, a single common factor was sufficient to account for the common component of the series. This common component was I(0) for the intercept-only case but I(1) for the intercept + trend case. The common component also dominated the variance of each $q_i$, implying that it was the factor contributing to the rejection of the null when the univariate and the majority of the panel tests were directly applied to the $q_{it}$ in the intercept only case and the non-rejection in the intercept + trend case. In fact, when the univariate ADF and KPSS tests were applied to the idiosyncratic components in the latter case, only one series was found to be I(0).
7. In sum, the support that was obtained for the absolute version of the PPP hypothesis from applying the first generation panel procedures directly to the $q_{it}$ appear to be due to ignoring the dependence between the series. The procedures where this dependence is accounted for either give very weak support to the PPP hypothesis (intercept-only case) or strongly favour the presence of a unit root in the series. A, rather informal, explanation for this outcome may be obtained by comparing the plots of the series for Germany, our largest trading partner, and the common component, $\hat{F}_t$. This is given in

**Figure 2**
**Plot of the Common Factor (F) and the DM-Based Real Exchange Rate**

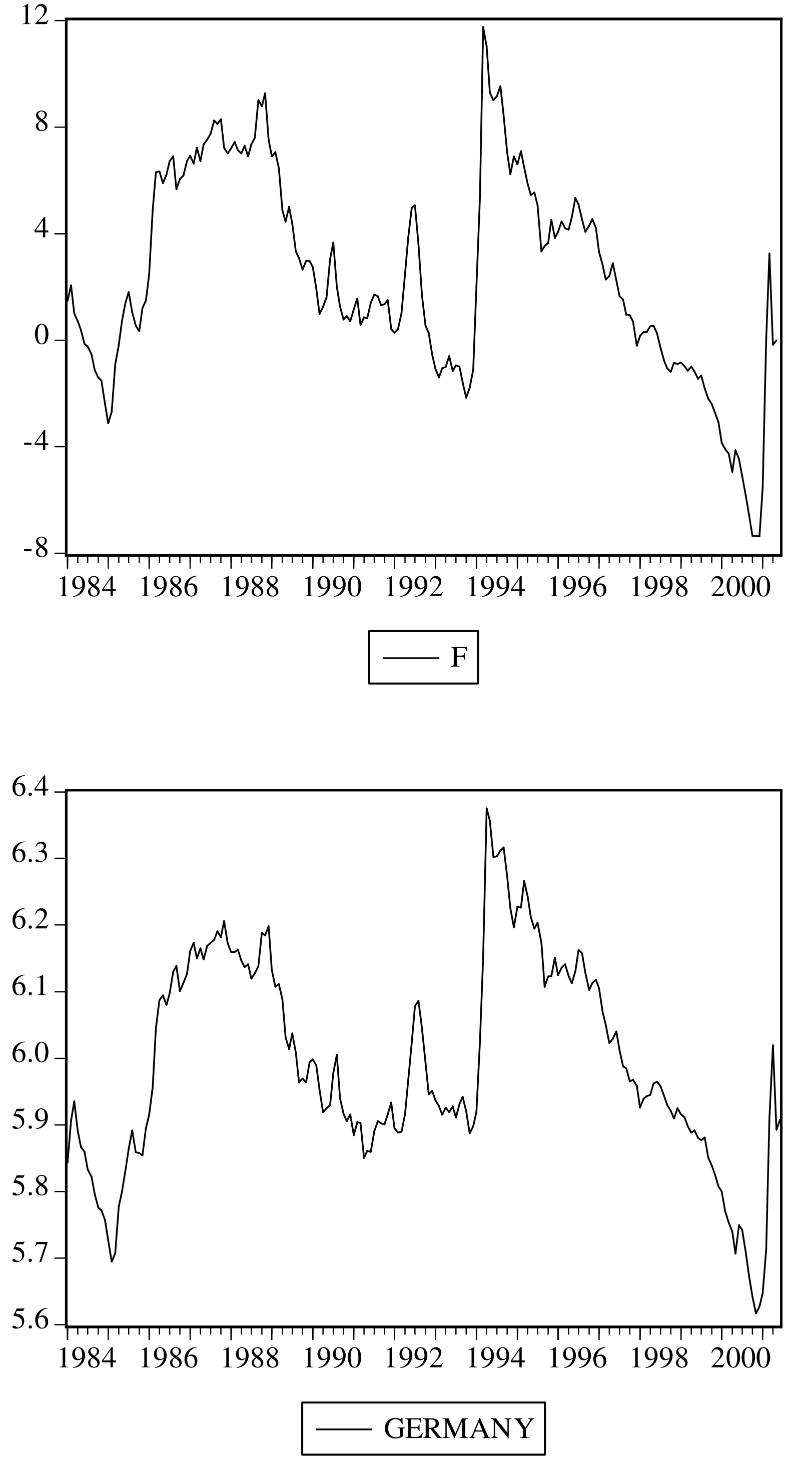


Figure 2. Note that the series are almost the same. Thus, it is not surprising to find that testing for a unit root in a panel of Turkish RERs when the majority of the series are from continental Europe and they resemble the German series does not provide any evidence supporting the PPP hypothesis. This strong co-movement in the series is, apparently, not sufficiently offset by cross-sectional heterogeneity, so that the null of a unit root is not rejected when the dependence between the series is taken into account, particularly when a trend terms in included.

8. What may be done, in future research, is to incorporate structural shifts in the deterministic terms with the testing procedures. But, due to the remarks in item (7), this may not give any new results other than the ones obtained, in a univariate framework, by Erlat (2003), which were favorable to the "quasi" PPP hypothesis.

**References**


Abuaf, N. and P. Jorion (1990): 'Purchasing Power Parity in the Long Run', *Journal of Finance*. 45(1), 157-174.

Bai, J. and S. Ng (2002): "Determining the Number of Factors in Approximate Factor Models', *Econometrica*, 70(1), 191-221.

Bai, J. and S. Ng (2004a): "A PANIC Attack on Unit Roots and Cointegration". *Economtrica*, 72(4), 1127-1177.

Bai, J. and S. Ng. (2004b): "A New Look at Panel Testing of Stationarity and the PPP Hypothesis'. D. Andrews and J. Stock (eds.): *Identification and Inference in Econometric Models: Essays in Honor of Thomas J. Rothenberg*. Cambridge: Cambridge University Press.

Baltagi, B.H. (2008): *Econometric Analysis of Panel Data*, 4th Ed. Chichester: John Wiley.

Banerjee, A., M. Marcellino and C. Osbat (2001): "Testing for PPP: Should We Use Panel Methods?", *Empirical Economics*, 30(1), 77-91.

Breitung, J. and S. Das (2005): "Panel Unit Root Tests Under Cross-Sectional Dependence", *Statistica Neerlandica*, 59(4), 414-433.

Breuer, J.B., R. McNown and M.S. Wallace (2001): "Misleading Inferences from Panel Unit Root Tests with an Illustration from Purchasing Power Parity", *Review of International Economics*, 9(3), 482-493.

Breuer, J.B., R. McNown and M.S. Wallace (2002): "Series-Specific Unit Root Tests with Panel Data", *Oxford Bulletin of Economics and Statistics*, 64(5), 527-546.

Caner, M. and L. Kilian (2001): "Size Distortions of Tests of the Null Hypothesis of Stationarity: Evidence and Implications for the PPP Debate", *Journal of International Money and Finance*, 20(2), 639-657.

Choi, I. (2001): "Unit Root Tests for Panel Data", *Journal of International Money and Finance,* 20, 249-272.

Choi, I. (2002): "Combination Unit Root Tests for Cross-Sectionally Correlated Panels", Working Paper, Hong Kong University of Science and Technology.

Erlat, H. (2002): "Long Memory in Turkish Inflation Rates", in A. Kibritçioğlu, L. Rittenberg and F. Selçuk (eds.): *Inflation and Disinflation in Turkey*. Brookfield, Vermont: Ashgate, 97-122.

Erlat, H. (2003): "The Nature of Persistence in Turkish Real Exchange Rates", *Emerging Markets Finance and Trade*, 39(2), 70-97.

Erlat, H. (2004): "Unit Roots or Nonlinear Stationarity in Turkish Real Exchange Rates", *Applied Economics Letters*, 11, 645-650.

Flores, R., P. Jorion, P.Y. Premont and A. Szafarz (1999): "Multivariate Unit Root Tests of the PPP Hypothesis", *Journal of Empirical Finance*, 6, 335-353.

Groen, J.J. (2000): "New Multi-Country Evidence on Purchasing Power Parity: Multivariate Unit Root Test Results", Econometric Institute Report EI 2000-9/A, Erasmus University Rotterdam.

Hadri, K. (2000): "Testing for Stationarity in Heterogeneous Panels", *Econometrics Journal*, 3, 148-161.

Hegwood, N. and D. Papell (1998): "Quasi Purchasing Power Parity", *International Journal of Finance and Economics*, 3, 279-289.

Higgins, M. and E. Zakrajsek (2000): "Purchasing Power Parity: Three Stakes Through the Heart of The Unit Root Null", Finance and Economics Discussion Series 2000-22, Federal Reserve Board.

Ho, T.W. (2002): "Searching Stationarity in the Real Exchange Rates: Application of the SUR Estimator", *Open Economies Review*, 13, 275-289.

Hurlin, C. and V. Mignon (2004): “Second Generation Panel Unit Root Tests”, Working Paper, LEO, University of Orleans.

Im, K.S., M.H. Pesaran and Y. Shin (2003): “Testing for Unit Roots in Heterogeneous Panels”, *Journal of Econometrics*, 115(1), 53-74.

Jonsson, K. (2005): “Cross-Sectional Dependency and Size Distortion in a Small-Sample Homogeneous Panel Data Unit Root Test”, *Oxford Bulletin of Economics and Statistics*, 67(3), 369-392.

Jorion, P. and R.J. Sweeney (1996): “Mean Reversion in Real Exchange Rates: Evidence and Implications for Forecasting”, *Journal of International Money and Finance*, 15(4), 535-550.

Kuo, B.S. and A. Mikola (2001): “How Sure Are We About Purchasing Power Parity? Panel Evidence with the Null of Stationarity of Real Exchange Rates”, *Journal of Money, Credit and Banking*, 33(3), 767-789.

Kwiatowski, D., P.C.B. Phillips, P. Schmidt and Y. Shin (1992): “Testing the Null Hypothesis of Stationarity Against the Alternative of a Unit Root: How Sure Are we that Economic Time Series Have a Unit Root?”, *Journal of Econometrics*, 54, 159-178.

Levin, A., C.F. Lin and C.S.J. Chu (2002): “Unit Root Tests in Panel Data: Asymptotic and Finite Sample Properties”, *Journal of Econometrics*, 108, 1-24.

Luintel, K.B. (2001): “Heterogeneous Panel Unit Root Tests and Purchasing Power Parity”, *Manchester School Supplement*, 42-56.

MacKinnon, J.G. (1996): “Numerical Distribution Functions for Unit Root and Cointegration Tests”, *Journal of Applied Econometrics*, 11, 601-618.

Maddala, G.S. and S. Wu (1999): “A Comparative Study of Unit Root Tests with Panel Data and a New Simple Test”, *Oxford Bulletin of Economics and Statistics*, 61, 631-652.

Moon, H.R. and B. Perron (2004): “Testing for a Unit Root in Panels with Dynamic Factors”, *Journal of Econometrics*, 122, 81-126.

Newey, W. and K. West (1994): “Automatic Lag Selection in Covariance Matrix Estimation”, *Review of Economic Studies*, 61, 631-653.

O’Connell, P. (1998): “The Overvaluation of Purchasing Power Parity”, *Journal of International Economics*, 44, 1-19.

Özdemir, N. (2002): *Panel Approaches to Testing the Persistence in Turkish Real Exchange Rates*. Unpublished thesis submitted to the Department of Economics, Middle East Technical University, in partial fulfillment of the degree of Master of Science.

Pesaran, M.H. (2006): "Estimation and Inference in Large Heterogeneous Panels with a Multifactor Error Structure", *Econometrica*, 74(4), 967-1012..

Pesaran, M.H. (2007): "A Simple Panel Unit Root Test in the Presence of Cross Section Dependence", *Journal of Applied Econometrics*, 22(2), 265-312.

Phillips, P.C.B. and D. Sul (2003): "Dynamic Panel Estimation and Homogeneity Testing Under Cross Section Dependence", *Econometrics Journal*, 6(1), 217-259.

Sarno, L. (2000a): "Real Exchange Rate Behaviour in the Middle East: A Re-examination", *Economics Letters*, 66, 127-136.

Sarno, L. (2000b): "Real Exchange Rate Behaviour In High Inflation Countries: Empirical Evidence from Turkey", *Applied Economics Letters*, 7, 285-291.

Sarno, L. and M.P. Taylor (2002): "Purchasing Power Parity and the Real Exchange Rate", *IMF Staff Papers*, 49(1), 65-105.

Schmidt, P. and J. Lee (1991): "A Modification of the Schmidt-Phillips Test", *Economics Letters*, 36, 285-289.

Shin, Y. (1994): "A Residual Based Test for the Null of Cointegration Against the Alternative of No Cointegration", *Econometric Theory*, 10(1), 91-115.

Srivastava, M.S. (2002): *Methods of Multivariate Analysis*. New York: John Wiley and Sons.

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

Zellner, A. (1962): "An Efficient Method of Estimating Seemingly Unrelated Regressions and Tests for Aggregation Bias", *Journal of the American Statistical Association*, 57(298), 348-368.