Re-examination of the long-run purchasing power parity: further evidence from Turkey

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Re-examination of the long-run purchasing power parity: further evidence from Turkey

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In this article, we re-examine the empirical validity of the Purchasing Power Parity (PPP) theory for the Turkish economy. For this purpose, an empirical model is constructed using some contemporaneous estimation techniques such as multivariate co-integration and vector error correction methodology. Our estimation results reveal that the PPP can strongly be supported as a long-run stationary steady-state relationship for the Turkish economy.

I. Introduction

The long-run determination of exchange rates has been of a special issue of interest for the researchers when constructing theories and policies in new open economy macroeconomics. Such researches conducted to explore the motives behind the course of policies would reveal the extent to which applying to discretionary policy tools can be succeeded in attaining *ex ante* policy targets. Of all these theoretical debates, Purchasing Power Parity (PPP) attracts a considerable attention especially for the post-1973 period following the collapse of the Bretton Woods system. In this article, our contribution to the economics literature is to assess some theoretical issues dealing with the PPP theory once again and to re-examine the validity of the PPP employing data from the Turkish economy. For this purpose, the rest of this article is organized as follows. The next section outlines some theoretical concepts for the PPP theory. Section III interests in alternative methods for testing purposes, and Section IV gives

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the empirical methodology applied in this article. Section V provides results, whilst the Section VI concludes.

II. Theoretical Concepts

The PPP relationship is based on the law of one price which states that under the frictionless goods arbitrage, the prices of individual traded goods should have been equalized when the prices are expressed in terms of the same currency of denomination (Sarno and Taylor, 2002a):

$$p_{i,t}^{d} = e_t + p_{i,t}^{f} \tag{1}$$

where $p_{i,t}^{d}$ is the domestic currency price of any good *i* at any time *t*, e_t the domestic currency price of foreign exchange at time *t*, and $p_{i,t}^{f}$ the relevant foreign currency price of good *i* at time *t*, all expressed in natural logarithms. Letting Equation 1 hold for every individual good leads to the assumption that it must hold for any identical basket of goods. Even if the law of one price fails for individual goods, it is possible that the deviations cancel out when averaged across a basket of goods (Froot and Rogoff, 1994).

The absolute PPP theory embedded in Equation 1 has been subject to some criticism in which it gives the exchange rate that equilibrates trade in goods and services while disregarding the capital account (Salvatore, 1998). Furthermore, due to the existence of nontradeable goods in the consumption bundles included in prices indices, whose prices in part depend on relative productivity levels discussed mainly by Balassa (1964) and Samuelson (1964), the PPP relationship would not completely give the exchange rate that equilibrates trade in goods and services (Jenkins and Snaith, 2005). Another shortcoming for the empirical purposes is to assume a causal relationship running from relative prices to the nominal exchange rate when the causation runs from a different way (Taylor, 2006). This is especially problematic when the exogeneity/endogeneity characteristic of the price levels and the nominal exchange rate have not been elaborately considered by the researchers to obtain a mutually stationary causal relationship between the variables. Furthermore, for the lack of PPP relationship, an important contribution comes from the pricing-to-market theorem of Krugman (1987) and Dornbusch

(1987) as well, which argues the price stickiness phenomenon in international trade following the imperfect competition conditions subject to the economic agents. The latter models assume a monopolisitic or an oligopolistic market structure which leads the producers to charge different prices in different country cases, whilst the exchange rate changes would not cause fluctuations in relative prices charged (Obstfeld and Rogoff, 2000). This is possible because there are many industries that can supply separate licences for the sale of their goods at home and abroad (Sarno and Taylor, 2002b). As a methodological problem to be considered, employing low-frequency data and model misspecification can also lead to biased results in favour of the slow convergence of real exchange rates to the long-run equilibrium (Taylor, 2001).

Considering all these, the relative form of the PPP relationship can be constructed by relating the change in the exchange rate over a period of time to the relative changes in the price levels in different countries over the same time period within a proportional relation.[1] In this sense, Taylor and Taylor (2004) express that the relative PPP would hold if the absolute PPP holds, but the absolute PPP does not necessarily hold when the relative PPP holds, since it is possible that common changes in nominal exchange rates may be happening at different levels of purchasing power for the two currencies.

III. Alternative Methods for Testing PPP

Taylor (1996) reveals that exchange rates and prices might be determined simultaneously which give them an endogenous characterics against each other. If so, single equation results may lead researchers to misleading conclusions resulted from simultaneity bias and invalid conditioning problems (Gokcan and Ozmen, 2001). Testing the PPP hypothesis requires that real exchange rate should be a long-run stationary/mean reverting process that is equal to a constant or that a linear combination of same order integrated variables in Equation 1 should be satisfied significantly with *a priori* assumed signs.[2]

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[1] Likewise, Rogoff (1996) touchs upon briefly the possible data problems taken place by the differences in the basket weights considered to construct the price indices on which the PPP relationship is based.

[2] Rather than imposing a linear specification of the PPP hypothesis, some contemporaneous recent literature dealing with mean reversion of real exchange rates also emphasize that the time-series behaviour of the real exchange rates can be beter approximated by a nonlinear adjustment process employing smooth transition autoregressive models. For a brief account of such a methodology, see, e.g. Kilian and Taylor (2003) and Taylor (2006).

Alternatively, we can also consider the half-life of the real exchange rates in search for a support for the PPP hypothesis, rather than conventional analyses of whether real exchange rate shocks are mean reverting or not, where the half-life is defined as the duration of time required for half the magnitude of a unit shock to the level of a series to dissipate. Too long a half-life of real exchange rates, in general accepted longer than 5 years, would indicate the purchasing power puzzle of the low-convergence speed in real exchange rates (Rogoff, 1996; Chen and Engel, 2004).

Of all these alternatives to measure the validity of the PPP hypothesis, in this article we try to examine whether a variable vector ($p_t^d \quad e_t \quad p_t^f$)′ can be represented by a stationary steady-state process which satisfies the coefficient restriction vector (1 -1 -1)′ with a significant feedback process that reveals the endogeneity of the variables.[3] If homogeneity and symmetry restrictions cannot be rejected, the strong form of the PPP can be supported. Otherwise, weak form of such a relationship would be brought out provided that symmetry restrictions can be satisfied.

IV. Estimation Methodology

We now test for a long-run stationary relationship within the *ex ante* endogenous variable vector related to the absolute PPP hypothesis, and for this purpose the multivariate co-integration and vector error correction techniques proposed by Johansen (1988) and Johansen and Juselius (1990) are used. Let us assume a z_t vector of nonstationary *n* endogenous variables and model this vector as an unrestricted vector autoregression (VAR) involving up to *k*-lags of z_t:

$$z_t = \Pi_1 z_{t-1} + \ldots + \Pi_k z_{t-k} + \varepsilon_t \tag{2}$$

[3] Taylor (1988) and Kim (1990) testing the long-run PPP relationship for some major currencies against the US dollar (1990) can be considered among the pioneering studies that use co-integrating and VRC techniques to reveal both the long-run stationary relationships leading to the PPP hypothesis and the deviations from the PPP relationship, that affect exchange rates. Furthermore, Sarno and Taylor (1998) and Taylor and Sarno (1998) in this sense are deserved to be examined in the contemporaneous economics literature.

where ε_t follows an i.i.d. process $N(0,\sigma^2)$ and z_t is (nx1) and the Π_i an (nxn) matrix of parameters. Equation 2 can be rewritten in a vector error correction (VEC) model of the form:

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + \ldots + \Gamma_{k-1} \Delta z_{t-k+1} + \Pi z_{t-k} + \varepsilon_t \tag{3}$$

where:

$$\Gamma_i = -I + \Pi_1 + \ldots \Pi_i \ (i = 1,2, \ldots, k\text{-}1) \tag{4}$$

and

$$\Pi = \mathrm{I} - \Pi_1 - \ldots - \Pi_k \tag{5}$$

Equation 3 can be arrived by substracting z_{t-1} from both sides of Equation 2 and collecting terms on z_{t-1} and then adding $-(\Pi_1\text{-}1)X_{t-1} + (\Pi_1\text{-}1)X_{t-1}$. Repeating this process and collecting of terms will yield Equation 3. This specification of the system of variables carries on the knowledge of both the short- and long-run adjustment to changes in z_t, via the estimates of Γ_i and Π. Following Harris (1995), $\Pi = \alpha\beta'$ where α measures the speed of adjustment coefficient of particular variables to a disturbance in the long-run equilibrium relationship and can be interpreted as a matrix of error correction terms, while β is a matrix of long-run coefficients such that $\beta' z_{t-k}$ embedded in Equation 3 represents up to (n-1) co-integrating relations in the multivariate model which ensures that z_t converge to their long-run steady-state solutions. Note that all terms in Equation 3 which involve Δz_{t-i} are I(0) while Πz_{t-k} must also be stationary for ε_t~I(0) to be white noise of an $N(0, \sigma_\varepsilon^2)$ process.

Dealing with the rank conditions, three alternative cases can be considered. If the rank of Π matrix equals zero, there would be no co-integrating relation between the endogenous variables, which means that there would be no linear combinations of the z_t that are I(0) leading to Π would be an (nxn) matrix of zeros. In this case, a VAR model consisted of a set of variables in first differences, thus carrying no long-run knowledge of any stationary relationship could be suggested to examine the variable system. If the Π matrix is of full rank when $r=n$, then all elements in z_t would be stationary in their levels. Of special interest here is the possibility that there exists r co-integrating vectors in $\beta' z_t \sim I(0)$ and (n-r) common stochastic trends when β has reduced rank, i.e. $0<r\leq(n-1)$. That is, first r columns of β are the linearly independent combinations of the endogenous variables settled in the vector z_t, which represents stationary relationships. Whereas, the latter (n-r) columns constitute the nonstationary vectors of I(1) common trends, which also require that the last (n-r) columns of α take insignificantly values highly close to zero, impeding feedback effects of deviations from long-run stationary equilibrium process. Thus, this method is equivalent to testing which columns of α are zero (Harris, 1995). Gonzalo (1994) indicates that this method performs better than other estimation methods even when the errors are non-normal distributed or when the dynamics are unknown and the model is over-parameterized by including additional lags in the error correction model.

We estimate the existence of potential co-integrating vector between the variables of interest by using two likelihood test statistics known as maximum eigenvalue for the null hypothesis of r versus the alternative of r+1 co-integrating relations and trace for the null hypothesis of r co-integrating relations against the alternative of n co-integrating relations, for $r = 0, 1, \ldots,$ n-1 where n is the number of endogenous variables. Following Johansen (1992), for the co-integration test, we restrict intercept and trend factor into our long-run variable space following the so-called Pantula principle. This requires a test procedure which moves through from the most restrictive model and at each stage compares the trace or max-eigen test statistics to its critical value, and only stop for the first time where the null hypothesis is not rejected.

V. Results

We now construct an empirical model for the Turkish economy using quarterly data for the period of 1987Q1 to 2006Q4. All the data used are taken from the electronic data delivery system of the Central Bank of the Republic of Turkey for the domestic variables and from the FRB of St. Louis electronic data delivery system for the foreign variables. Following Johansen and Juselius (1992) and Johansen (1995), we use seasonally unadjusted values in natural logarithms.

For the domestic ($p_{i,t}^{d}$) and foreign price level ($p_{i,t}^{f}$) data, the gross domestic product deflators from the Turkish and the US economy are used, whilst the spot exchange rate of YTL/US$ (e_t) is considered for the nominal exchange rate variable. Besides, we add a set of centred seasonal dummies which sum to zero over a year as exogenous variable so that the linear term from the dummies disappears and is taken over completely by the constant term, and only the seasonally varying means remain (Johansen, 1995).

We then examine the (non)stationarity characteristics of the variables. Granger and Newbold (1974) indicate that using nonstationary time series steadily diverging from long-run mean will produce biased SEs, which causes to unreliable correlations within the regression analysis leading to unbounded variance process. In this way, the standard OLS regression will produce a good fit and predict statistically significant relationships between the variables considered, however, none really exists (Mahadeva and Robinson, 2004). Dickey and Fuller (1979) provide one of the commonly used test methods known as Augmented Dickey–Fuller (ADF) test of detecting whether the time series are of stationary form. This can be formulated as follows:

$$\Delta y_t = \alpha + \beta t + (\rho - 1) y_{t-1} + \sum_{i=1}^{k} \eta_i \Delta y_{t-i} + \varepsilon_t \qquad (6)$$

where y_t is the variable of interest and t is a time trend. The k-lagged differences are to ensure a white noise error series and the number of lags is determined by a test of significance on the coefficient n_i. The null hypothesis of the ADF test is the presence of a unit root (ρ =1) against the alternative stationary hypothesis. For any y_t to be stationary, (ρ - 1) should be negative and significantly different from zero. We compare the estimated ADF statistics with the simulated MacKinnon (1996) critical values. For the case of stationarity, we accept that these statistics must be larger than the critical values in absolute value and have a minus sign.

However, due to the evidence yielded by, e.g. DeJong et al. (1989) Dickey–Fuller-type tests may have low power against plausible stationary alternatives, and therefore the ADF tests are supplemented by the tests proposed by Kwiatkowski et al. (1992) known as KPSS tests. The KPSS tests are designed to test the null hypothesis of stationarity against the unit root alternative. The results are given in Table 1.

Table 1. Unit Root Tests

Variable	τ_t	$\Delta\tau_t$	$Z(\tau_t)$	$\Delta Z(\tau_t)$
p_t^d	1.87	-10.15	0.31	0.09
e_t	0.71	-6.31	0.24	0.05
p_t^f	-2.42	-4.82	0.23	0.13

Above, τ_t are the test statistics with allowance for constant and trend terms for the ADF test. 'Δ' denotes the first difference operator. $Z(\tau_t)$ are the relevant KPSS statistics. Unit root test results indicate that the nonstationarity cannot be rejected for all variables in the level form. However, for the first differences of all the variables unit root hypothesis is strongly rejected. Thus all the series are integrated of order 1 which have an invertible ARMA representation after applying to first differencing. In Table 2, the co-integration test results given the nonstationary time-series characteristics of the variables are presented.

From Table 2, both LR tests verify the existence of one potential co-integrating factor with the largest eigenvalue lying in the long-run variable space. Rewriting the normalized equation upon the domestic price level under the assumption of $r = 1$ yield the following equation:

$$\beta' z_t = p_t^d - 1.01 e_t - 9.22 p_t^f + 0.06 trend \sim I(0) \tag{7}$$

In Equation 7, all the signs match *a priori* expectations and have statistical significance. A one-to-one relationship between domestic price level and nominal exchange rate can easily be noticed, and null of homogeneity for the nominal exchange rate cannot be rejected by the LR tests. However we find a highly large coefficient of foreign price level in magnitude, such a restriction in addition to the homogeneity restriction on the nominal exchange rate is accepted by the LR tests as well. Adjustment coefficients indicating feedback effects of disturbances from the steady-state functional form are found highly different from zero in a statistically significant way, which verify the endogenous characteristics of the variables in the long-run variable space. Finally, estimation results fit well to the data generating process as to the diagnostics.

Table 2. Co-integration test

Null hypothesis	$r = 0$	$r \leq 1$	$r \leq 2$	
Eigenvalue	0.51	0.17	0.12	
λ_{trace}	78.81*	24.18	9.58	
5% cv	42.92	25.87	12.52	
λ_{max}	28.81*	15.74	6.26	
5% cv	25.82	19.39	12.52	
Unrestricted co-integrating coefficients				
p_t^d	e_t	p_t^f	*trend*	
4.831790	-4.868700	-44.53728	0.188034	
1.259811	0.562840	90.97460	-0.708451	
-15.79781	14.31109	101.8962	-0.300238	
1 Co-integrating equation[a]				
p_t^d	e_t	p_t^f	*trend*	
1.000000	-1.007639	-9.217553	0.062925	
	(0.10129)	(2.67264)	(0.02436)	
Adjustment coefficients (the letter '*D*' denotes the difference operator)				
$D(p_t^d)$	$D(e_t)$	$D(p_t^f)$		
-0.208690	-0.210883	-0.001815		
(0.02695)	(0.04347)	(0.00085)		
Multivariate statistics for testing stationarity				
	p_t^d	e_t	p_t^f	
$\chi^2(2)$	6.465312	6.008656	7.060492	
Probs.	0.039453	0.049572	0.029298	

Homogeneity and symmetry restrictions
$b(1,2) = -1$, $\chi^2(1) = 0.01$ (prob. 0.95); b(1,2) = -1, b(1,3) = -1, $\chi^2(2) = 4.98$ (prob. = 0.14)

VEC Res. serial correlation LM test
H_0: no residual correlation at lag order h
LM(1) = 3.43 (prob. 0.95) LM(4) = 8.81 (prob. 0.48)

VEC res. normality test
H_0: residuals are multivariate normal
Skewness $\chi^2(3) = 1.37$ (prob. 0.71) Kurtosis $\chi^2(3) = 4.98$ (prob. 4.98)
Jarque-Bera (6) = 6.35 (prob. 0.39)

Notes: * Rejection of the hypothesis at the 0.05 level
[a] SE in parenthesis.

VI. Concluding Remarks

In this article, we try to construct an empirical model to re-examine the empirical validity of the PPP theory for the Turkish economy. Using some contemporaneous estimation techniques such as multivariate co-integration and VEC methodology, our estimation results give strong support to the validity of the PPP theory for the Turkish economy. Future papers should be elaborately constructed to reveal whether the estimation results in this article can be confirmed and whether they are in fact of the stylized facts for the Turkish economy. Complementary papers should also consider nonlinearities in the real exchange rates to test the validity of the PPP theory.

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