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DO PPP AND UIP NEED EACH OTHER IN A FINANCIALLY OPEN ECONOMY? THE TURKISH EVIDENCE

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Abstract

This paper investigates the empirical validity of the capital enhanced equilibrium exchange rates (CHEERs) model for the Turkish quarterly data from 1986:1 to 1999:4. The results of the Johansen cointegration analyses for the variable system containing Turkish and US inflation rates, interest rates, and exchange rate suggest the existence of two stationary relationships explaining the long run evolution of Turkish interest rates and inflation rates, respectively. The results of the structural model obtained by data-acceptable over-identifying restrictions over the cointegration space suggest the non-rejection of the hypothesis that the first vector contains uncovered interest parity (UIP) and the second vector contains purchasing power parity (PPP) with proportionality and symmetry conditions. Consistent with the CHEERs approach, each of the international parity hypotheses is strongly rejected when formulated independently. This is a theory-consistent result for a financially open economy for which equilibrium conditions of asset and commodity markets may not be independent of each other.

Keywords: PPP, UIP, Exchange rates, cointegration, Turkey

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I. INTRODUCTION

The purchasing power parity (PPP) and uncovered interest parity (UIP) hypotheses are among the most controversial issues in the international macroeconomics literature. The PPP hypothesis postulates that exchange rates adjust to price differentials in open economies to restore international commodity market equilibrium. The UIP, on the other hand, considers international asset markets, and asserts that exchange rates adjust to interest rate differentials. As a disequilibrium in one market may have repercussions on the other, the two international parity conditions may not be independent of each other in the long run.

The literature on both PPP and UIP is large and several comprehensive reviews of them are available (see, e.g. McCallum 1994, MacDonald 1999, Obstfeld and Rogoff 2000 and Taylor 2001). Empirical tests of each of these hypotheses have often yielded conflicting conclusions. There are ample explanations for this. The explanations for the failure of the PPP include imperfect competition, pricing to market, the composition of price indices, information costs, transport costs and trade barriers. Most explanations for the failure of UIP contain the existence of time varying risk premium, expectational errors and limited international capital mobility. Invalid conditioning, non-linear dynamics, the low power of the conventional unit root tests and temporal aggregation are among the empirical modelling issues suggested as the reasons for the mixed results especially for the validity of the PPP hypothesis.

In the literature, equilibrium exchange rates are often defined as either in terms of the PPP or UIP, but seldom both. For a financially open economy, the PPP may be postulated to depict the interdependence of commodity markets in terms of consumption prices, and UIP is a forward-looking equilibrium mechanism for the international financial markets. As asset and commodity markets may not be independent of each other, defining equilibrium exchange rates in terms of only one international parity condition may be seriously misleading[1].

Johansen and Juselius (1992), Juselius (1995) and Juselius and MacDonald (2000) propose an approach combining both international parities. This approach allowing for interactions among prices, interest rates and exchange rates is referred to as capital enhanced equilibrium exchange rates, or CHEERs by MacDonald (2000). As argued by MacDonald (2000, p.18), "this approach captures the basic Casselian view of PPP, …, that an exchange rate may be away from its PPP determined rate because of non-zero interest differentials".

This study aims to investigate the long run relationships between the variables in a system containing Turkish and US inflation rates, interest rates, and exchange rate employing Johansen cointegration procedure. This system allows also to test the empirical validity of the CHEERs model by combining the UIP and PPP hypotheses. The Turkish economy, during the sample covered by this study (1986-1999) can be interpreted as financially open with flexible exchange rates, liberalized international capital flows and domestic banking system offering deposits also in terms foreign currencies. This may make Turkey a natural candidate for investigating the validity of the PPP and the UIP hypotheses[2].

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[1] Defining equilibrium exchange rates appropriately may be crucially important also for the design of an exchange rate based stabilization program. If the evolution of exchange rates is not independent of interest rate differentials, then a PPP based exchange rate targeting policy may not be sustainable. This is basically because, the adjustment of exchange rates to capital flows due to risk adjusted interest parities may lead to targeted exchange rates substantially diverging from the equilibrium rates for a financially open economy.

[2] The recent studies investigating the validity of the PPP or/and UIP for the Turkish data includes Metin (1994), Mustafaoğlu (1999), Sarno (2000) and Erlat (2001). Metin (1994) tests the UIP and PPP jointly and finds that neither of the parities is supported by the Turkish annual data pertaining fixed and flexible exchange rate periods. Mustafaoglu (1999) finds that the real exchange rates are not stationary even if an endogenous break is taken account. The results by Erlat (2001) suggest that the absolute PPP hypothesis cannot be rejected when the recent fractional integration and single/multiple break point estimation methods are employed. Sarno (2000) employs non-linear modelling approach and provides a support for the validity of PPP.

The plan of the rest of this study is as follows. In Section 2, we present a brief overview of the theoretical and empirical relationships. Section 3 presents and evaluates the empirical results. Section 4 concludes.

II. THEORETICAL AND EMPIRICAL RELATIONSHIPS

The PPP hypothesis stems from the "law of one price" which states that, measured in a common currency, freely traded identical commodities should have the same price everywhere in the absence of transaction and transportation costs. That is:

$$p_t = e_t + p_t^* \qquad (1)$$

where e is the log of the nominal exchange rate (domestic currency per unit of foreign currency), p and p* are the logs of the domestic and foreign price levels, respectively. Rearranging (1) gives the PPP hypothesis (the strong or absolute form), postulating that the exchange rate equals to the price differential between two open economies:

$$e_t = p_t - p_t^* \qquad (2)$$

Equation (2) can be obtained from

$$e_t = \gamma_0 + \gamma_1 p_t - \gamma_2 p_t^* \qquad (3)$$

under a maintained hypothesis that $\gamma_0 = 0$ and $\gamma_1 = \gamma_2 = 1$ (symmetry and proportionality condition).

The lack of absolute price level data constructed for an internationally standardized basket of goods to test the absolute PPP for almost any country often enforces researchers to retreat to the testing of relative PPP (Rogoff, 1996). The relative (or weak) version of the PPP relaxes the restriction that $\gamma_0 = 0$, and often defines the evolution of exchange rates in a growth rate form:

$$\Delta e_t = \Delta p_t - \Delta p_t^* \qquad (4)$$

where Δ is the first difference operator.

The cointegration of variables in the systems defining the parities with theory-consistent unitary coefficients (or the stationarity of real exchange rates $q_t = e_t - p_t + p_t^*$) can be interpreted as evidence supporting the PPP hypothesis. A cointegration between a set of variables can be quite consistent with alternative conditioning restrictions for the parameters of the long-run relationship. Although going from (1) to (2) maintains the endogeneity of exchange rate, the theory may not be inconsistent with the joint endogeneity of all the variables in the system (except the case that one of the countries is small enough to effect price level of the other). As Taylor (1996, p. 8) notes, "There is no a priori reason to have exchange rates on the left-hand side and prices on the right". Similarly, Isard (1995, p. 59) defines the PPP "as a theory about the relationship between endogenous variables". Thus single equation results may be seriously misleading due to a simultaneity bias and/or invalid conditioning.

Besides the simultaneity issue, there are ample alternative explanations for the conflicting results for the validity of the PPP hypothesis. The explanations for the failure of the PPP include[3], imperfect competition, pricing to market, the choice and the composition of price indices, information costs, transport costs and trade barriers. Non-linear dynamics (Taylor and Sarno 1998), the low power of the conventional unit root tests over short time spans of data (Lothian and Taylor (1996) and Edison and Melick (1999)) and temporal aggregation (Taylor 2001) are among the empirical modeling issues suggested as the reasons for the mixed results for the validity of the PPP.

The UIP hypothesis, considering international capital flows, states that the rates of return on domestic and foreign assets expressed in the same currency are equal:

$$i_t = i_t^* + \Delta e_t^e \tag{5}$$

where i_t and i_t^* are domestic and foreign nominal interest rates with maturity t+m, respectively, $\Delta e_t^e = E_t \Delta e_{t+m}$ is the expected rate of exchange rate change during t+m, and E_t is the conditional expectations operator on the basis of information available at time t. Assuming that expectations are formed rationally:

$$\Delta e_t^e = \Delta e_t + v_t \qquad (6)$$

where v_t is a white noise error. Equations (5) and (6) give a rational expectations-cum-uncovered interest parity relationship:

$$i_t = i_t^* + \Delta e_t + v_t \qquad (7)$$

which can be obtained from

$$i_t = \delta_1 i_t^* + \delta_2 \Delta e_t + \varepsilon_t \qquad (8)$$

under $\delta_1 = \delta_2 = 1$ and ε_t is zero-mean stationary. Note that, ε_t can also be defined as $v_t + u_t$, with u_t being a time-varying risk premium.

From (8) it can be inferred that a cointegration of the variables with unitary coefficients supports the UIP hypothesis. Compared to the PPP, the UIP has been subject to relatively less scrutiny in the empirical literature. However, the evidence appears to be conflicting also for the UIP[4]. Equation (8) suggests that the UIP hypothesis can be rejected in the presence of non-stationary time varying risk premium and systematic expectation errors. Limited international capital mobility, changes in the term structure of interest rates, non-linear dynamics are amongst the other explanations. As for the PPP, there is no generally accepted theory prior restriction on the endogeneity/exogeneity status of the variables forming the UIP. Thus, an invalid conditioning may be another reason for the failure of studies testing the UIP employing a single-equation method.

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[3] See MacDonald (1999), Obstfeld and Rogoff (2000) and Taylor (2001) for recent reviews.
[4] McCallum (1994) and Flood and Rose (1999) provide the recent accounts. Note that, under a maintained hypothesis that the Fisher parity holds for each of the countries, equation (8) gives a relationship between real exchange rate and real interest rate differential. See Edison and Melick (1999) and MacDonald and Nagayasu (2000) for recent applications.

The PPP is based on the arbitrage in goods market, hence postulated as an adjusting mechanism for the current account (ca) equilibrium. Equilibrium in capital account (ka), on the other hand, may need adjustments in the variables determining the UIP. By definition, balance of payments (bop) consists of the sum of ca and ka. As a disequilibrium in one market may have repercussions on the other, the two international parity conditions may not be independent of each other in the long run evolution of the bop equilibrium.

Johansen and Juselius (1992), Juselius (1995), MacDonald and Marsh (1997) and Juselius and MacDonald (2000) propose an approach taking into account both asset and good market adjustment dynamics by combining both international parities. This approach allowing for interactions among prices, interest rates and exchange rates is referred to as capital enhanced equilibrium exchange rates, or CHEERs by MacDonald (2000).

The main idea of the CHEERs approach is that non-stationary deviations from the PPP and UIP forms a stationary relationship consistent with the interdependence of adjustments in asset and good markets towards equilibrium. Juselius (1995) has suggested expressing this as:

$$(\omega_1(PPP)_t - \omega_2(UIP)_t) \sim I(0) \qquad (9)$$

where ω_1 and ω_2 enter to allow for the effects of temporal aggregation and a possible weak correspondence between theoretical and observed variables. The approach involves testing for structurally identified cointegration relationships between the variables forming the parities. Following section presents the results of the application of this approach to the Turkish data.

III. DOES PPP AND UIP NEED EACH OTHER? THE TURKISH EVIDENCE

We start with a five-variable system z_t = [Δp_t, Δp_t^*, Δe_t, i_t, i_t^*] which allows us to test the validity of the CHEERs approach for the Turkish data. In the system, p and p* are the logs of the Turkish and US consumer price indices, respectively, e is the log of the TL/$US spot exchange rate, i = log (1 + R/100), i* = log(1 + R*/100), R is the 3 month time deposit rate in Turkey (TL) and R* is the 3 month LIBOR rate ($US), and Δ denotes the first difference operator. All series[5] are seasonally unadjusted. The sample period is from 1986:1 to 1999:4 covering the recent floating exchange rate regime and domestic and international capital market liberalization.

The integration properties of the variables are investigated by conducting augmented Dickey-Fuller [ADF] tests with the lag length (k) selected to remove any manifest serial correlation. The results recorded in Table 1 suggest that each of the variables in z_t except Δe_t is integrated of order 1, I(1). The evidence for Δe_t is less conclusive[6]. From its ADF statistics Δe_t might be I(0). Considering the possible low power of the univariate tests, we employed also a multivariate stationary test. The results of the Johansen and Juselius multivariate stationary test reported later in Table 3 suggest the strong rejection of the stationarity null for each of the variables including Δe_t in the system. Since four of the variables are I(1) and neither of them is I(2) or higher, the necessary condition for a valid cointegration inference appears not to be violated even by the ADF evidence.

[5]The data are from the IMF's International Financial Statistics CD-Rom (December 2000) and the Central Bank of the Republic of Turkey database. The data can be obtained by request from the authors at their internet addresses.

[6] Note that the possible stationarity of Δe_t does not necessarily preclude it from being contained in a cointegration space. This is because cointegration tests may still be used even if some series are stationary as Dickey and Rossana (1994) note.

Table 1: Augmented Dickey-Fuller (ADF) test Statistics

	Levels		First Differences
Series	λ_t	λ_m	λ_m
Δp_t	-2.79 (3)	-2.85 (3)	**-8.70** (2)
Δp_t*	-2.67 (2)	-2.26 (2)	**-8.93** (1)
Δe_t	**-3.52** (4)	**-3.45** (4)	**-8.36** (1)
i_t	-2.32 (2)	-1.52 (2)	**-8.20** (1)
i_t*	-2.24 (2)	-1.89 (2)	**-3.24** (1)

Notes: All the test regressions contain a constant term. The equations for λ_t includes also a linear trend. MacKinnon (1991) critical values are -3.50 for λ_t, -2.92 for λ_m. Bold values indicate the rejection of the unit root null at the 5 % level. Numbers in parentheses are the lags (k) used in the augmentation of the regressions.

Table 2: Tests of the cointegration rank

H_0: r	λ_i	λ_{max}	λ_{max} (0.95)	λ_{trace}	λ_{trace} (0.95)
0	0.549	40.56**	33.6	100.26**	70.5
1	0.468	32.14**	27.4	59.71**	48.9
2	0.238	13.89	21.1	27.56	31.5
3	0.166	9.26	14.9	13.67	17.9
4	0.083	4.41	8.1	4.41	8.1

Note: ** denotes the significant tests at the 5 % level

For the 5x1 system z_t we consider a reparameterised VAR(k) process:

$$\Delta z_t = \Pi z_{t-1} + \Gamma_1 \Delta z_{t-1} + ... + \Gamma_{k-1} \Delta z_{t-k+1} + \mu + D_t + \varepsilon_t \tag{10}$$

where μ is a vector of constants, D is a matrix of centered seasonal dummies and ε_t is a multivariate disturbance term. Given that z_t is an I(1) system, (10) represents a vector equilibrium correction mechanism (VECM) if the rank of Π, denoted by r, is such that $0 < r < 5$. For $0 < r < 5$, $\Pi = \alpha\beta'$, where α and β are 5 x r matrices of full column rank. While the columns of the matrix β represent the cointegrating vectors, α gives the matrix of adjustment coefficients. Under these conditions, the VECM for long-run endogenous variables can be written as:

$$\Delta z_{jt} = \alpha\beta' z_{t-1} + \Gamma_1 \Delta z_{t-1} + ... + \Gamma_{k-1} \Delta z_{t-k+1} + \mu + \delta D_t + \varepsilon_t. \tag{11}$$

A necessary condition for the weak exogeneity of a variable in the system, say, z_{jt}, is that no cointegration vector is significant ($\alpha = 0$) in the Δz_{jt} equation. Johansen and Juselius (1992) and Johansen (1995) provide a maximum likelihood procedure for the cointegration analysis of an I(1) system and it is this method we employ in this paper.

Table 2 reports the eigenvalues (λ_i), the maximal eigenvalue (λ_{max}) and trace eigenvalue (λ_{trace}) statistics for the Johansen procedure applied to the five variable system $z_t = [\Delta p_t, \Delta p_t^*, \Delta e_t, i_t, i_t^*]$. The vector autoregression (VAR) contains five lags[7] which is plausable for quarterly data with possible stochastic seasonality, three centered seasonal dummies and an unrestricted constant. Table 3 records various residual diagnostics to test the empirical adequacy of the system with k = 5. Each equation passes all the diagnostics except the normality test. The results for the univariate Jarque-Bera test suggest that it is basically excess kurtosis causing the rejection of normality for the Turkish inflation and exchange rate change equations. The residual

[7] The sequential likelihood ratio (LR) test of system reduction from VAR(5) to VAR(4) yielded $X^2(25) = 53.96$ ($p = 0.001$). Thus the reduction from VAR(5) to VAR(4) appears not to be data-acceptable. This choice is supported also by the Akaike Information Criterion (AIC).

non-normality may not be alarming as cointegration results appear robust to excess kurtosis (Gonzalo, 1994). According to these results, the VAR(5) seems to be a valid approximation of the data generation process, and thus a congruent system for cointegration inference.

The λ_{max} and λ_{trace} statistics reported in Table 2 strongly support the hypothesis that there are two cointegration vectors in the system. Table 4 records some multivariate tests about the properties of the system variables. The test of long-run exclusion investigates whether any of the variables can be excluded from the cointegration space. It is formulated as a zero row in β matrix, i.e. H_{β}: $\beta_{ij} = 0$, $j = 1,\ldots,r$ saying that the variable z_i does not enter the cointegration space. The weak exogeneity test investigates whether any variable affects the remaining variables while it is not affected by them in the long run adjustment process. This test is formulated as a zero row in α matrix, i.e. H_{α}: $\alpha_{ij} =0$, $j = 1,\ldots,r$ where H_{α} is a hypothesis that the variable z_i does not adjust to deviations from the equilibrium. Finally, the multivariate stationarity statistic tests the restriction that the coefficients of the designated variable and the rest of the variables in the maintained cointegration vector are unity and zero, respectively.

According to the results in Table 4, all variables are significant for the cointegration space at the 10% level. The results of the stationarity test strongly suggest the rejection of the stationarity null for each of the variables in z_t. Consistent with the fact that Turkey can be interpreted as a small open economy in the international commodity and financial markets, each of the foreign variables (US interest and inflation rates) appears to be weakly exogenous for the parameters of the cointegration space.

Table 3: System evaluation and single equation residual diagnostics

Multivariate tests:

Residual Autocorrelation	LM_1	$\chi^2(25) = 33.1$
Residual Autocorrelation	LM_4	$\chi^2(25) = 18.0$
Normality: LM		$\chi^2(10) = 48.1$

Univariate tests:	$\Delta^2 p_t$	$\Delta^2 p_t^*$	$\Delta^2 e_t$	Δi_t	Δi_t^*
ARCH(5)	1.79	4.02	3.16	1.06	7.37
Jarque-Bera(5)	**12.13**	2.77	**30.90**	2.15	7.17
Skewness	1.29	0.19	1.39	-0.36	-0.52
Excess Kurtosis	6.12	3.54	10.33	2.47	4.48
R^2	0.95	0.63	0.65	0.94	0.64

Note: Test statistics in bold face are significant at the 5 % level. See Hansen and Juselius (1995) for the details of the tests.

Table 4: Multivariate LR tests about the properties of the system variables

	Δp_t	Δp_t^*	Δe_t	i_t	i_t^*
Long-run exclusion ($\chi^2_{(r)}$)	23.55	9.99	17.02	23.20	4.54
Stationarity ($\chi^2_{(p-r)}$)	16.12	22.87	20.52	18.24	24.97
Weak exogeneity ($\chi^2_{(r)}$)	17.22	1.08	3.54	20.04	1.48

Notes: Under r = 2 and p-r = 3.

Table 5: Cointegration Analysis

	Standardised eigenvectors β		Standardised adjustment coefficients α		
Variable	β_1	β_2	Eq.	α_1	α_2
Δp_t	-3.169	1.000	$\Delta^2 p_t$	**0.244** (3.674)	**-0.668** (-4.622)
Δp_t*	11.705	-0.558	$\Delta^2 p_t$*	-0.007 (-0.797)	0.016 (0.831)
Δe_t	-1.015	-1.228	$\Delta^2 e_t$	**0.518** (1.889)	0.734 (1.230)
i_t	1.000	0.176	Δi_t	**-0.383** (-5.931)	**-0.312** (-2.213)
i_t*	-1.388	-0.471	Δi_t*	-0.014 (-1.125)	-0.019 (-0.705)

Notes: Values in parentheses are the t -values. Bold values are significant at the 5% level.

Table 5 reports the significant cointegration vectors and the corresponding adjustment coefficients. The first vector normalized by i_t and the second vector normalised by Δp_t can be interpreted as representing the long run Turkish interest rate and inflation rate equations, respectively: :

$$i_t = 1.388 i_t^* + 1.015 \Delta e_t + 3.169 \Delta p_t - 11.7 \Delta p_t^*$$

$$\Delta p_t = 0.588 \Delta p_t^* + 1.228 \Delta e_t - 0.176 i_t + 0.471 i_t^*$$

The coefficient estimates are consistent with expected sign priors. The t-statistics for standardized adjustment coefficients are to test the weak exogeneity of the corresponding variable for the parameters of the long run equation[8]. The adjustments of both interest rates and inflation rates to deviations from the corresponding disequilibrium appears to be relatively fast (less than three quarters) as indicated by the magnitude of adjustment coefficients. The first and second cointegrating vectors seem to contain the UIP and PPP relations, respectively.

Johansen and Juselius (1992, 1995) show that, for a q-dimensional system with r cointegration vectors, restrictions on the cointegration structure can be tested by formulating

$$\beta = \{ H_1 \psi_1, .., H_r \psi_r\}$$

where H_i are (q x s_i) design matrices, ψ_i are (s_i x 1) matrices for s_i free parameters. The hypothesis that the first vector describes PPP with unrestricted interest rates ($\beta_1 = [\psi_{11}, \psi_{12}, -\psi_1, \psi_1, -\psi_1]$) and the second vector describes UIP with unrestricted inflation rates ($\beta_2 = [\psi_2, -\psi_2, -\psi_2, \psi_{21}, \psi_{22}]$) can be tested by defining the design matrices:

$$H_1 = \begin{matrix} 0 & 1 & 0 \\ 0 & 0 & 1 \\ -1 & 0 & 0 \\ 1 & 0 & 0 \\ -1 & 0 & 0 \end{matrix} , \qquad H_2 = \begin{matrix} 1 & 0 & 0 \\ -1 & 0 & 0 \\ -1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{matrix}$$

The LR test of these restrictions, distributed as $\chi^2(2)$, is 0.30 and this structure is not rejected with a p-value of 0.86. This result supports the existence of UIP and PPP in the first and second vectors, respectively, with proportionality and symmetry conditions.

The hypothesis that the first vector includes only the UIP and the second includes only the PPP can also be tested. The restricted vectors are:

$$\beta_1 = [0, 0, -1, 1, -1] \text{ and } \beta_2 = [1, -1, -1, 0, 0]$$

[8] Note that these t-statistics has a meaning only if the corresponding cointegration vector

with the corresponding H matrices

$H_3' = [0, 0, -1, 1, -1]$ and $H_4' = [1, -1, -1, 0, 0]$.

This hypothesis is strongly rejected with the LR statistic of 28.6 (distributed as $\chi^2(6)$). Thus, each of the international parity hypotheses is rejected when formulated in isolation. As the adjustments in both asset and commodity markets may interdependent in a financially open economy, it may be plausible to consider the two parity conditions jointly. The design matrices

$$H_5 = \begin{matrix} 0 & 1 \\ 0 & -1 \\ -1 & 0 \\ 1 & 0 \\ -1 & 0 \end{matrix} \qquad H_6 = \begin{matrix} 1 & 0 \\ -1 & 0 \\ -1 & 0 \\ 0 & 1 \\ 0 & -1 \end{matrix}$$

corresponds to testing the hypothesis that the first vector contains the UIP with unitary coefficients restricting two inflation rates to have equal and opposite signs $[\psi_{11}, -\psi_{11}, -1, 1, -1]$, while the second vector contains the PPP with unitary coefficients restricting two interest rates to have equal and opposite sings $[1, -1, -1, \psi_{21}, -\psi_{21}]$. The LR statistic of these restrictions is 6.37 (distributed as $\chi^2(4)$) suggesting non-rejection at the 10% level.

Table 6 reports a structural representation of the cointegration space obtained under the data-acceptable overidentifying restrictions defined by H_5 and H_6. The restricted cointegrating vectors can be written as:

$$i_t - i_t^* - \Delta e_t - 3.958\,(\Delta p_t - \Delta p_t^*)$$

$$\Delta p_t - \Delta p_t^* - \Delta e_t + 0.139\,(i_t - i_t^*)$$

defines a plausible economic relationship (see, Hansen and Juselius (1995)).

That is, the deviations from the UIP can be explained by the inflation differentials, while the deviations from the PPP can be explained by the interest rate differentials.

The results of the Johansen procedure applied to the system z_t = [Δp_t, Δp_t^*, Δe_t, i_t, i_t^*] strongly suggest that neither the PPP nor the UIP hypotheses alone can be valid for the Turkish data. This is a consistent result with Juselius and MacDonald (2000) which rejects strongly the stationarity hypothesis of the pure parity conditions and achieves the stationarity when the interdependence between the parities is allowed. So, what can be deduced from this analysis is that disequilibria in asset and commodity markets should be taken into account jointly in understanding the evolution of the international parities in a financial open economy.

Table 6: A structural representation of the cointegrating space

	Standardised eigenvectors β			Standardised adjustment coefficients α	
Variable	β_1	β_2	Eq.	α_1	α_2
Δp_t	-3.958	1.000	$\Delta^2 p_t$	**0.158** (2.750)	**-0.747** (-4.560)
Δp_t^*	3.958	-1.000	$\Delta^2 p_t^*$	-0.003 (-0.334)	0.020 (0.933)
Δe_t	-1.000	-1.000	$\Delta^2 e_t$	**0.407** (1.718)	0.590 (0.872)
i_t	1.000	0.139	Δi_t	**-0.318** (-5.578)	**-0.383** (-2.354)
i_t^*	-1.000	-0.139	Δi_t^*	-0.007 (-0.642)	-0.019 (-0.634)

Notes: Values in parentheses are t - values. Bold values are significant at 5% level.

IV. CONCLUDING NOTES

This study investigated the validity of two important international parity hypotheses, the UIP and PPP, for the Turkish quarterly data over 1986:1- 1999:4. The PPP and UIP hypotheses are formulated as equilibrium conditions for international commodity and capital markets, respectively. As a disequilibrium in one market may have spillover effects on the other, the characterisation of one parity condition as the general equilibrium level for the whole economy may be seriously misleading. An approach which does not preclude the possible interdependence of asset and commodity markets and thus allowing for interactions among prices, interest rates and exchange rates is provided by the capital enhanced equilibrium exchange rates, or CHEERs model.

The results of the Johansen cointegration analyses for the variable system containing Turkish and US inflation rates, interest rates, and exchange rate suggest the existence of two stationary relationships in the system. The first cointegration appears to explaining the long run evolution of Turkish interest rates whilst the second representing the Turkish inflation rate equation. Consistent with the fact that Turkey can be interpreted as a small country in the international commodity and capital markets, all the foreign variables are found to be weakly exogenous for the parameters of the long run relationships. The weak exogeneity of exchange rates for the long run inflation equation suggests the importance of exchange rate pass-through in designing a disinflation policy. The value of the estimated adjustment coefficients suggests that the adjustments of interest rates and inflation rates to deviations from the long-run equilibrium is relatively fast. The rapid adjustment can be interpreted as reflecting the high cost of being out of equilibrium (or, reflecting that the cost of adjustment is low). This is a plausible result for a financially open economy with a sustained high inflation during the sample period.

The data appear to support the hypothesis that the first vector contains UIP and the second vector contains PPP with proportionality and symmetry conditions. However, each of the international parity hypotheses is strongly rejected when formulated as independent of each other. The results further suggest that the deviations from the PPP can be explained by the interest rates differentials whilst the deviations from the UIP can be explained by the inflation rates differentials. Thus, the Turkish evidence can be interpreted as lending a strong support to the CHEERs approach.

The interdependence of international asset and commodity markets and the consequent interaction between the UIP and PPP has a crucial implication for an exchange rate targeting policy and an exchange rate based stabilization programme. These policies may not be sustainable if they are designed under a maintained hypothesis that the equilibrium exchange rate is determined only by commodity market clearing PPP condition. This is basically because, the adjustment of exchange rates to capital flows due to interest parities may lead to targeted exchange rates substantially diverging from the equilibrium rates for a financially open economy.

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