Testing international parity hypothesis in a multivariate identified co-integrating system: the Turkish evidence

Levent, Korap

Istanbul University Institute of Social Sciences, Besim Ömer Paşa Cd. Kaptan-ı Derya Sk. 34452 Beyazıt /ISTANBUL

2008

Testing international parity hypothesis in a multivariate identified co-integrating system: the Turkish evidence

Levent KORAP

Economist, Marmara University Department of Economics

ABSTRACT

In this paper, a multivariate co-integrating model is constructed upon the Turkish economy to examine the validity of the purchasing power parity and the uncovered interest parity theories simultaneously. Estimation results obtained from the identified co-integrating vectors support *a priori* modelling expectations and yield evidence to the existence of both parities when integrated within each other. However, no evidence is obtained in favor of the two international exchange rate determination parity hypotheses when formulated in isolation. A policy inference derived from the paper can be summarized such that, since the market mechanisms seem to closely affect the long-run course of the nominal exchange rate, exchange rate based stabilization programs should be appreciated by economic agents in a cautious way.

Key words: Purchasing Power Parity ; Uncovered Interest Parity ; Turkish Economy ;

ÖZET

Bu çalışmada, çok değişkenli bir eş-bütünleşim modeli Türkiye ekonomisi üzerine satın alma gücü paritesi ve korunmamış faiz paritesi kuramlarının eşanlı incelenmesi amacıyla oluşturulmaktadır. Tanımlanmış eş-bütünleşik vektörlerden elde edilen tahmin sonuçları *önsel* modelleme beklentilerini desteklemekte ve her iki paritenin birbirleri içerisine bütünleştirilmeleri durumunda varlığına yönelik bulgular üretmektedir. Bununla birlikte, birbirlerinden ayrı olarak formülleştirilmeleri durumunda iki uluslar arası döviz kuru belirlenme paritesi doğrultusunda bulgu elde edilememektedir. Çalışmadan türetilen bir

politika çıkarsaması, piyasa mekanizması parasal döviz kurunun uzun dönemli gelişim yolunu yakından etkiler bir şekilde gözlendiği için, döviz kuru temelli istikrar programlarının iktisadi birimler tarafından ihtiyatlı bir şekilde değerlendirilmesi gerekliliği olarak özetlenebilir.

Anahtar kelimeler: Satın Alma Gücü Paritesi ; Korunmamış Faiz Paritesi ; Türkiye Ekonomisi ;

INTRODUCTION

One of the main recent issues of interest in contemporaneous macroeconomic theories is to examine the fundamental building blocks of exchange rates and interest rates based on the theoretical underpinnings of the exchange rate determination. Revealing the course of exchange rates in a long-run steady-state relationship will help researchers conduct the empirical investigations for testing the coherence of international macroeconomic theories such as purchasing power parity (PPP) and uncovered interest parity (UIP) as well as the theories explaining the determination of exchange rates. Thus, testing the integratedness of such relationships assuming open economy conditions may yield crucial policy conclusions, so that researchers and policy makers can use the knowledge of various arbitrage possibilities and equilibrium conditions in assets and goods markets to design appropriate stabilization policies and to extract the stylized facts of the economies in an ever-increasing complexity of today's gobalizing world economy.

In this paper, our aim is to test whether the PPP and the UIP theories can be supported by the Turkish data, and to extract the long-term information content derived from the interaction between the PPP and the UIP theories, which relates these two contemporaneous international parity conditions to each other. Since investigating the two international parity conditions may not be independent of each other in the long run evolution of the balance of payment equilibrium, such a methodology will enable researchers to relate these parity conditions to the balance of payment components, that is, to current account through an adjustment mechanism in the goods market and also to capital account through adjustments determining the UIP condition. For this purpose, the next section provides a theoretical background combining these parity conditions in the goods and the assets markets. Section 2 highlights the preliminary data issues and estimation methodology over which an empirical model is estimated in section 3. The last section summarizes results to conclude the paper.

I. THEORETICAL BACKGROUND

Integratedness of the PPP and the UIP conditions has been well-studied both theoretically and empirically in some papers such as Johansen and Juselius (1992), Juselius (1995), MacDonald (2000), Caporale et al. (2001) and Özmen and Gökcan (2004). Using the sticky-price model *á la* Dornbusch (1976), let us assume a long-run steady-state economic relationship through which the PPP and the UIP conditions tend to hold. Under the perfect capital mobility, the UIP will take the form of:

$$\Delta s_t^e = (i_{t,k} - i_{t,k}^*) \tag{1}$$

where Δs_t^e is the expected depreciation rate of exchange rate, $i_{t,k}$ the *k*-period yield on the domestic instrument, and $i_{t,k}^*$ the corresponding rate on the foreign instrument. Eq. 1 assumes that expected change in exchange rate is a function of the gap between nominal exchange rate (s_t) and equilibrium exchange rate (s_t^{eq}) in their natural logarithms with a proportion of constant κ which relates this gap to the expected change in the nominal exchange rate:

$$\Delta s_t^e = -\kappa(s_t - s_t^{eq}) \tag{2}$$

Thus the larger the deviation of nominal exchange rate from the hypothetical long-run equilibrium exchange rate, the lower the expected exchange rate depreciation in the future periods to restore the equilibrium conditions. Furthermore, due to the PPP relationship, equilibrium exchange rate will be a function of domestic (p_t) and foreign price levels (p_t^*):

$$s_t^{eq} = (p_t - p_t^*) \tag{3}$$

Rearranging Eqs. (1)-(3) yields:

$$s_t = (p_t - p_t^*) - (1/\kappa)(i_t - i_t^*) \quad (4)$$

In line with the Dornbusch (1976) sticky price exchange rate determination model, the determination of nominal exchange rate in Eq. (4) is a function of both price level and interest differentials.[1] Özcan and Gökmen (2004) note that such a formulation of system of equations given above has been of special interest for researchers and policy makers since the determination of exchange rate takes account of not only international price differentials but also of interest rate differentials. This case provides additional knowledge in the exchange rate determination process while considering the adjustment process of exchange rates due to the risk-adjusted interest parities for financially open economies.

II. DATA AND METHODOLOGY

A. DATA

In this section, an empirical model has been constructed for the Turkish economy to examine the existence of the PPP and the UIP conditions simultaneously for the period 1987Q1 – 2007Q2 using quarterly observations. All the data used are taken from the electronic data delivery system of the Central Bank of the Republic of Turkey (CBRT) for domestic variables and from the FRB of St. Louis for the foreign variables, and indicate seasonally unadjusted values in natural logarithms except the domestic and foreign interest rates, which are used in their linear forms.

For domestic price level (p_t) and foreign price level (p_t^*) data, the gross domestic product (GDP) deflators from the Turkish and the US economies are used, while the spot exchange rate of YTL/US\$ (s_t) has been considered for the nominal exchange rate variable. The Treasury interest rate data (i_t) are used for the domestic short-term interest rates, which are the maximum rate of interest on the Treasury bills whose maturity at most twelve months, while one-year Treasury bond rate for the US economy (i_t^*) is used for the foreign interest rate data. Two impulse-dummy variable which take on values of unity from 1994Q1 till 1994Q4 and

---

[1] For informative purposes, note that:

$$i_t - i_t^* = \kappa(s_t - p_t + p_t^{eq}) \Rightarrow s_t = (-1/\kappa)(i_t - i_t^*) + (p_t - p_t^*)$$

from 2001Q1 till 2001Q4, concerning the financial crises occurred in 1994 and 2001, are considered exogeneous variables, as well.

B. TESTING UNIT ROOTS ALLOWING FOR ENDOGENOUS BREAKS

Spurious regression problem analysed by Granger and Newbold (1974) indicates that using non-stationary time series steadily diverging from long-run mean will produce biased standard errors and unreliable correlations within the regression analysis. This means that the variables must be differenced (*d*) times to obtain a covariance-stationary process. However, conventional tests for identifying unit roots in a time series are criticized strongly in the contemporaneous economics literature when they have been subject to structural breaks which yield biased estimations. Perron (1989) in his seminal paper on this issue argues that conventional unit root tests used by researchers do not consider that a possible known structural break in the trend function may tend too often not to reject the null hypothesis of a unit root in the time series when in fact the series is stationary around a one time structural break. Contrary to the general evidence of many earlier papers which conclude that the US post-war GNP series can be represented by a unit root process, Perron (1989) finds that if the first oil shock in 1973 is treated as a structural breakpoint in the trend function, then the unit root hypothesis of the US post-war GNP series can be rejected in favor of a trend stationary hypothesis.

Selecting the date of structural break, that is, assuming that the time of break is known *a priori*, however, may not be the most efficient methodology. The actual dates of structural breaks may not be coincided with the dates chosen exogenously. To address this issue, several methodologies including Perron (1990), Zivot and Andrews (1992) and Banerjee, Lumsdaine and Stock (1992) have been suggested to allow for the determination of the date of structural breaks endogenously. Considering these issues, in our paper, we follow the Zivot and Andrews (1992) (henceforth ZA) methodology, allowing the data to indicate breakpoint endogenous rather than imposing a breakpoint from outside the system.

The ZA methodology as a further development on Perron (1989) methodology can be explained by considering three possible types of structural breaks in a series, i.e., Model A assuming shift in intercept, Model B assuming change in slope and Model C assuming change in both intercept and slope. For any given time series y_t, ZA (1992) test the equation of the form:

$$y = \mu + y_{t\text{-}1} + e_t \qquad (5)$$

Here the null hypothesis is that the series y_t is integrated without an exogenous structural break against the alternative that the series y_t can be represented by a trend-stationary I(0) process with a breakpoint occurring at some unknown time. The ZA test chooses the breakpoint as the minimum *t*-value on the autoregressive y_t variable, which occurs at time 1 < TB < T leading to $\lambda = TB / T$, $\lambda \in \left| 0.15, 0.85 \right|$, by following the augmented regressions:

Model A:

$$y_t = \mu + \beta t + \theta DU_t(\lambda) + \alpha y_{t\text{-}1} + \sum_{i=1}^{k} c_j \, \Delta y_{t\text{-}j} + \varepsilon_t \qquad (6)$$

Model B:

$$y_t = \mu + \beta t + \gamma DT_t(\lambda) + \alpha y_{t\text{-}1} + \sum_{i=1}^{k} c_j \, \Delta y_{t\text{-}j} + \varepsilon_t \qquad (7)$$

Model C:

$$y_t = \mu + \beta \mathrm{t} + \theta DU_t(\lambda) + \gamma DT_t(\lambda) + \alpha y_{t\text{-}1} + \sum_{i=1}^{k} c_j \, \Delta y_{t\text{-}j} + \varepsilon_t \qquad (8)$$

where DU_t and DT_t are sustained dummy variables capturing a mean shift and a trend shift occuring at the break date respectively, i.e., $DU_t(\lambda) = 1$ if $t > T\lambda$, and 0 otherwise; $DT_t(\lambda) = t - T\lambda$ if $t > T\lambda$, and 0 otherwise. Δ is the difference operator, k is the number of lags determined for each possible breakpoint by one of the information criteria, and ε_t is assumed to be i.i.d. error term. The ZA method runs a regression for every possible break date sequentially and the time of structural changes is detected based on the most significant *t*-ratio for α. To test the unit root hypothesis, the smallest *t*-values are compared with a set of asymptotic critical values estimated by ZA. We must note that critical values in the ZA methodology are larger in absolute sense than the conventional ADF critical values since the

ZA methodology is not conditional on the prior selection of the breakpoint. Thus, it is more difficult to reject the null hypothesis of a unit root in the ZA test. For the appropriate lag length, we consider the Schwarz's Bayesian information criterion (SBIC)-minimizing value.

Table 1: Zivot-Andrews Unit Root Test[a,b]

Var.	Intercept		Trend		Both	
	k	min t TB	k	min t TB	k	min t TB
p_t	0	-3.619 (1992Q1)	0	-2.391 (1992Q2)	0	-3.630 (1992Q1)
s_t	1	-1.537 (2001Q2)	1	-3.861 (2001Q3)	1	-3.853 (2001Q2)
p_t^*	3	-4.080 (1997Q2)	3	-4.135 (2002Q4)	3	-4.046 (2004Q1)
i_t	0	-4.503 (1996Q1)	0	-4.480 (1994Q1)	0	-4.480 (1994Q1)
i_t^*	1	-2.565 (2004Q2)	1	-2.659 (2003Q4)	1	-3.160 (2001Q1)

[a] Estimation with 0.15 trimmed. Lag length is determined by Schwarz's Bayesian information criterion. min t is the minimum *t*-statistic calculated.

[b] Critical values – intercept: -5.43 (1%), -4.80 (5%); trend: -4.93 (1%), -4.42 (5%); both: -5.57 (1%), -5.08 (5%)

In Tab. 1, ZA unit root tests alowing endogenous break in the time series used indicate that unit root null hypothesis cannot be rejected for all the series.

C. ESTIMATION METHODOLOGY

In order to test for a stationary relationship among the variables, we apply to the multivariate co-integration techniques proposed by Johansen (1988), Johansen and Juselius (1990), and Johansen (1995) and search for whether it is possible to extract any steady-state knowledge from the long-run variable space. Let us assume a z_t vector of non-stationary n endogenous variables and model this vector as an unrestricted vector autoregression (UVAR) involving up to k-lags of z_t:

$$z_t = \Pi_1 z_{t-1} + \Pi_2 z_{t-2} + \ldots + \Pi_k z_{t-k} + \varepsilon_t \qquad (9)$$

where ε_t is the N(0, σ^2) disturbance term, assuming an expected value with a normally distributed zero-mean and constant variance, and z is ($nx1$) and the Π_i is an (nxn) matrix of parameters. Eq. 9 can be rewritten leading to a vector error correction (VEC) model:

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + \Gamma_2 \Delta z_{t-2} + \ldots + \Gamma_{k-1} \Delta z_{t-k+1} + \Pi z_{t-k} + \varepsilon_t \quad (10)$$

where:

$$\Gamma_i = -I + \Pi_1 + \ldots + \Pi_i \;\; (i = 1, 2, \ldots, k\text{-}1) \quad (11)$$

and:

$$\Pi = I - \Pi_1 - \Pi_2 - \ldots - \Pi_k \quad (12)$$

This specification of the system of variables carries on the knowledge of both the short- and the 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 as a matrix of error correction terms, while β is a matrix of long-run coefficients such that $\beta' z_t$ embedded in Eq. 10 represents up to (n-1) cointegrating relations in the multivariate model which ensure that z_t converge to their long-run steady-state solutions. Note that all terms in Eq. 10 which involve Δz_{t-i} are $I(0)$ while Πz_{t-k} must also be stationary for $\varepsilon_t \sim I(0)$ to be white noise of an N(0, $\sigma_\varepsilon{}^2$) process. Following Johansen (1992), an intercept and a linear trend are restricted into the long run variable space in line with the Pantula principle.

III. ESTIMATION RESULTS

A. RANK TEST RESULTS

In line with these econometric model specification issues, we now test the empirical validity of both parities in a multivariate co-integrating framework. For this purpose, a UVAR model is constructed, using an endogenous variable vector $(p_t, s_t, p_t^*, i_t, i_t^*)'$ of the potential long-run co-integrating space. For the lag length of the UVAR model, we consider the sequential modified LR statistics employing small sample modification, which compare the modified LR statistics to the 5% critical values starting from the maximum lag chosen and decreasing the lag one at a time until first getting a rejection. Considering the maximum lag of 5, the reduction of the system from 5 to 4 lags is accepted by an LR statistic 20.77 but is first rejected when we test the reduction from 4 to 3 lag orders by an LR statistic 45.26:

Table 2: Co-integration Rank Tests

Null hypothesis	r=0	r≤1	r≤2	r≤3	r≤4
Eigenvalue	0.63	0.49	0.28	0.22	0.03
λ trace	174.44*	97.39*	46.18*	20.78	2.06
5% critical value	88.80	63.88	42.92	25.87	12.52
λ max	77.05*	51.21*	25.41	18.72	2.06
5% critical value	38.33	32.12	25.82	19.39	12.52

* denotes rejection of the hypothesis at the 0.5 level.

Unrestricted Co-integrating Coefficients

p_t	s_t	p_t^*	i_t	i_t^*	*trend*
8.200483	-10.23726	-142.6694	1.375030	-47.27569	1.076261
-4.873663	8.265032	125.9455	-0.666015	11.10795	-1.008706
-5.471619	3.041515	-64.50218	8.029622	-23.39144	0.633207
-18.60886	18.22960	152.9311	-8.457413	20.52780	-0.720105
4.816002	-0.447238	-8.373522	-4.722794	45.57301	-0.457079

Unrestricted Adjustment Coefficients (alpha) ('D' indicates the first difference operator)

D(p_t)	-0.030753	-0.001141	0.014153	-0.003198	-0.001395
D(s_t)	-0.038370	-0.015281	-0.001496	-0.021056	-0.002487
D(p_t^*)	-0.000120	-0.000590	0.000229	0.000128	0.000106
D(i_t)	-0.067739	-0.051782	-0.033531	0.005957	-0.016609
D(i_t^*)	0.001499	-0.001340	0.001117	-0.000613	-0.000138

In Tab. 2, we find that trace test suggest three and max-eigen test two potential co-integrating vectors lying in the long-run variable space. For we *a priori* hypothesize two international macroeconomics parity conditions in the long-run variable space, we accept that two stationary relationships, which are found common by both likelihood statistics, can be identified by means of economics theory to obtain independent co-integrating vectors. As Caporale et al. (2001) emphasize, no clear-cut assessment has just been made to impose economic relationships. But a cursory examination of the first two rows of the unrestricted co-integrating coefficients indicates that the first vector seems to satisfy the PPP relationship, and that both the first and the second rows reveal that domestic and foreign interest rates seem to be in a positive stationary relationship, which can give support to the UIP hypothesis. When the unrestricted adjustment coefficients are considered, for which a value highly close to zero would mean weakly exogenous characteristic of the variable, the negative non-zero value for nominal exchange rate, domestic price level and foreign interest rate in the first vector and for domestic interest rate in the second vector support the existence of both co-integrating relationships in our long-run variable space. Considering these preliminary investigation of Tab. 2, we need to normalize the co-integrating vectors to extract more information from the data. If we briefly write down the *ex-ante* relationships in a theoretical sense, we search for evidence of the form:

$$\beta' = \begin{bmatrix} \beta_1' \\ \beta_2' \end{bmatrix} = \begin{bmatrix} b_{11}\ b_{12}\ b_{13}\ b_{14}\ b_{15} \\ b_{21}\ b_{22}\ b_{23}\ b_{24}\ b_{25} \end{bmatrix} = \begin{bmatrix} 1\ \text{-}1\ \text{-}1\ 0\ 0 \\ 0\ 0\ 0\ 1\ \text{-}1 \end{bmatrix} \tag{13}$$

where β_1' and β_2' are the vectors carrying the knowledge of unrestricted coefficients and b_{ij}'s refer to the coefficient of j^{th} endogenous variable within i^{th} co-integrating vector. Above β_1' and β_2' are assigned to the PPP and the UIP relationships with appropriate homogeneity and symmetry restrictions, respectively. Eq. 14 and Eq. 15 below report the unrestricted normalized vectors on domestic price level and domestic interest rate under the assumption r=2:[2]

$$\beta'_{x1} = p_t - 0.56s_t - 4.54p_t^* + 0.21i_t - 11.07i_t^* + 0.02trend + 21.78 \tag{14}$$

$$\beta'_{x2} = 4.49 - 2.05s_t - 10.94p_t^* + i_t - 53.64i_t^* - 0.01trend + 52.81 \tag{15}$$

[2] Normalizations are carried out by EViews 5.1.

Above β'_{x1} and β'_{x2} are the estimated co-integrating vectors with no *a priori* hypothesized restrictions. We can easily notice that the signs of the unrestricted coefficients verify the theoretical expectations for both vectors. We should specify that any linear combination of the stationary vectors is itself a stationary vector, and thus, the estimates produced for any particular column of the long-run cointegrating coefficients are not necessarily unique. Such a case requires restrictions imposed on the co-integration space and motivated by economic arguments to obtain unique vectors lying within that space (Harris, 1995).

B. IDENTIFICATION

Applying joint restrictions to all of the separate β vectors may not be the most appropriate way to start an econometric analysis unless economic theory is informative on the hypotheses that should be tested. Thus we test whether the knowledge of first vector can be attributed to identification of the PPP relationship by leaving domestic and foreign interest rates unrestricted. Then we test whether the second vector can be attributed to the identification of a stationary relationship between domestic and foreign interest rates, assuming other variables as unrestricted. Since we are now cabaple of identifying the co-integrating vectors, we can obtain the standard errors, thus, the *t*-statistics for the unrestricted variables in each co-integrating vector:

$$\beta' x_1 = p_t - s_t - p_t^* - 0.61 i_t + 20.37 i_t^* - 0.09 trend + 14.08 \tag{16}$$

t-stats. (-1.05) (3.86)

$$\beta' x_2 = 67.51 p_t - 86.51 s_t - 1058.55 p_t^* + i_t - i_t^* + 6.91 trend + 5170.59 \tag{17}$$

t-stats. (4.61) (-6.00) (-6.75) (5.79)

Eq. 16 and Eq. 17 give the identified vectors obtained from the co-integrating space. We give evidence in favor of the restrictions leading to that the absolute PPP and UIP hypotheses hold for the Turkish data using $\chi^2(1)$=2.51 (prob. 0.11) under the H_0 hypothesis.[3] Note that under

---

[3] Following Harris(1995), the degrees of freedom for this identification procedure can be obtained by using the formula $v = \sum_i (n - r + 1 - s_i)$ where n is the number of co-integrating vectors and s_i is the number of unrestricted parameters in each vector. In our case, $(n - r + 1) = (5 - 2 + 1) = 4$ and for the first vector $v_1 = 4 - 3 = 1$, for the second vector $v_2 = 4 - 4 = 0$. Thus the degrees of freedom considered for the identification of the co-integrating system equals 1, i.e. $v_1 + v_2 = 1 + 0 = 1$.

the PPP restrictions in the firt vector, the remaining variables retain the signs implied by the UIP hypothesis, and that under the UIP restrictions in the second vector, the remaining variables retain the signs implied by the PPP hypothesis. Following MacDonald et al. (2000) and Özmen and Gökcan (2004), however, when we test the equality of domestic and foreign interest rates with opposite signs in the vector, i.e. $\beta_1' = [1\ -1\ -1\ b_{14}\ -b_{14}]$, and of domestic and foreign price levels with opposite signs in the second vector, i.e. $\beta_2' = [b_{21}\ -b_{21}\ -b_{21}\ 1\ -1]$, we reject H_0 hypothesis of coefficient restrictions using $\chi^2(6) = 38.47$ (prob. 0.00) against $\chi^2(6)$-table value = 12.59. We then test whether the first vector includes only the long-run stationary knowledge of the PPP relationship, i.e. $\beta_1' = [1\ -1\ -1\ 0\ 0]$ and the second vector includes only the UIP relationship, i.e. $\beta_2' = [0\ 0\ 0\ 1\ -1]$, and estimate the same system below:

$$\beta' x_1 = p_t - s_t - p_t^* - 0.19 trend + 18.95 \quad (18)$$

t-stats. (-7.39)

$$\beta' x_2 = i_t - i_t^* + 0.24 trend - 10.91 \quad (19)$$

t-stats. (7.35)

yielding $\chi^2(6) = 38.54$ (prob. 0.00). Eq. 18 and Eq. 19 indicate that identifying assuptions for the coefficient restrictions of independent vectors have been rejected. These estimation results support the findings of Özmen and Gökcan (2004) and indicate that no evidence is found in favor of the two international parity hypotheses when forulated in isolation.

Finally, we find no vector serial correlation problem using LM(1) = 30.64 (prob. 0.12) and LM(4) = 30.34 (prob. 0.21), where LM(1) and L(4) are the 1st and 4th order vector error correction (VEC) system residual serial correlation lagrange multiplier statistics under the null of no serial correlation. However, vector normality tests reject the normality of VEC model residuals using $JB\chi^2(10) = 51.94$ (prob. 0.00) where JB is the Jarque-Bera VEC residual normality statistic. But Gonzalo (1994) reveals that Johansen multivariate co-integration methodology performs better than other estimation methods even when the errors non-normally distributed.

CONCLUDING REMARKS

In this paper, a structural vector error correction model of the Turkish economy is tried to be constructed, and the validity of the purchasing power parity (PPP) and uncovered interest parity (UIP) are examined, which assume simultaneous equilibrium in the long run by means of using multivariate identified co-integrating analysis. We find that when we integrate both parities within each other by permitting interest rates in the PPP vector and price levels and nominal exchange rate in the UIP vector to lie unrestrictedly in addition to the symmetry and homogeneity restrictions required for identification of each parities, we accept the validity of both parities in the Turkish economy. However, no evidence is found in favor of the two international parity hypotheses when formulated in isolation. A policy conclusion derived from the paper can be summarized such that exchange rate based stabilization programs should be appreciated by economic agents in a cautious way, since the market mechanisms seem to closely affect the long run course of the nominal exchange rate.

ACKNOWLEDGEMENT

The author is grateful to Gökhan Karabulut of the Istanbul University Department of Economics for the philosophy doctorate Open Economy Macroeconomics courses including the issues argued in this paper.

REFERENCES

BANERJEE, A, R.L. LUMSDAINE, and J.H. STOCK; (1992), “Recursive and Sequential tests of the Unit Root and Trend Break Hypothesis”, **Journal of Business and Economic Statistics**, 10, pp. 271-87.

CAPORALE, M.G., S. KALYVITIS, and N. PITTIS; (2001), “Testing for PPP and UIP inan FIML Framework: Some Evidence for Germany and Japan”, **Journal of Policy Modeling**, 23, pp. 637-50.

DORNBUSCH, R.; (1976), “Expectations and Exchange Rate Dynamics”, **Journal of Political Economy**, 84(6), pp. 1161-176.

GONZALO, J.; (1994), “Five Alternative Methods of Estimating Long-run Equilibrium Relationships”, **Journal of Econometrics**, 60, pp. 203-33.

GRANGER, C.W.J. and P. NEWBOLD; (1974), “Spurious Regression in Economics”, **Journal of Econometrics**, 60, pp. 203-33.

HARRIS, R.I.D.; (1995), **Using Cointegration Analysis in Econometric Modelling**, Prentice Hall.

JOHANSEN, S.; (1988), “Statistical Analysis of Cointegration Vectors”, **Journal of Economic Dynamics and Control**, 12, pp. 231-54.

JOHANSEN, S.; (1992), “Determination of Cointegration Rank in the Presence of a Linear Trend”, **Oxford Bulletin of Econoics and Statistics**, 54(3), pp. 383-97.

JOHANSEN, S.; (1995), **Likelihood-based Inference in Co-integrated Vector Autoregressive Models**, Oxford University Press.

JOHANSEN, S. and K. JUSELIUS; (1990), “Maximum Likelihood Estimation and Inference on Cointegration-with Applications to the Demand for Money”, **Oxford Bulletin of Economics and Statistics**, 52, pp. 169-210.

JOHANSEN, S. and K. JUSELIUS; (1992), “Testing Structural Hypotheses in a Multivariate Cointegration Analysis of the PPP and the UIP for UK”, **Journal of Econometrics**, 53, pp. 211-44.

JUSELIUS, K.; (1995), “Do Purchasing Power Parity and Uncovered Interest Parity Hold in the Long-run? An Example of Likelihood Inference in a Multivariate Time-series Model”, **Journal of Econometrics**, 69, pp. 211-40.

MACDONALD, R; (2000), “Concepts to Calculate Equilibrium Exchange Rates: An Overview”, **Economic Research Group of the Deutsche Bundesbank Discussion Paper**, No. 3/00.

ÖZMEN, E. and GÖKCAN, A.; (2004), “Deviations from PPP and UIP in a Financially Open Economy: The Turkish Evidence”, **Applied Financial Economics**, 14(11), pp. 779-84.

PERRON, P.; (1989), “The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis”, **Econometrica**, 57, pp. 1361-401.

PERRON, P.; (1990), “Testing for a Unit Root in a Time Series with Changing Mean”, **Journal of Business and Economic Statistics**, 8, pp. 153-62.

ZIVOT, E. and D.W.K. ANDREWS; (1992), “Further Evidence on Great Crash, the Oil Price Shock and the Unit Root Hypothesis”, **Journal of Business and Economic Statistics**, 10, pp. 251-70.