Uşak Üniversitesi Sosyal Bilimler Dergisi (2008) 1/2, 87-101

TESTING EXCHANGE RATE DETERMINATION MODEL FOR YTL/US$: EVIDENCE FROM HIGH FREQUENCY DATA

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Abstract

In this paper, exchange rate determination mechanism of YTL/US$ is examined for the 1986M01-2007M08 period of monthly frequency data. Using a theoretical approach based on the monetary model exchange rate determination mechanism, estimation results obtained from contemporaneous multivariate co-integration methodology indicate that YTL/US$ nominal exchange rate is co-integrated with the fundamentals suggested by economics theory. Estimation results reveal that there exists a positive relationship between nominal exchange rate and relative money supply differential and that an increase in the relative income differential would lead to an appreciation of the domestic currency against the US$. Besides, relative interest differential variable has a positive relationship with the nominal exchange rate as *a priori* hypothesized. However, no significant effect of relative inflation differentials on the nominal exchange rate has been found.
Key Words: Exchange Rates; Economic Fundamentals; Co-integration

Özet

Bu çalışmada, YTL/US$ döviz kuru belirlenme mekanizması aylık gözlem aralığı kullanılarak 1986M01-2007M08 dönemi için incelenmektedir. Parasal model döviz kuru belirlenme mekanizmasına dayalı olarak oluşturulan kuramsal bir yaklaşım doğrultusunda, çağdaş çok değişkenli eş-bütünleşim yöntemi kullanılarak elde edilen tahmin sonuçları YTL/US$ parasal döviz kurunun iktisat kuramı tarafından önerilen temellerle eş-bütünleşik bir ilişki içerisinde olduğunu göstermektedir. Tahmin bulguları parasal döviz kuruyla göreceli para arzı farkı arasında pozitif bir ilişkinin varlığını ve göreceli gelir farkındaki bir artışın yerli paranın US$ karşısında değerlenmesine neden olduğunu ortaya koymaktadır. Ayrıca, göreceli faiz farkı değişkeni *önsel* olarak varsayıldığı gibi parasal döviz kuruyla pozitif bir ilişki içerisinde bulunmuştur. Bununla birlikte, göreceli enflasyon farkının parasal döviz kuru üzerinde herhangi anlamlı bir etkisi bulunamamıştır.
Anahtar Kelimeler: Döviz Kurları; İktisadi Temeller; Eş-bütünleşim

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I. Introduction

Modeling determinants of exchange rates provides policy makers significant knowledge of whether various parity conditions between exchange rates constructed in modern open economy macroeconomics can be incorporated within each other to explain the long-run course of exchange rates. Such a phenomenon has been of special importance especially for developig countries since policy makers tend to canalize use of exchange rates to gain an *ex-ante* designed macroeconomic growth performance as well as to break the inertial nature of prices dominated in the economy in fighting inflation. Thus, investigation of the long-run course of exchange rates enables researchers to examine efficiency of the discretionary policies and to extract expected consequences of policy implementations based on exchange rate determination.

The Turkish economy as a small open developing country which was subject to chronic two-digits inflationary framework over a 20 years period till the early 2000s can be considered an interesting case study to examine the issue of exchange rate determination. By the beginning of 2000, an anti-inflationary stabilization program based on a quasi-currency board was established to fight domestic inflation and policy makers aimed at mainly forming the expectations of economic agents in pricing behavior following the policy based on nominal exchange anchor. Although seemed to be successful in bringing inflation down as the one-half of the initial level for the first 10 months realization, the subsequent two economic crisis periods ended the program with a depreciating real income. Following such developments, the Turkish economy has still been trying to establish an inflation targeting framework supported by free-floating exchange rate system.

Following the seminal paper by Meese and Rogoff (1983), many papers have been attributed to modeling the behavior of exchange rates so as to see whether monetary fundamentals are able to explain long-run course of exchange rates. Among many others, MacDonald and Taylor (1993), McNown and Wallace (1994), Mark (1995), MacDonald and Marsh (1997), Kilian (1999), Groen (2000), Bahmani-Oskooee and Kara (2000), Mark and Sul (2001), Civcir (2003) and Rapach and Wohar (2004) examine the validity of the exchange rate determination model. In this paper, our aim is to examine empirical validity of the monetary model of exchange rate determination for the Turkish economy. We first highlight the construction of a simple flexible price monetary exchange rate model and give a variant of this model examined in the economics literature. Then an empirical model using contemporaneous multivariate co-integration tecniques is constructed for the Turkish economy. Finally, the last section summarizes results and concludes.

II. Model Construction

Following Neely and Sarno (2002) we begin our analysis by examining the flexible price monetary model (FPMM). Model is constructed in line with the assumptions based on the quantity theory of money (QTM) and the purchasing power parity (PPP) relating the changes in price level and exchange rate to the money supply changes. McNown and Wallace (1994) express that if the demand for money is stable, monetary approach is a richer formulation than the PPP combining money demand variables with money supplies in the determination of exchange rate. Thus the model assumes that determination of supply of and demand for money leads to the existence of a stable money demand function. As Neely and Sarno (2002) noted, perfect capital mobility assumption implicit in the model also requires that the real interest rate be exogenous in the long run and be determined in the world markets.

Consider that equilibrium in the monetary markets for the domestic and foreign country requires:

$$m_t = p_t + \alpha y_t - \beta i_t \tag{1}$$

$$m_t^* = p_t^* + \alpha^* y_t^* - \beta^* i_t^* \tag{2}$$

where m_t, p_t, y_t, and i_t denote the measure of money supply, price level, real income and the interest rate at any time t respectively, which are all in natural logarithms except the interest rate, while those carrying an asterisk represent the identical foreign variables. The coefficients α and β are the positive constants used for the income elasticity of demand for money and interest rate semi-elasticity, respectively.

The second building block of the monetary model assumes that absolute PPP would hold and that prices in two currencies would tend to be equalized *via* exchange rate movements resulted from goods market arbitrage. Writing down such a relationship below in Eq. 3 yields:

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

where s_t respresents the domestic price of foreign currency, i.e., nominal exchange rate, in natural logarithms. Subtracting Eq. 2 from Eq. 1, solving for $(p_t - p_t^*)$ and inserting the result into Eq. 3 yield the FPMM of nominal exchange rate determination:

$$s_t = (m_t - m_t^*) - (\alpha y_t - \alpha^* y_t^*) + (\beta i_t - \beta^* i_t^*) \quad (4)$$

Let us assume as a simplifying assumption for the ease of applying to the modern time series estimation techniques that the income elasticities and interest rate semi-elasticities of money demand equal each other for the home and foreign countries:

$$s_t = (m_t - m_t^*) - \alpha (y_t - y_t^*) + \beta (i_t - i_t^*) \quad (5)$$

In line with Eq. 5 we expect a positive relationship between nominal exchange rate and relative money supply, and a negative relationship between relative income level and nominal exchange rate. Thus the larger the home relative to the foreign money supply the larger would be the nominal exchange rate, and the larger the home relative to the foreign real income level the lower would be the nominal exchange rate.

Such a specification would differ from the Mundell-Fleming model in that the latter approach assumes that there would be a negative relationship between relative income level and exchange rate since the depreciating trade balance following a boom in real income thus in imports volume would require a depreciation of domestic currency in order to restore equilibrium. Whereas, FPMM assumes that increases in domestic real income *ceteris paribus* would lead to an excess demand for domestic money and in turn agents would reduce their expenditures in order to increase their real money balances leading to a fall in prices. Appreciation of domestic currency *via* the PPP would then restore the equilibrium.

Based on the the FPMM given in Eq. (5), a variant of the model can be drived in line with the assumptions considered through the economics theory. Dornbusch (1976) and Frankel (1979) make a difference between FPMM which can also be called the ‘Chicago’ theory and sticky price monetary model (SPMM) attributed to the so-called Keynesian theory assuming that prices are sticky, at least, in the short run. Following Neely and Sarno (2002) let us assume that domestic policy makers decide to apply a restrictive monetary policy that leads to a contraction in domestic real money supply due to the price stickiness dominated in the economy in the short-run. This in turn creates an upward pressure on domestic interest rates to clear the money market. Increases in relative interest rates attract capital inflows to the domestic economy and consequently the domestic currency appreciates and the nominal exchange rate decreases by a greater extent than the decrease in equilibrium exchange rate. Consequently, a negative relationship occurs between the exchange rate and nominal interest differential.

Neely and Sarno (2002) express that a short-run equilibrium is achieved when the expected rate of depreciation is just equal to the interest rate differential, i.e., when the uncovered interest parity (UIP) condition holds. In the medium run, however, domestic prices begin to fall in response to the fall in the money supply leading to a rise in the real money supply in turn decreasing domestic interest rates. Therefore, the exchange rate depreciates slowly toward long-run PPP. Comparing FPMM and SPMM of exchange rate determination Frankel (1979) considers the 'Chicago' theory as a realistic assumption when variation in inflation differential is large such as witnessed in German hyperinflation of the 1920s, while Keynesian SPMM would be more realistic when variation in inflation differential is small. In line with these assumptions and following Cheung and Chinn (1998) we can write down the SPMM of exchange rate determination as in Eq. (6):

$$s_t = (m_t - m_t^*) - \alpha (y_t - y_t^*) - (1/\theta)(i_t - i_t^*) + [\beta + (1/\theta)](\pi - \pi^*) \qquad (6)$$

where π and π^* represent the domestic and foreign inflations, respectively. Eq. (6) differs from Eq. (5) in that the former assumes slow adjustment of goods prices at rate θ and instantaneous adjustment of asset prices thus yielding the overshooting characteristic, whereas the latter relies on the assumptions that the prices are perfectly flexible and that PPP holds continuously.[1]

III. Empirical Model

III. 1. Preliminary Data Specification

We now construct a model of exchange rate determination of the YTL/US$ for the Turkish economy. We consider data for the investigation period 1986M01-2007M08 using monthly frequency data. For the nominal exchange rate data (*s*), the spot new Turkish lira per US dollar, i.e. YTL/US$ exchange rate, is used. Money supply measures (*m*) and real income data (*y*) are represented by the M2 broad money supplies and 2000: 100 based industrial production indices, respectively. For the interest rate data (*i*), 12-

months time deposit rates are considered. Finally, inflation rates (p) are based on the consumer price indices with the base year 2000: 100. All the data take the form of seasonally unadjusted values in their natural logarithms except the 12-months interest rates and annualized inflation which are in their linear forms. The nominal exchange rate of YTL/US$, money supply and interest rate variables are taken from the electronic data delivery systems of the Central Bank of the Republic of Turkey for the Turkish data and FRB of St. Louis for the US data, while inflation and industrial production data for both the Turkish and the US economies are obtained from the OECD electronic statistical database.

Spurious regression problem analyzed by Granger and Newbold (1974) indicate that using nonstationary time series steadily diverging from long-run mean causes to unreliable correlations within the regression analysis leading to unbounded variance process. However, for the mean, variance and covariance of a time series to be constant over time, conditional probability distributions of the series must be invariant with respect to the time, and if only so can the conventional procedures of OLS resgressions be applied using a stationary process for the variables. 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 for any X variable as follows:

$$\Delta X_t = \alpha + \beta t + (\rho\text{-}1)X_{t\text{-}1} + \sum_{i=1}^{k} \Phi \Delta X_{t\text{-}i} + \varepsilon_t \qquad (7)$$

of which the null hypothesis is the presence of a unit root (ρ=1) against the alternative (trend)stationary hypothesis. For X_t to be stationary, (ρ-1) should be negative and significantly different from zero. Moreover, while the assumption that X_t follows an autoregressive (AR) process may seem restrictive, Said and Dickey (1984) demonstrate that the ADF test is asymptotically valid in the presence of a moving average (MA) component, provided that sufficient lagged difference terms are included in the test regression. The estimated ADF statistics are compared with the simulated MacKinnon (1991, 1996) critical values, which employ a set of simulations to derive asymptotic results and to simulate critical values for arbitrary sample sizes. For the case of stationarity, we expect that these statistics must be larger than the critical values in absolute value and have a minus sign.

However, 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 the KPSS tests. The KPSS tests are designed to test the null hypothesis of stationarity against the unit root alternative. Kwiatkowski et al. (1992) argue that such tests should complement the ADF-type tests to test the non-stationarity of the variables. The KPSS test statistic is computed based on the residuals of the regression of any Y_t series onto the exogenous variable ζ_t which follows a random walk process (Mahadeva and Robinson, 2004):[2]

$$Y_t = \zeta_t + \varepsilon_t \tag{8}$$

using an auxiliary equation for ζ_t:

$$\zeta_t = \zeta_{t-1} + \upsilon_t, \tag{9}$$

where ε_t represents a stationary process and υ_t has been subject to an expected value with a normally distributed zero-mean and constant variance process. For the KPSS test H_0 and H_1 hypothesis can be indicated as follows:

$$H_0: \sigma_\upsilon^2 = 0 \tag{10}$$

$$H_1: \sigma_\upsilon^2 > 0 \tag{11}$$

Kwiatkowski et al. (1992) propose the following test statistic for the unit root test:

$$\text{KPSS} = T^{2} \sum_{t=1}^{T} S_t^2 \,/\, \sigma^2 \tag{12}$$

where:

$$S_t = \sum_{i=1}^{t} \hat{u}_t,\ \hat{u}_t = Y_t - \hat{Y}_t \text{ and } \sigma^2 = \lim_{t\to\propto} T^{-1}\text{Var} \sum_{t=1}^{T} \varepsilon_t \tag{13}$$

Following these theoretical issues, unit root test results are reported below:

Tab. 1. Unit Root Tests

Variable	τ_c^{ADF}	τ_t^{ADF}	τ_c^{KPSS}	τ_t^{KPSS}
s	-2.08	1.01	2.06	0.38
Δs	-9.70^{*}	-10.06^{*}	0.41^{*}	0.13^{*}
m	-1.03	0.92	2.09	0.36
Δm	-9.54^{*}	-9.79^{*}	0.45^{*}	0.15^{*}
y	-2.12	-2.45	0.75	0.21
Δy	-17.76^{*}	-17.72^{*}	0.14^{*}	0.13^{*}
i	-1.41	-1.94	0.74	0.45
Δi	-12.04^{*}	-12.08^{*}	0.21^{*}	0.05^{*}
π	-0.32	-1.18	0.84	0.46
$\Delta\pi$	-5.84^{*}	-7.83^{*}	0.20^{*}	0.04^{*}
5% cri. val.	-2.87	-3.43	0.46	0.15

Above, τ_c and τ_t are the test statistics with allowance for only constant and constant&trend tems in the unit root tests, respectively. 'Δ' represents the first difference operator. An asterisk denotes that the data are of stationary form. The results of ADF unit root tests reveal that the null hypothesis that there is a unit root cannot be rejected for all the variables in the level form, but inversely, for the first differences the stationary alternative hypothesis can be accepted. Likewise, the KPSS tests under the null hypothesis of stationarity indicate that all the variables are difference stationary.

III. 2. Multivariate Tests for Co-integration

In order to test for a long-run stationary relationship derived from the variable space expressed above, we apply to the multivariate co-integration methodology proposed by Johansen (1988) and Johansen and Juselius (1990). This methodology constructs an error correction mechanism among the same order integrated variables enabling that a stationary combination of the variables do not drift apart without bound even though all have been individually subject to a non-stationary I(d) process, therefore ruling out the possibility that estimated relationships tend to be spurious. Besides, this technique is superior to the regression-based techniques, e.g. Engle and Granger (1987) two-step methodology, for it enables researchers to capture all the possible stationary relationships lying within the long-run variable space. Following MacDonald and Taylor (1993), however, using ordinary least squares to estimate a co-integrating relationship for an n-

dimensioned vector does not clarify whether one is dealing with a unique co-integrating vector or a linear combination of the potential (n-1) distinct co-integrating vectors that may be lying within 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 (VAR) involving up to k-lags of z_t:

$$z_t = \Pi_1 z_{t-1} + \Pi_2 z_{t-2} + \dots + \Pi_k z_{t-k} + \varepsilon_t \tag{14}$$

where ε_t follows an i.i.d. process N(0, σ^2) and z is (nx1) and the Π_i an (nxn) matrix of parameters. Eq. 14 can be rewritten leading to a vector error correction (VEC) model of the form:

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + \Gamma_2 \Delta z_{t-2} + \dots + \Gamma_{k-1} \Delta z_{t-k+1} + \Pi z_{t-k} + \varepsilon_t \tag{15}$$

where:

$$\Gamma_i = -I + \Pi_1 + \dots + \Pi_i \;\; (i = 1,\dots,k\text{-}1) \text{ and } \Pi = I - \Pi_1 - \Pi_2 - \dots - \Pi_k \tag{16}$$

Eq. 15 can be arrived by subtracting z_{t-1} from both sides of Eq. 14 and collecting terms on z_{t-1} and then adding $-(\Pi_1 - 1)X_{t-1} + (\Pi_1 - 1)X_{t-1}$. Repeating this process and collecting of terms would yield Eq. 15 (Hafer and Kutan, 1994). 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 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 Eq. 15 represents up to (n-1) co-integrating relations in the multivariate model which ensure that z_t converge to their long-run steady-state solutions.

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 that 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 exist r co-integrating vectors in $\beta' z_{t-k} \sim$ I(0) and (n-r) common stochastic trends when Π has reduced rank, i.e., $0 < r \leq (n\text{-}1)$. That is, first r columns of β are the linearly independent combinations of the endogenous variables settled in vector z_t, which represents stationary relationships. Whereas, the latter (n-r) columns constitute the non-stationary vectors of I(1) common trends, which require also 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. Further, this method does not suffer from problems associated with normalisation (Johansen, 1995). We thus first determined the lag length of our unrestricted VAR model using Schwarz (SC) information criterion. Considering the maximum lag length 12 for the unrestricted VAR model of monthly frequency data, the SC information criterion suggests using 2 lags to construct the unrestricted VAR model.[3] As a next step, we estimate the long run co-integrating relationships between the variables 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 $r = 0,1, \ldots ,n\text{-}1$ where n is the number of endogenous variables. For the co-integration test, intercept and trend factor have been restricted into the long run variable space following the *so-called* Pantula principle. Johansen (1992) and Harris (1995) suggest the need to test the joint hypothesis of both the rank order and the deterministic components. They propose to identify the sub-hypotheses, which give different limit distributions, and construct a test statistic and a critical region for each of these sub-hypotheses. The critical region for the test of the original null hypothesis is then the intersection of the critical regions constructed for each of the sub-hypotheses or, in other words, the hypothesis in question is only rejected if all sub-hypotheses are rejected. The test procedure begins by considering the most restrictive model and at each stage compare the LR test statistics to its critical value and only stop the first time the null hypothesis is not rejected. However, a critical point to be considered here may be that assuming quadratic deterministic trends allowing for linear trends in the short run VEC model may be economically difficult to justify especially if the variables are entered in log-linear form, since this would imply an implausible ever-increasing or decreasing rate of change. In line with these model specification issues, we restrict a long-run deterministic trend in the co-integrating space, but no deterministic trend is allowed for the short-run dynamics:

Tab. 2. Co-integration Test Results

Null hypothesis					
	r=0	r≤1	r≤2	r≤3	r≤4
Eigen value	0.17	0.10	0.06	0.04	0.03
λ trace	108.9*	60.77	33.07	17.29	7.33
5% cri. val.	88.80	63.88	42.92	25.87	12.52
Prob.	0.00	0.09	0.33	0.39	0.31
λ max	48.09*	27.69	15.78	9.96	7.33
5% cri. val.	38.33	32.12	25.82	19.39	12.52
Prob.	0.00	0.16	0.56	0.62	0.31
* denotes rejection of the hypothesis at the 0.05 level.					
Standardized eigenvectors					
s_t	$(m_t - m_t^*)$	$(y_t - y_t^*)$	$(i_t - i_t^*)$	$(\pi_t - \pi_t^*)$	*trend*
1.000	-0.652	3.449	-0.034	0.010	-0.013
-3.907	1.000	-21.82	0.299	-0.283	0.109
0.107	-0.033	1.000	-0.001	-0.001	-0.004
-905.2	1695.5	1241.5	1.000	-1.146	-35.53
-221.7	264.9	-327.4	-3.359	1.000	4.476
Weak exogeneity test statistics					
	s_t	$(m_t - m_t^*)$	$(y_t - y_t^*)$	$(i_t - i_t^*)$	$(\pi_t - \pi_t^*)$
LR test $\chi^2(1)$	10.82	8.903	0.536	1.966	0.247
Probs.	(0.000)	(0.003)	(0.464)	(0.161)	(0.619)
Multivariate statistics for testing stationarity					
	s_t	$(m_t - m_t^*)$	$(y_t - y_t^*)$	$(i_t - i_t^*)$	$(\pi_t - \pi_t^*)$
LR test $\chi^2(4)$	34.10	36.68	36.03	10.25	10.39
Probs.	(0.000)	(0.000)	(0.000)	(0.049)	(0.042)

In Tab. 2, we find that both rank statistics indicate that a unique co-intagrating vector lies in the long-run variable space which represents the existence of a stationary relationship between the variables of interest. Normalizing the first vector with the largest eigenvalue on the nominal exchange rate yields a theoretically plausible co-integrating equation:

$$s_t = 0.65(m_t - m_t^*) - 3.45(y_t - y_t^*) + 0.03(i_t - i_t^*) - 0.01(\pi_t - \pi_t^*) + 0.01\,trend \quad (17)$$

t-stat. (2.419) (-2.673) (4.742) (0.168) (-1.115)

Results from Eq. 17 give support to the monetary model of exchange rate determination constructed above. The relative money supply has a positive and relative real income has a negative significant long-run relationship with nominal exchange rate. Following Karfakis (2003), therefore, a positive monetary shock would raise permanently the level of exchange rate and an increase in the relative income would lead to an appreciation of the domestic currency against the US$ in the long-run. Karfakis attributes such an estimation result to that any policy which boosts economic growth would mean a strong domestic currency. Besides, interest differential variable has a positive and significant sign as *a priori* hypothesized. However, no significant effect of inflation differentials on the nominal exchange rate has been found. In the long-run variable space, we cannot reject the weak exogeneity of relative income, relative interest and inflation differentials and accept the endogeneity of exchange rate and relative money supply in the long-run co-integrating variable space. We can easily notice from Tab. 2 that non-stationary time-series characteristics of the variables are verified by the multivariate statistics for testing stationarity derived from the co-integration analysis in the sense that no variable alone can represent a stationary relationship in the co-integrating vector. We must finally note that we obtain one significant co-integrating vector as reported in Eq. 18 carrying highly similar characteristics to the findings obtained above when the trend factor has been excluded from the co-integrating space:

$$s_t = 0.95(m_t - m_t^*) - 2.25(y_t - y_t^*) + 0.032(i_t - i_t^*) - 0.01(\pi_t - \pi_t^*) \qquad (18)$$
$$t\text{-stat.} \quad (35.70) \qquad (-2.582) \qquad (4.759) \qquad (0.101)$$

IV. Concluding Remarks

Modeling determinants of exchange rates using economic fundamentals produces significant knowledge of monetary equilibrium combining some other contemporaneous monetary theories revealing equilibrium conditions for goods and assets markets. Such a phenomenon has been of special importance especially for developig countries since policy makers tend to canalize use of exchange rates to gain an *ex-ante* designed macroeconomic growth performance as well as to break the inertial nature of prices dominated in the economy in fighting inflation. In this paper, we try to investigate the exchange rate determination mechanism for the Turkish economy. Our empirical findings employing multivariate Johansen-Juselius type co-integrating approach for the 1986M01-2007M08 period of monthly frequency data indicate that YTL/US$ nominal exchange rate is co-integrated with the fundamentals suggested by economics theory. Complementary papers should be elaborately constructed to investigate equilibrium conditions in goods and assets markets separately and assess the out-of-sample forecasting performances of the exchange rate models.

Endnotes

[1] For a more comprehensive investigation of exchange rate determination models, see Sarno and Taylor (2003).

[2] Any deterministic linear trend can be included into Eq. (8) to test trend-stationarity.

[3] We also estimated sequential modified LR statistics and Akaike information criterion for the model. The former statistics proposed three co-integrating vectors lying in the long-run variable space and the latter statistics proposed two co-integrating vectors. Since we theoretically attributed one steady-state economic relationship to the long-run variable space, we chose the lag length proposed by the SC information criterion.

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