



EXCHANGE RATE DETERMINATION OF TL/US\$:A CO-INTEGRATION APPROACH

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Abstract

In our paper, we investigate exchange rate determination mechanism of TL/US\$ for the 1987Q1-2006Q4 period using quarterly observations. Following a large literature review we first highlight various approaches explaining monetary model exchange rate determination based on economic fundamentals and then construct an empirical model revealing both long-run stationary relationships and short-run dynamic adjustment processes of the nominal exchange rate for the Turkish economy. Our findings employing multivariate Johansen-Juselius type co-integrating approach indicate that nominal exchange rate is co-integrated with the fundamentals suggested by economics theory. Besides, short-run deviations from the fundamental-based equilibrium course of the nominal exchange rate have permanent effects on the long-run equilibrium exchange rate and so have been stemmed from the existence of some form of hysteresis effects dominated in the nominal exchange rate.

Keywords: Exchange Rates; Sticky Price Monetary Model; Flexible Price Monetary; Economic Fundamentals; Random Walk; Co-integration; Hysteresis; Turkish Economy

Jel Classification F31; F41; F47.

Özet

Çalışmamızda, TL/US\$ döviz kuru belirlenme mekanizması 1987Q1-2006Q4 döneminde üçer aylık veriler kullanılarak incelenmektedir. Geniş bir yazın taramasından yararlanılmak suretiyle öncelikle iktisadi temeller dahilinde oluşturulmuş döviz kuru belirlenme mekanizmasına yönelik çeşitli yaklaşımlar aydınlatılmış ve daha sonra parasal döviz kurundan kaynaklanan hem uzun dönem durağan ilişkileri hem de kısa dönemli devinimsel uyum süreçlerini Türkiye ekonomisi koşullarında ortaya koyan uygulamalı bir model oluşturulmuştur. Çok değişkenli Johansen-Juselius eş-bütünleşim yaklaşımı kullanılarak elde ettiğimiz sonuçlar parasal döviz kurunun iktisadi kuramı tarafından önerilen temeller ile eş-bütünleşik olduğunu göstermiştir. Ayrıca, parasal döviz kurunun iktisadi temeller dahilindeki denge yolundan kısa dönemli sapmaları uzun dönem denge döviz kuru üzerinde kalıcı etkiler göstermekte ve bu nedenle parasal döviz kuru içerisinde yerleşik bulunan histeresis türü etkilerin varlığı altında ortaya çıkmaktadır.

Anahtar Kelimeler: Döviz kurları; Katı Fiyat Parasal Modeli; Esnek Fiyat Parasal Modeli; İktisadi Temeller; Rassal Yürüyüş; Eş-bütünleşim; Histeresis; Türkiye Ekonomisi

Jel Sınıflaması: F31; F41; F47.

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I. INTRODUCTION

Determination of exchange rates has been of great concern for policy makers in today's liberalized and integrated world economies. Examining the course of exchange rates would help researchers conduct empirical investigations for testing the coherence of international macroeconomic theories such as purchasing power parity (PPP) and uncovered interest parity (UIP) as well as theories explaining the determination of exchange rates assuming open economy conditions. Such researches would reveal the extent to which discretionary economic policies can succeed in attaining the *ex-ante* policy targets and provide knowledge of forecasting performances of different modeling approaches following the developments in estimation techniques as for the future courses of economic variables. Especially, for a small and open developing country such as Turkey policy design process by the policy makers should be inclusive of the stylized facts based on these researches. Kesriyeli (1994), Metin (1994), Telatar and Kazdagli (1998), Bahmani-Oskooee and Kara (2000), Gokcan and Ozmen (2001), Dulger and Cin (2002), Civcir (2003a), Civcir (2003b), Civcir (2003c), Yazgan (2003), Erlat (2003), Ozdemir (2004) and a recent paper by Saatcioglu et al. (2007) give some empirical findings upon these issues of interest for the Turkish economy.

Of all these contemporaneous theoretical developments, a vast literature has been attributed to modeling the behavior of exchange rates so as to see whether monetary fundamentals are able to explain long-run course and short-run dynamics of exchange rates. What is of considerable interest in the economics literature is also to examine how well the out-of-sample forecasts fit to the actual data when assessing various estimation methods for forecasting purposes. Following the seminal paper by Meese and Rogoff (1983) indicating that fundamental based structural models of exchange rate do not beat the performance of the *naïve* random walk models in out-of-sample forecasts, there has been an extensive controversy upon these issues of interest. Researchers tend to explore whether the models based on structural relations or driven by *naïve*-random walks or considering more recent multivariate co-integration techniques, both assuming non-stationarity of data in the level form and preserving long-run knowledge of economic relations, must be of special interest and to the extent that they produce more accurate estimates models have been accepted to be superior when compared with the others.



However, some recent literature based on monetary model exchange rate determination yield estimation results supporting fundamental based models in the construction of exchange rates and give evidence in favor of out-of-sample forecasting performances for the fundamental based models against those of the random walk models particularly over relatively long horizons. In this sense, MacDonald and Taylor (1993) give some support to the flexible price monetary model (FPMM) forecasts of Deutsche mark-US dollar exchange rate in a dynamic error correction framework for the 1976-1990 period that are superior to those generated by a random walk forecasting model. McNown and Wallace (1994) search for co-integrating relationships leading to the monetary model of exchange rate determination for Israel, Chile and Argentina experiencing rapid monetary inflations. Their findings using data for the post-1973 period till the late-1980s not only support long-run relations among the variables of the monetary model but are also sensitive to the model specification as for the appropriate signs of the model in the sense that co-integrating vectors with plausible estimates are only obtained for Chile and Argentina but not for Israel. Likewise, Moosa (2000) examining the 1919-1923 German hyperinflation period gives evidence to the monetary model of exchange rate determination. Mark (1995) employing US dollar prices of the Canadian dollar, the Deutsche mark, the Swiss franc and the Japanese yen finds that long-horizon changes in the logarithm of spot exchange rates are predictable. He gives evidence to that the out-of-sample point predictions from fundamental based models generally out-perform the driftless random walk model at the longer horizons. Kilian (1999) and Berkowitz and Giorgianni (2001) present a criticism to Mark's (1995) methodology dealing with the data generating process used for predictability. Chinn and Meese (1995) support Mark's conclusions to a great extent and find that fundamental-based error correction models out-perform the random walk model for long-term prediction horizons. Neely and Sarno (2002) give a brief summary of the seminal papers of Meese and Rogoff (1983) and Mark (1995) upon exchange rate determination mechanism considered in the economics literature.

MacDonald and Marsh (1997) also indicate that fully dynamic out-of-sample forecasts from simultaneous equations models incorporating meaningful long-run equilibrium and short-run dynamic relationships are capable of significantly out-performing those of a random walk model considering US, Japan, UK and German data for the 1974-1990 period. Cheung and Chinn (1998) apply to a methodology using some consistency tests of evaluating



exchange rate forecast rationality for Japanese yen, German Mark and Canadian dollar exchange rates against the US dollar for 1983-1993 and 1987-1993 periods, for which consistency requires that the forecast and the actual series (i) be the same order of integration, (ii) be cointegrated and (iii) yield co-integrating vector consistent with long-run unitary elasticity of expectations. They indicate that the first requirement generally holds, however co-integration fails to hold the longer the horizon. Of the co-integrating pairs, besides, the third requirement is not generally rejected.

Following recent developments in the methodology of estimation techniques, researchers tend to use panel estimation techniques and so examine whether results from panel based studies are able to yield more powerful results in favor of fundamental based models in the long run. In this sense, Groen (2000) employs data for 14 bilateral exchange rates and monetary fundamentals with respect to either the US dollar or Deutsche mark for the 1973-1994 period. However co-integration tests on the time series of individual countries lack of giving evidence for the monetary model, he finds that panel based testing both produces more powerful results supporting monetary model and indicates co-integrating relation between exchange rates and monetary fundamentals. Mark and Sul (2001) examine the long-run relationship between nominal exchange rates and monetary fundamentals in a panel of 19 countries for the 1973-1997 period. They estimate that co-integration between exchange rates and long run fundamentals predicted by economics theory is generally approved by the data and that panel based forecasts indicate significant predictive power of monetary fundamentals for future exchange rate movements. Using the data set of Mark and Sul (2001), Rapach and Wohar (2004) test the long run monetary model of exchange rate determination as well. They first support the evidence that country-by-country estimates of co-integrating coefficients diverge widely from the values predicted by the monetary model and give little evidence for the co-integration between nominal exchange rates and monetary fundamentals in the floating period. But when they apply to the panel estimates of co-integrating coefficients they obtain supportive results in line with the monetary model as for the signs and magnitudes. However, when they analyze the cross-country homogeneity restrictions in the panel estimation procedure they reveal that such assumptions are not supported by the data. Rapach and Wohar (2002) emphasize that results in Groen (2000) and Mark and Sul (2001) should be re-examined for robustness to various subpanels and to formally test for heterogeneity across panel members.



Besides, some recent papers find evidence for the fundamental based models. Abhyankar et al. (2005) using data for the US, Canada, Japan and the UK for the modern floating exchange rate period lend some support to the predictive ability of the exchange rate monetary fundamentals model. Their results indicate that the gain from using the information in fundamentals in order to predict the exchange rate out-of-sample is found substantial. Karfakis (2006) using US and euro area observations for the 1999-2004 period reveals that the exchange rate is co-integrated with money and income differentials and that estimated error correction model out-performs the random walk forecasts. Also Nwafor (2006) supports flexible price monetary model by estimating a long-run equilibrium relationship between Nigerian naira and the US dollar.

In our paper, our aim is to examine the empirical validity of monetary model of exchange rate determination for the Turkish economy and to compare the out-of-sample forecasting performances of the results with those of a *naïve* random walk model. For this purpose, the outline of the paper is as follows. We first highlight the construction of a simple flexible price monetary exchange rate model as well as some extensions examined in the economics literature. Then, an empirical model for the Turkish economy is constructed which also assesses the out-of sample forecasting performance against *naïve* random walk model. Finally, the last section summarizes results and concludes.

II. MODEL CONSTRUCTION

II.1. A Simple Monetary Model of Exchange Rate Determination

Following an excellent paper by Neely and Sarno (2002) we begin our analysis by examining the flexible price monetary model (FPMM) developed in the 1970s mainly by Frenkel (1976), Mussa (1976), Bilson (1978a) and Bilson (1978b). 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 the price level and exchange rate to the money supply changes. McNown and Wallace (1994) express that if the demand for money is stable, the 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 the 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:

$$m_t = p_t + \alpha y_t - \beta i_t \quad (1)$$

$$m_t^* = p_t^* + \alpha^* y_t^* - \beta^* i_t^* \quad (2)$$

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:

$$s_t = p_t - p_t^* \quad (3)$$

where s_t represents 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:¹

¹ Following Nwafor (2006), expectations can be introduced in Eq. 4. Since the nominal interest rate consists of real interest rate (r) and the expected inflation (π^e):

$$i_t = r_t + \pi_t^e$$

$$i_t^* = r_t^* + \pi_t^{e*}$$

and supposing that real interest rates are equalized in home and foreign countries:

$$i_t - i_t^* = \pi_t^e - \pi_t^{e*}$$

Thus FPMM could be re-arranged such as:

$$s_t = (m_t - m_t^*) - (\alpha y_t - \alpha^* y_t^*) + (\theta \pi_t^e - \theta \pi_t^{e*})$$

As for the sign of the relative expected inflation, since an increase in π_t^e decreases the demand for money and increases demand for domestic and foreign assets we expect that an increase in expected relative inflation would lead to a depreciation of domestic currency, that is, to a rise in s_t .



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

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.

II. 2. Extensions of the Model

Based on the the FPMM given in Eq. (5) some other versions of the model can be derived 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. Goldberg (2000) considers such a transmission mechanism by an example based on a one time increase in domestic money.

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 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 1920's, 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):

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

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.

MacDonald and Taylor (1993) indicate a possibility that information from the stationary long-run equilibrium, i.e. co-integrating, vector of nominal exchange rate conditioned upon monetary fundamentals may lead to reducing the change in the exchange rate when s_t tends to be above its equilibrium level in the long-run supporting overshooting hypothesis. Bahmani-Oskooee and Kara (2000) employing the Turkish data estimate that lagged error-correction term (of the value 1.56) produced in autoregressive distributed lag (ARDL) co-integrating model suggested by Pesaran and Shin (1999) and Pesaran, Shin

and Smith (2001) does not carry a negative but a positive significant coefficient. They interpret this result such that since the relative money supply elasticity of nominal exchange rate is found greater than unity, the Turkish lira would tend to be overshoot in the long-run. Goldberg (2000) also reveal the requirements for and indicate the possibility of that nominal exchange rate may be subject to undershooting as well as to overshooting its long-run value on impact in response to a change in money supply. We will come back below to these issues of interest recalling some form of hysteresis phenomenon dominated in exchange rate determination mechanism when applying to the Turkish data for empirical purposes.

Some other extensions of the model can be obtained through Dornbusch (1976) by assuming the effect of tradables and non-tradables indicated in Eq. (7):

$$s_t = (m_t - m_t^*) - \alpha(y - y_t^*) - (1/\theta)(i_t - i_t^*) + [\beta + (1/\theta)](\pi - \pi^*) + \kappa[(p^T - p^N) - (p^{T*} - p^{N*})] \quad (7)$$

where p^T and p^N are the relative price of tradables and non-tradables, respectively. Cheung and Chinn (1998) report that this model is motivated by the failure of the PPP to hold for broad price indices such as the consumer price index and GDP deflator. It makes an explicit recognition of tradable and non-tradable goods and posits the assumption that the PPP only holds for tradable goods. Besides, Hooper and Morton (1982) allow cumulated trade balances in affecting the nominal exchange rate such as Eq.(8):

$$s_t = (m_t - m_t^*) - \alpha(y - y_t^*) - (1/\theta)(i_t - i_t^*) + [\beta + (1/\theta)](\pi - \pi^*) - \eta tb + \eta^* tb^* \quad (8)$$

where tb represents the trade balance. However, in our paper our aim is to test whether the Turkish data support empirically SPMM of exchange rate determination in Eq. (6) above and we will leave further extensions of this model to the future researches.

III. EMPIRICAL MODEL

III. 1. Preliminary Data Specification

We now construct a model of exchange rate determination of the TL/US\$ for the Turkish economy. We consider data for the investigation period of 1987Q1-2006Q4 using



quarterly observations. All the data take the form of seasonally unadjusted values in their natural logarithms except the interest rates, which are in their linear forms, and are taken from the electronic data delivery system of the Central Bank of the Republic of Turkey (CBRT) for the domestic variables and from the FRB of St. Louis electronic data delivery system for the external variables. For the exchange rate data the spot Turkish lira per US dollar, i.e. TL/US\$ exchange rate, is used. Money supply measures are represented by the M2 broad money supplies, and real gross domestic product (GDP) data are used for real income variables. Short-term interest rates are considered, and for this purpose, the Treasury interest rate for domestic interest rates, which is the maximum rate of interest on the Treasury bills whose maturity are at most twelve months or less, and the one-year Treasury constant maturity rate for the US economy are used. Price measures are based on annual inflations calculated as the difference between GDP deflator and four period lagged value in their natural logarithms.

As a next step, we investigate the time series properties of the variables of interest. Spurious regression problem introduced by Yule (1926), and further analyzed by Granger and Newbold (1974), Phillips (1986) and Phillips and Durlauf (1986) indicate that using non-stationary 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 requires that conditional probability distributions of the series tend to be invariant with respect to the time, and if only so can the conventional procedures of OLS regressions be applied using a stationary process for the variables. Therefore, at first by using the augmented Dickey-Fuller (ADF) unit root tests of Dickey and Fuller (1979, 1981) under the null hypothesis for the presence of a unit root against the stationary alternative hypothesis, we check for the stationarity condition of our variables and compare the estimated ADF statistics 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 are larger than the MacKinnon critical values in absolute value and that they have a minus sign. The lags used for the ADF stationarity test are augmented up to a maximum of 10 lags and the choice of the optimum lag was decided on the basis of minimizing the Schwarz Information Criterion (SC). ‘*’ indicate the rejection of a unit root.

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 Kwiatkowski,

Phillips, Schmidt, and Shin (KPSS) unit root test of Kwiatkowski et al. (1992) is also conducted to examine the time-series characteristics of the variables. The KPSS test differs from the ADF unit root test in that the series considered is assumed to be stationary under the null in the KPSS test. Yavuz (2004) highlights the properties of the ADF type and KPSS tests and tries to compare them by using Turkish stock exchange data:

Table 1: Unit Root Tests

Variable	τ_C	τ_T	$Z(\tau_C)$	$Z(\tau_T)$
s_t	-2.04	0.53	1.21	0.22
Δs_t	-5.76*	-6.29*	0.42*	0.13*
$(m_t - m_t^*)$	-1.89	0.36	1.22	0.22
$\Delta(m_t - m_t^*)$	-5.03*	-5.46*	0.46*	0.15*
$(y_t - y_t^*)$	-1.92	-2.34	0.52	0.19
$\Delta(y_t - y_t^*)$	-2.92*	-2.90	0.11*	0.11*
$(i_t - i_t^*)$	-2.57	-2.79	0.54	0.28
$\Delta(i_t - i_t^*)$	-8.72*	-8.81*	0.26*	0.08*
$(\pi_t - \pi_t^*)$	-0.11	-1.42	0.72	0.25
$\Delta(\pi_t - \pi_t^*)$	-7.22*	-7.39*	0.08*	0.03*

Above, τ_C and τ_T are the test statistics with allowance for only constant and constant&trend terms in the unit root tests, respectively, and $Z(\tau_C)$ and $Z(\tau_T)$ are the relevant KPSS statistics. ‘ Δ ’ denotes the first difference operator. 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 of difference-stationary form.

III. 2. Some Methodological Issues

Recent developments in time series estimation techniques provide new insights for researchers in applying to macroeconomic models for policy purposes. Since the world economies have been occasionally subject to institutional and structural changes *à la* Lucas (1981) the traditional large-scaled simultaneous equation macroeconomic models are not able to detect most econometric long-run relationships and lead to controversies as for the stability of structural relationships between macroeconomic aggregates. Also the classic paper



by Nelson and Plosser (1982) reveal that many macroeconomic time series data have a stochastic trend plus a stationary component, that is, they are difference stationary processes, and following Enders (2004) we can state that numerous economic theories suggest the importance of distinguishing between temporary and permanent movements in a series. In this sense, economic theory assumes that at least some subsets of economic variables do not drift through time independently of each other and some combination of the variables in these subsets reverts to the mean of a stable stochastic process (Anderson et al., 1998).

Granger (1986) and Engle and Granger (1987) indicate that even though economic time series may be non-stationary in their level forms, there may exist some linear combination of these variables that converge to a long run relationship over time, which also requires that there must be Granger causality in at least one direction in an economic sense as one variable can help forecast the others. That is, if the series are individually stationary after differencing but a linear combination of their levels is stationary then the series are said to be co-integrated. In such a case, they cannot move too far away from each other in a theoretical sense (Dickey, Jansen and Thornton, 1991). Engle and Yoo (1987) using some Monte Carlo experiments estimate that forecasts taken from co-integrating relations will have a finite limiting forecast error variance leading to that they are tied together because the co-integrating relations must hold exactly in the long run (Hoffman and Rasche, 1996). However, as emphasized in Anderson et al. (1998), since the long-horizon forecast of the error correction term is always zero incorporating co-integrating relationships for forecasting purposes may reduce mean square forecast error only at short horizons.

Therefore, adopting an error-correction model to the data enables researchers to track both short- and long-run dynamics and provides long-run stability by the introduction of error correction term in order to adjust for departures from equilibrium (Strauss-Kahn, 1991). Otherwise, by analysing only the differences of economic time series all information about potential long-run relationships between the levels of economic variables would be lost (Hendry, 1986). Contemporaneous co-integration techniques, e.g. proposed by Johansen (1988) and Johansen and Juselius (1990) as a further development to co-integration methodology, enabling also to testing that more than one stationary long-run equilibrium relation can be lying in the long-run variable space, take account of the non-stationarity characteristics of the most economic aggregate time series. Whereas, employing conventional estimation techniques based on an OLS estimation would not possibly lead to a constant mean

and a finite variance and therefore diverge after a shock. In line with these developments in econometrics theory, contemporaneous economics theories make use of these estimation tools in constructing and testing the theories based on model specification issues conditioned upon econometrics.

III. 3. Johansen-Juselius Multivariate Co-integration Techniques

In order to test for a long-run stationary relationship derived from the variable space expressed above, we apply to the multivariate co-integration and vector error correction (VEC) techniques 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-dimensional 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 :

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

where ε_t follows an i.i.d. process $N(0, \sigma^2)$ and z is (nx1) and the Π_i an (nxn) matrix of parameters. Eq. 9 can be rewritten leading us to a vector error correction (VEC) model of the form:

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



where:

$$\Gamma_i = -I + \Pi_1 + \dots + \Pi_i \quad (i = 1, 2, \dots, k-1) \quad \text{and} \quad \Pi = I - \Pi_1 - \Pi_2 - \dots - \Pi_k \quad (11)$$

Eq. 10 can be arrived by subtracting z_{t-1} from both sides of Eq. 9 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. 10 (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. 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.

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 $(n \times n)$ 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 for us 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., $0 < r \leq (n-1)$. 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 or when the dynamics are unknown. 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 sequential modified LR statistics employing small sample modification, which compare the modified LR statistics to the 5% critical values starting from the maximum lag and decreasing the lag one at a time until first getting a rejection. Considering the maximum lag of 8 for the unrestricted VAR model of quarterly frequency data, in our case reduction of system is first rejected when applying to 5 lag orders by the sequential modified LR statistics. Besides, the minimized Akaike information criterion suggests using the five lag orders. However, Schwarz (SC) and Hannan-Quinn (HQ) statistics suggest one lag order to be considered.

On this point, we should specify that for the appropriate lag length to ensure that the residuals are Gaussian, i.e.. they do not suffer from autocorrelation, non-normality, etc., considering the presence of co-integrating relationships used in the next sub-section, Cheung and Lai (1993) find that Monte Carlo experience carried out using data generating processes (DGPs) suggests that tests of co-integration rank are relatively robust to over-parametrizing, while setting too small a value of lag length such as lag length one or two generally suggested by SC statistics also producing serial correlation problem severely distorts the size of the maximum likelihood tests (Cheung and Lai, 1993; Harris, 1995). Gonzalo (1994) also reveals that the cost of over-parametrizing by including more lags in the maximum likelihood based error correction model (ECM) is small in terms of efficiency lost, but this is not the case if the ECM is underparametrized.

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 relationships and trace for the null hypothesis of r co-integrating relations against the alternative of n co-integrating relations, for $r = 0, 1, \dots, n-1$ where n is the number of endogenous variables. Briefly to say, to test the null hypothesis that there are at most r cointegrating vectors and thus $k-r$ unit roots amounts to:

$$H_0: \lambda_i = 0, \quad i = r+1, \dots, n \quad (12)$$

where only the first r eigenvalues are non-zero. This restriction can be imposed for different values of r and then the log of the maximised likelihood function for the restricted model is



compared to the log of the maximised likelihood function of the unrestricted model and a standard LR test computed. Using the trace statistic we can test the null hypothesis:

$$\lambda_{\text{trace}} = -2 \log (Q) = -T \sum_{i=r+1}^n \log (1-\lambda_i) \quad \text{and} \quad r = 0, 1, 2, \dots, n-2, n-1 \quad (13)$$

where $Q = (\text{restricted maximised likelihood} / \text{unrestricted maximised likelihood})$, T is the sample size. Another test of the significance of the largest λ_i is the maximal-eigenvalue statistic:

$$\lambda_{\text{max}} = -T \log (1-\lambda_{r+1}) \quad \text{and} \quad r = 0, 1, 2, \dots, n-2, n-1 \quad (14)$$

which tests that there are r co-integration vectors against the alternative that $r+1$ exist as expressed above.

For the co-integration test, we restrict intercept and trend factor into our 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 is to move through from the most restrictive model and at each stage to compare the trace or max-eigen 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. Doornik et al. (1998) also indicate that restricting the trend factor into the co-integration space is preferable. In line with these model specification issues we assume that a long-run deterministic trend would be restricted in the co-integrating space, but no deterministic trend is allowed for the short-run dynamics.



In the next section we give the estimation results of the model including also a set of centered seasonal dummy variables included into the model, which sum to zero over a year as exogeneous variable. Johansen (1995) indicates that the linear term from the dummies in this way disappears and is taken over completely by the constant term, and only the seasonally varying means remain. For instance, the second quarter takes the value of 0.75 while the sum of the remaining three three quarters' dummies is -0.75.

III. 4. Estimation Results

In line with data and econometric model specification issues highlighted so far, the TL/US\$ exchange rate determination model is estimated using the multivariate co-integration methodology of the same order integrated variables. In Tab. 2 below which reports the results of Johansen co-integration test using max-eigen and trace tests based on critical values taken from Osterwald-Lenum (1992) are given the estimation results in which a constant and long-run deterministic trend are restricted but no deterministic trend is assumed for dynamic VEC model. So doing we leave the 2004Q1-2006Q4 period for out-of-sample forecasting purposes:



Table 2: Co-integration Test

Null hypothesis	r=0	r≤1	r≤2	r≤3	r≤4
Eigenvalue	0.68	0.42	0.36	0.22	0.12
λ trace	144.41*	78.90*	46.97*	21.53	7.11
5% Critical Value	88.80	63.88	42.92	25.87	12.52
λ max	65.51*	31.93	25.44	14.42	7.11
5% Critical value	38.33	32.12	25.82	19.39	12.52

* denotes rejection of the hypothesis at the 0.05 level.

Unrestricted Co-integrating Coefficients

s_t	$(m_t - m_t^*)$	$(y_t - y_t^*)$	$(i_t - i_t^*)$	$(\pi_t - \pi_t^*)$	TREND
-14.47349	12.91783	-22.02630	6.342022	21.57344	0.210501
2.489526	1.714840	-2.774192	15.866650	-46.67813	-0.723257
-4.806262	3.443205	-42.51449	3.431268	-1.771258	0.180151
7.252148	-5.901042	10.32024	3.265402	12.92902	-0.213329
11.63431	-5.811610	6.008131	1.054633	-24.54527	-0.722579

Unrestricted Adjustment Coefficients (alpha)

D(s)	-0.017220	0.017531	-0.020810	-0.012986	-0.011404
D($m_t - m_t^*$)	-0.011632	-0.003169	-0.009759	0.010143	-0.000960
D($y_t - y_t^*$)	0.004803	0.002519	0.016990	0.001770	-0.000675
D($i_t - i_t^*$)	-0.099848	-0.046367	-0.042281	-0.033356	-0.029249
D($\pi_t - \pi_t^*$)	-0.028303	0.015654	0.003541	-0.006503	0.003069

1 Co-integrating Equation: Log likelihood 503.9703

Standard errors in () & t-statistics in []

s_t	$(m_t - m_t^*)$	$(y_t - y_t^*)$	$(i_t - i_t^*)$	$(\pi_t - \pi_t^*)$	TREND	C
1.000000	-0.892517	1.521838	-0.438182	-1.490549	-0.014544	-5.72
	(0.08829)	(0.34047)	(0.16971)	(0.45823)	(0.01319)	
	[-10.1095]	[4.46979]	[-2.58191]	[-3.25283]	[-1.10249]	

Adjustment coefficients

D(s_t)	D($m_t - m_t^*$)	D($y_t - y_t^*$)	D($i_t - i_t^*$)	D($\pi_t - \pi_t^*$)
0.249238	0.168361	-0.069522	1.445142	0.409638
(0.14257)	(0.07889)	(0.08008)	(0.45231)	(0.09749)
[1.748180]	[2.13414]	[-0.86813]	[3.19504]	[4.20187]

Multivariate Statistics for Testing Stationarity

	s_t	$(m_t - m_t^*)$	$(y_t - y_t^*)$	$(i_t - i_t^*)$	$(\pi_t - \pi_t^*)$
$\chi^2(4)$	45.96004	43.62834	37.96759	40.76771	38.49888

Homogeneity and Symmetry Restrictions on Co-integrating Coefficients

b(1,1)=1, b(1,2)=-1, b(1,3)=1, b(1,5)=-1, $\chi^2(3) = 4.27216$ $\chi^2(3)$ -table = 7.81473

b(1,1)=1, b(1,2)=-1, b(1,3)=1, b(1,4)=-1, b(1,5)=-1, $\chi^2(4) = 9.41221$ $\chi^2(4)$ -table = 9.48773

Rewriting the exchange rate equation under the assumption of r=1:

$$s_t = 0.89(m_t - m_t^*) - 1.52(y_t - y_t^*) + 0.44(i_t - i_t^*) + 1.49(\pi_t - \pi_t^*) + 0.01TREND + 5.72 \quad (15)$$

From the Tab. 2, trace test indicates 3 and max-eigen test 1 potential co-integrating vectors lying in the long-run variable space. In this situation some researchers may tend to work with more than one co-integrating vectors such as Dulger and Cin (2002), but this would require the identification of the each vector to which different economic interpretations can be attributed. Besides, following Harris (1995) some other identification issues to obtain independent vectors from each other would be required, which are able to yield independent coefficient and t-statistics. Thus in our paper we choose to work with the common vector suggested by both LR statistics with the largest eigenvalue. However, the co-integrating vector produced in the third row of unrestricted co-integrating coefficients has similar long run characteristics to the first vector as for their signs except the relative inflation variable.

When we normalize the first vector on the nominal exchange rate in Eq. 15 we find a theoretically plausible co-integrating equation. The relative money supply has a positive and relative real income has a negative significant long-run relationship with nominal exchange rate. These findings are common for all the models explaining monetary exchange rate determination. Besides, inflation differentials lead to a depreciation of the domestic currency as expected. The interest differential variable has a positive and significant sign recalling the FPMM of exchange rate determination. Thus, no support is given to the assumption that relative interest differential in favor of domestic interest rates leads to an appreciation of the domestic currency. These results are in line with the findings of Bahmani-Oskooee and Kara (2000) but contradict Dulger and Cin (2002). Goldberg (2000) relates the positive sign of interest rate in the monetary model also to the under-or-overshooting of the exchange rate under perfect or imperfect capital mobility.

An interesting finding in Tab. 2 is that the adjustment coefficient of nominal exchange rate carries a positive rather than a negative sign in a statistically significant way. Bahmani-Oskooee and Kara (2000) find a positive lagged error correction term of the monetary exchange rate equation produced by ARDL co-integrating model used and interpret this as exchange rate staying above its long-run value or overshooting itself in the long-run. They relate this result to the finding that relative money supply elasticity which has the value of 1.56 is greater than unity indicating that 1% increase in Turkish relative money supply would cause a long-run depreciation of the lira by 1.56% supporting the overshooting hypothesis. In our paper, however, nothing is related to under-or-overshooting of nominal exchange rate, but what is of the stylized fact for the Turkish economy and also giving support to Bahmani-



Oskooee and Kara (2000) is that short-run deviations from the fundamental-based equilibrium course of the nominal exchange rate have permanent effects on the long-run equilibrium exchange rate and so these deviations have been stemmed from the existence of some form of hysteresis effects dominated in the nominal exchange rate. The other adjustment coefficients estimated reveal that all the variables except the real income differential which is found weakly exogenous have an endogenous characteristic in our co-integrating system and these lead us to the feedback effects onto these variables requiring to construct dynamic vector error correction models adjusting to long-run stationary equilibrium relationship.

The hypothesis $sp(H_5) \subset sp(\beta)$ presented by Johansen and Juselius (1992) assumes r_1 of the r co-integrating relations known as specified by the matrix H_5 . The remaining r_2 relations are chosen without restrictions. This type of hypothesis for $r=1$ differs from widely used testing procedure of the Dickey-Fuller type of univariate testing in the sense that the former as a χ^2 test hypothesizes stationarity as the null hypothesis, whereas the usual latter univariate tests formulate the null of nonstationarity. Our former estimation results from the ADF and KPSS unit root testing procedures above which may yield incorrect inferences when the data generating process has shifted by the time are verified by the Johansen-Juselius type of testing stationarity in the sense that no variable alone can represent a stationary relationship in the co-integrating vector.

Finally, we accept the null hypothesis of homogeneity and symmetry restrictions applied in Tab. 2. If we explicitly write down these equations:

$$s = (m-m^*)-(y-y^*)+0.68(i-i^*)+(\pi -\pi^*)-0.00223TREND+5.16 \quad (16)$$

producing $\chi^2(3) = 4.27216$ against $\chi^2(3)$ -table = 7.81473, and:

$$s = (m-m^*)-(y-y^*)+(i-i^*)+(\pi -\pi^*)-0.00416TREND+4.97 \quad (17)$$

producing $\chi^2(4) = 9.41221$ against $\chi^2(4)$ -table = 9.48773. We must note that no VEC residual serial correlation problem occurs using LM-statistics $LM(1)=20.56682$ (0.7165) and $LM(4)=25.66830$ (0.4255) under the null of no serial correlation of the 1st or 4th order using prob. values in parantheses from χ^2 with 25 d.o.f. Under the null of no heteroskedasticity or no misspecification VEC residual heteroskedasticity test gives that $\chi^2(825)=833.1149$ using

the prob value of 0.4147. However, under the null hypothesis of multivariate normal residuals VEC residual normality test produces excess kurtosis and non-normality using skewness $\chi^2(5)=0.95$ (0.97), kurtosis $\chi^2(5)=57.60$ (0.00) and JB $\chi^2(10)=58.55$ (0.00) assuming Cholesky orthogonalization of Lütkepohl (1991), indicating significant outliers in the model. We find non-normality for residuals, but no problem in our model through Gonzalo (1994).

Having established the long run co-integrating equilibrium model, we now estimate the parsimonious VEC model whitening the error structure. We include two variables which have probs. above 5% due to multicollinearity problems:

Table 3: Parsimonious VEC Model

White Heteroskedasticity-Consistent Standard Errors & Covariance				
Variable	Coefficient	Std. Error	t-Statistic	Prob.
C	-0.199561	0.080716	-2.472384	0.0176
EC(-1)	0.157609	0.058500	2.694159	0.0101
D(s)(-1)	0.938654	0.297098	3.159404	0.0029
D(s)(-3)	0.339831	0.212954	1.595791	0.1180
D(s)(-5)	0.395634	0.173019	2.286650	0.0273
D(m-m [*])(-1)	-0.404985	0.269477	-1.502859	0.1404
D(m-m [*])(-2)	0.949123	0.304915	3.112750	0.0033
D(m-m [*])(-5)	0.523546	0.220041	2.379314	0.0220
D(y-y [*])(-1)	0.275938	0.118211	2.334281	0.0244
D(y-y [*])(-3)	0.287498	0.112441	2.556868	0.0143
D(i-i [*])(-2)	0.193941	0.067023	2.893673	0.0060
D(i-i [*])(-4)	0.139550	0.067001	2.082799	0.0434
D(π - π [*])(-1)	-0.969412	0.329385	-2.943094	0.0053
D(π - π [*])(-3)	-0.371398	0.158238	-2.347090	0.0237
D(π - π [*])(-5)	-0.517268	0.226670	-2.282035	0.0276

Parsimonious VEC model produces diagnostics Adj.R²=0.44, F=4.15, DW=2.14, LM(1)=0.68 (0.41), LM(4)=0.44 (0.78), Nor=5.12(0.08), White=1.33(0.23), where Adj.R² is the adjusted squared multiple correlation coefficient, LM the Breusch-Godfrey Serial Correlation LM Test, White the White's Heteroskedasticity Test, and Nor the Jarque-Bera Normality Test. Probs. are given in parantheses. The error correction coefficient carrying the long-run knowledge of monetary exchange rate determination from the co-integrating equation has a value in line with the adjustment coefficient given in Tab. 2. The parsimonious error correction model has good diagnostics as well.

III. 5. Out-of-Sample Predictability

We now try to compare the out-of-sample forecasting performances from the benchmark random walk (RW) and random walk with drift (RW^t) models, for which following Enders (2004) the latter is governed by two nonstationary process, i.e. a linear deterministic trend and a stochastic trend, against our fundamental based monetary exchange rate error correction model (FBMM). The out-of sample forecasts are evaluated for the 2004Q1-2006Q4 period. We compute fully dynamic predictions in this period from one to eleven forecast periods of the quarterly observations so that previously forecasted values are used in forming the forecasts of the current period. Thus, such a forecasting methodology will differ from static forecasts which calculate a sequence of one-step forecasts using the actual rather than forecasted values in estimation process. Following Mark(1995) and Neely and Sarno (2002) we will use the Theil's U statistics, that is, the ratio of root mean squared errors (RMSEs) from two competing models considering monetary versus random walk with/without drift models. A ratio less than one would give support to the superior forecasting performance of the fundamental based model, while a ratio greater than one would support the forecasting ability of RW model when compared with the fundamental based monetary model. The results can be seen in Tab. 4 and Tab. 5 below:

TABLE 4: FBMM / RW OUT-OF-SAMPLE FORECASTING COMPARISONS

Forecast horizon	1	2	3	4	5	6	7	8	9	10	11
FBMM / RW	0.77	0.63	0.57	0.65	0.68	0.67	0.72	0.82	0.91	1.04	1.22

TABLE 5: FBMM / RW^t OUT-OF-SAMPLE FORECASTING COMPARISONS

Forecast horizon	1	2	3	4	5	6	7	8	9	10	11
FBMM / RW ^t	0.69	0.55	0.50	0.56	0.59	0.58	0.62	0.70	0.78	0.87	1.02

Tab. 4 and Tab. 5 reveal that fundamental based monetary model out-performs the benchmark random walk models for the forecast horizons of one to nine or one-to ten quarters and then random walk model has superiority in forecasting against the fundamental based model.



These results indicate that assuming the knowledge from co-integrating relationship tends to improve the success of forecasting inside a 2.5 years forecasting horizon.

IV. CONCLUDING REMARKS

Determination 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. Especially in today's liberalized and integrated world economies, policy design process should be inclusive of the stylized facts based on such researches and policies should be constructed in line with the requirements extracted through testing the economics theory.

In our paper, we try to investigate the exchange rate determination mechanism for the Turkish economy. Our empirical findings following a large literature review explaining monetary model exchange rate determination and employing multivariate Johansen-Juselius type co-integrating approach indicate that TL/US\$ nominal exchange rate is co-integrated with the fundamentals suggested by economics theory. We find that short-run deviations from the fundamental-based equilibrium course of the nominal exchange rate have permanent effects on the long-run equilibrium exchange rate and so have been stemmed from the existence of some form of hysteresis effects dominated in the nominal exchange rate. Results obtained also indicate that fundamental based monetary model out-performs the benchmark random walk with/without drift models for the forecast horizons between one to nine and one to ten quarters thus leading us to the conclusion that knowledge from co-integrating relationship tends to improve the success of forecasting performance for the exchange rate determination model.

Complementary papers should be elaborately constructed to investigate equilibrium conditions in goods and assets markets separately and such future papers will help the researchers confirm whether estimation results in this paper are in fact of the stylized facts for the Turkish economy.

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