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The Monetary Model of the Exchange Rate under High Inflation

The Case of the Turkish Lira/US Dollar

Irfan CIVCIR*

1. Introduction

The monetary model of exchange-rate determination suggests a strong link between the nominal exchange rate and monetary fundamentals. The monetary model implies that the price level of a country is determined by its supply and demand for money and that the price level in different countries should be the same when expressed in the same currency. This makes it an attractive theoretical tool for understanding fluctuations in exchange rates over time. It also provides a long-run benchmark for the nominal exchange rate between two currencies and thus a clear criterion for determining whether a currency is significantly „overvalued" or „undervalued."

A number of early studies on industrial nations found little evidence of cointegration among nominal exchange rates and monetary fundamentals during the post-Bretton Woods float (see for example, (Meese, 1986), (Baillie – Selover, 1987), (McNown – Wallace, 1989), and (Baillie – Pecchenino, 1991)). The lack of empirical evidence for a stable long-run relationship among nominal exchange rates and monetary fundamentals implies that the monetary model has little practical relevance. A similar situation exists in the literature on empirical purchasing-power parity (PPP). Long-run PPP posits a stable long-run relationship between nominal exchange rates and relative price levels, but the empirical support for such relationships are limited when using data from the modern float. Given that PPP is a building block of the monetary model, it is not surprising that it is difficult to find evidence of cointegration between nominal exchange rates and monetary fundamentals during the modern float.

However, recent studies using long spans of data and/or panel data find support for long-run PPP for the post-Bretton Woods era, including (Frankel – Rose, 1996), (Papell, 1997), and (Taylor – Sarno, 1998), (Abuaf – Jorion, 1990), (Glen, 1992), (Lothian – Taylor, 1996, 2000), and (Taylor, 2001).

* Ankara University, Faculty of Political Sciences (civcir@politics.ankara.edu.tr)
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In regard to the monetary model, recent studies by Groen (2000), Mark and Sul (2001) and Rapach and Wohar (2001) test for a stable long-run relationship between nominal exchange rates and monetary fundamentals using panel cointegration tests for the post-Bretton Woods float. Interestingly, these studies find strong evidence of cointegration among nominal exchange rates, relative money, and relative real output using panel cointegration tests. Mark and Sul (2001) actually find support for a very simple long-run monetary model that imposes basic homogeneity restrictions. They also find that nominal exchange rate forecasts based on the monetary model are generally superior to forecasts of a naive random-walk model. The recent findings of Groen (2000) and Mark and Sul (2001) again renew hope in the ability of monetary fundamentals to track nominal exchange rates.

Previous studies on high-inflation countries show that monetary fundamentals are important in determining exchange-rate behavior (see (McNown – Wallace, 1994), (Bahmani-Oskooee – Kara, 2000), and (Moosa, 2000)).

In this paper, we test the monetary model on a high-inflation, developing country. The test of the monetary model is motivated by the findings of PPP in Turkish data (see (Civcir, 2002) and (Erlat, 2001)). In particular, we apply augmented Dickey-Fuller and Phillips-Perron unit-root tests and Johansen (1995) cointegration tests to monthly data for 1987:1–2000:12 in order to test the long-run validity of the monetary model of exchange-rate determination.

Our estimation results exhibit considerable support for the monetary model of U.S. dollar exchange-rate determination for Turkey. We find evidence of a theoretically consistent long-run link between nominal exchange rates and monetary fundamentals. Our findings are noteworthy given the lack of empirical support in much of the existing literature for the long-run relationship among exchange rates and monetary fundamentals implied by the monetary model. After finding support for the long-run monetary model, we consider two additional topics.

First, we test for the weak exogeneity of the nominal exchange rates and monetary fundamentals from the estimated vector-error-correction--models (VECM). This gives us insight into the adjustment process through which the long-run equilibrium relationship between exchange rates and monetary fundamentals is maintained. Second, we calculate misalignment from the estimated long-run relationship to evaluate whether the Turkish lira (TL) was overvalued before the eve of the 2001 financial crisis in Turkey (the country experienced a profound money-market crisis). Calculated misalignment shows substantial overvaluation of the TL *before* the crisis.

The rest of the paper is organized as follows: Section 2 presents the flexible-price and sticky-price monetary models of exchange-rate determination augmented with relative price differentials. Section 3 outlines data and testing strategy. Section 4 reports test results for the long-run monetary model, including unit-root and cointegration and weak-exogeneity tests and implied misalignment. Section 5 summarizes our main findings.

2. The Monetary Model

In this section, we focus on three versions of the monetary models, namely the Flexible Price Monetary Model (FPM), the Sticky Price Monetary Model (SPM), and the Sticky Price Monetary Model augmented with relative price differential.

The first building block of the monetary model assumes that PPP holds continuously:

$$s_t = p_t - p_t^* + c \tag{1}$$

where c is a constant, s is the logarithm of exchange rate expressed in units of domestic currency per foreign currency, and p and p^* are, respectively, domestic and foreign price levels. If $c = 0$, equation 1 implies that absolute PPP holds, and if $c \neq 0$, equation 1 implies that relative PPP holds.

The second building block of the model assumes a stable money-demand function in domestic and foreign countries. The money-market equilibrium conditions for domestic and foreign countries are assumed to depend on the logarithm of real income, y, and the logarithm of price level, p, and the nominal interest rate, i. An identical relationship can also be assumed for the foreign country (asterisks denote foreign variables). Monetary equilibrium in the domestic and foreign country can be computed in equation 2 and 3:

$$m_t = p_t + \beta_2 y_t - \beta_3 i_t \tag{2}$$

$$m_t^* = p_t^* + \beta_2^* y_t^* - \beta_3^* i_t^* \tag{3}$$

where m_t and m_t^* are the domestic and foreign money supply, respectively. β_2 is the income elasticity of demand for money and β_3 is the interest-rate semi-elasticity. On rearranging equation 2 and 3 for domestic and foreign price levels, and substituting them into equation 1, gives us a flexible-price monetary model of the exchange-rate equation of Bilson (1978), Frankel (1978), and Hodrick (1978):

$$s_t = \beta_1 (m_t - m_t^*) - \beta_2 (y_t - y_t^*) + \beta_3 (i_t - i_t^*) + c + \varepsilon_t \tag{4}$$

where βs are parameters and c is an arbitrary constant and ε_t is a disturbance term. Equation 4 assumes that an equilibrium exchange rate is driven by relative excess money supplies.[1]

In equation 4 the nominal interest rate is made up of two components, namely the real interest rate and the expected inflation rate, that is:

$$i_t = r_t + \pi_t^e \tag{5}$$

$$i_t^* = r_t^* + \pi_t^{e*} \tag{6}$$

[1] This specification assumes equal and opposite signs on relative money, income and interest rates, that is $\beta_i = -\beta_i^*$. The validity of these restrictions should be tested before estimating the model; however, due to degrees-of-freedom considerations, it is usually assumed away.

where r_t and r_t^* are the domestic and foreign real interest rate and π_t and π_t^e are the expected rates of domestic and foreign inflation, respectively. Assuming that the real interest rates are equalized in both countries, we have:

$$i_t - i_t^* = \pi_t^e - \pi_t^{e*} \tag{7}$$

Thus, equation 4 can be rewritten as:

$$s_t = \beta_1 (m_t - m_t^*) - \beta_2 (y_t - y_t^*) + \beta_3 (\pi_t^e - \pi_t^{e*}) + c + \varepsilon_t \tag{8}$$

Equation 8 is the Flexible Price Monetary Model (FPM). The coefficient of the relative money supply is positive and equal to one based on the neutrality of money. The rationale is that for a given percentage increase in the money supply, prices will increase by the same percentage. If PPP holds continuously, this would mean a depreciation of the domestic currency (s_t increase) by the same amount, in order to restore equilibrium. However, the prediction of a negative coefficient for relative income is opposite to what the Mundell-Fleming approach predicts. In the Mundell-Fleming model, higher real income will increase imports; this will worsen the trade balance and will require a depreciation of the domestic currency in order to restore equilibrium. In the FPM, a rise in domestic real income creates an excess demand for the domestic currency. Agents will then decrease their expenditures in order to increase their real money balances. This will lead to a fall in prices. Then by virtue of PPP, an appreciation of the domestic currency will ensure that equilibrium is restored.

Furthermore, an increase in the expected long-run inflation results in agents switching from domestic currency to bonds (both domestic and foreign). Thus the demand for domestic currency decreases, causing a depreciation of the domestic currency (an increase in s_t) and thus the coefficient of the relative expected rate of inflation is positive.

Frankel (1979) develops a SPM of the exchange rate that incorporates a short-run interest rate to capture liquidity effects. Frankel assumes that the expected rate of depreciation for the exchange rate is a positive function of the gap between the current exchange rate and the long-run equilibrium rate and the expected long-run inflation differential between the domestic and foreign countries. The yield is:

$$E(\dot{s}_t) = -\lambda(s_t - \bar{s}_t) + \pi_t^e - \pi_t^{e*} \tag{9}$$

where λ is the speed of adjustment to equilibrium. This equation states that the current exchange rate is expected to return to its long-run equilibrium at the rate of λ. In the long-run, $s_t - \bar{s}_t$, then the expected rate of currency depreciation, will equal the difference of domestic to foreign inflation. Combining equation 5, 6 and 9 gives:

$$s_t - \bar{s}_t = -\frac{1}{\lambda} [(i_t - \pi_t^e) - (i_t^* - \pi_t^{e*})] \tag{10}$$

Equation 10 shows that the gap between the current exchange rate and its long-run equilibrium exchange rate is proportional to the real interest differentials between the two countries. Thus, if the foreign real interest rate is higher than the domestic real interest rate, there will be capital outflows from domestic bonds to foreign bonds until the real interest rates are equalized.

The long-run PPP relationship in SPM is represented by:

$$\bar{s}_t = \bar{p}_t - \bar{p}_t^* \tag{11}$$

In the long run, the interest differential must be equal to the long-run expected inflation differential:

$$\bar{r}_t - \bar{r}_t^* = \pi_t^e - \pi_t^{e*} \tag{12}$$

Thus equation 10 can be rewritten as:

$$s_t - \bar{s}_t = -\frac{1}{\lambda}\left[(\bar{i}_t - i_t) - (\bar{i}_t^* - i_t^{e*})\right] \tag{13}$$

The above equation states that the exchange rate will overshoot its long-run equilibrium rate whenever the relative nominal interest differential increase above its equilibrium level.

Combining equation 4, 12 and 13 gives:

$$s_t = \beta_1(\bar{m}_t - \bar{m}_t^*) - \beta_2(\bar{y}_t - \bar{y}_t^*) + \beta_3(\pi_t^e - \pi_t^{e*}) + c + \varepsilon_t \tag{14}$$

Equation 14 is actually identical to the reduced equation of FPM, thus the SPM reduces to a FPM in the long run.

The short-run dynamic of the SPM is obtained by substituting equation 14 into 13, which constitutes the SPM of Dornbusch (1976) and Frankel (1979):

$$s_t = \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^*) + \beta_3(i_t - i_t^*) + \beta_4(\pi_t^e - \pi_t^{e*}) + c + \varepsilon_t \tag{15}$$

In equation 15, the FPM is nested within the reduced equation of SPM. According to equation 15, the signs of the coefficients of β_1, β_2 and β_4 are the same as that for FPM. The β_3 coefficient is negative; an increase in the domestic interest rate leads to a capital inflow, which increases the demand for the domestic currency and, in turn, leads to the appreciation of the domestic currency.

The monetary model of exchange rate traces movements in the exchange rate by examining monetary variables, with the crucial assumption that PPP is maintained between countries for broad price indices. The most recent study done by Civcir (2002) provides evidence for the weak form of PPP for Turkey where symmetry restrictions on the domestic and foreign price holds but unitary coefficients on the price is rejected. Further, Erlat (2001) uses sequential unit-root tests with shifts in trend and constant and fractional integration techniques and finds empirical support for PPP. Given these findings on PPP, we add relative prices to equation 15. This additional variable allows for movements in the relative prices of tradables to nontradables within and across countries (see (Cheung – Chinn, 1998) and (Husted – MacDonald, 1999)). The relative price variable may be determined by any number of factors. In the Balassa (1964) and Samuelson (1964) model, relative prices are driven by relative differentials in productivity in tradable and nontradable sectors. Relative prices may also be affected by demand-side factors (see (DeGregorio – Wolf, 1994)). In the long run, the rising preference for services, which are largely nontradable, may induce a secular trend in the relative price of nontradables. Hence, these Balassa-Samuelson and demand-side effects are proxied with a relative price variable.[2]

Our empirical monetary-exchange-rate model is augmented with relative prices of tradable to nontradables and can be written as:

$$s_t = \beta_1 m_t^d + \beta_2 y_t^d + \beta_3 i_t^d + \beta_4 \pi_t^d + \beta_5 P_t^{dTN} + c + \varepsilon_t \tag{16}$$

where $m_t^d = (m_t - m_t^*)$, $y_t^d = (y_t - y_t^*)$, $i_t^d = (i_t - i_t^*)$, $\pi_t^d = (\pi_t - \pi_t^*)$, and $P_t^{dTN} =$ $= [(P_t^T - P_t^N) - (P_t^{T*} - P_t^{N*})]$. The tradable price variable is proxied with producer price index (PPI) and nontradable price is proxied with consumer price index. Further, it is assumed that expected rate of inflation is equal to the actual rate of inflation.

3. Data and Methodology

3.1. Data

Most series are from the Central Bank of Turkey and IMF's *International Financial Statistic*s, and span the 1987:1–2000:12 periods. The exchange rate is average-of-month data, expressed in TL per USD unit. For the broad deflator, the consumer price index (CPI) *IFS* line 64 is used. The „tradable" price deflator is proxied by PPI or WPI (wholesale price index) data reported in *IFS* line 63.[3] The measure of money supply is the monthly average of broad money (M2). Monthly average industrial production was used as a proxy for real output. Short-term interest rates are here monthly average interbank rates for Turkey and monthly average federal funds rate for the United States. PPI and CPI are used as proxies for the relative price of tradables and nontradables, respectively. All variables are in natural logs except for interest rates.

[2] In principle, one would like to substitute out for the determinants of the relative price variable in the square brackets, especially since the price of tradables is likely to be endogenous with respect to the exchange rate. Unfortunately, sectoral-productivity data is not available at a monthly frequency for Turkey.

[3] In the empirical literature, tradable goods prices are usually proxied by WPI; see, for example, (Goldfajn – Waldes, 1996), (Hinkle–Nsengiyumva, 1999), and (Chinn, 2000)

3.2 Methodology

All variables in the models above can be considered to be endogenous and the possibility of short-run deviations from, and adjustments to, the long--run cointegration relationship makes VECM applicable.

Johansen cointegration analysis involves estimating the following VECM in reduced form:

$$\Delta z_t = \sum_{i=1}^{k} \Gamma_i \Delta z_{t-i} + \Phi z_{t-i} + \Psi d + \varepsilon_t \tag{17}$$

where z_t is a vector of nonstationary (in levels) variables, the matrix Φ has reduced rank equal to r and can be decomposed as $\Phi = \alpha\beta'$, where α and β are p x r full-rank matrices, and contains adjustment coefficients and the cointegrating vectors respectively. In the equation 17, d is the vector of deterministic variables which may include a constant term, the linear trend, seasonal dummies and impulse dummies. Finally, the error term is a normal process. Following Hendry and Doornik (1994) and Doornik et al. (1998), impulse indicator variables are entered unrestrictedly to the cointegration space.

In order to test for the number of cointegration relationships amongst the variables, Johansen (1988) and Johansen and Juselius (1990) provide two different tests to determine the number of cointegrating vectors, namely trace and maximum eigenvalue tests. In the trace test, the null hypothesis is that there are at most r cointegrating vectors and it is tested against a general alternative. In the maximum eigenvalue test, the null hypothesis of the r cointegrating vector is tested against r+1 cointegrating vectors.

Once we determine the number of relationships, r, we can conduct hypothesis testing on both loadings and cointegrating vectors. Restrictions can be imposed on the coefficients to test theory-based hypothesis for the long--run value of variables.

One problem with the Johansen procedure is that it is not able to precisely identify the parameters in α and β matrices. One can make concrete conclusions about unique long-run relationships between variables if there is one cointegrating vector found; otherwise, theoretical restrictions should be used to identify the long-run relationships.

4. Monetary Model Test Results

4.1 Unit Root Test Results

Before conducting an analysis of the long-run relationships between exchange rate and monetary fundamentals, we first investigate the time-series properties of s_t, m_t^d, y_t^d, i_t^d, π_t^d and P_t^{dTN} using augmented Dickey-Fuller (1979) unit-root tests. *Table 1* illustrates the augmented Dickey-Fuller test results for our data.[4] Columns A and B of Table 1 show unit-root tests re-

TABLE 1 ADF (k) Unit Root Test Results

LEVELS						FIRST DIFFERENCES			
Variables	k	A	B	$F3$	$F1$	Variables	k	A	B
s	12	-2.108	1.107			Δs	7	-4.374**	-4.142**
m^d (M1)	12	-1.643	1.300	4.887	8.206	Δm^d (M1)	7	-5.8990**	-5.5593**
m^d (M2)	12	-1.838	-0.130	3.634	5.885	Δm^d (M2)	7	-4.6472**	-4.6564**
y^d	12	-1.397	-1.397	4.613	0.367	Δy^d	7	-6.7452**	-6.7337**
i^d (TD3)	12	-2.995	-2.372	7.977	4.540	Δi^d (TD3)	7	-4.2638**	-4.3973**
i^d (IB)	12	-3.241	-2.887	2.334	7.667	Δi^d (IB)	7	-6.5833**	-6.6172**
π^d (CPI)	12	-0.867	-1.579	5.577	4.176	$\Delta \pi^d$ (CPI)	7	-4.7144**	-3.7911**
π^d (WPI)	12	-0.990	-1.473	1.026	2.840	$\Delta \pi^d$ (WPI)	7	-4.5790**	-3.6609**
P^{dTN}	12	-1.772	0.694	4.119	3.071	ΔP^{dTN}	7	-6.4151**	-6.2880**
1% Crt.Val*		-4.026	-3.478	8.730	6.700	1% Crt.Val*		-4.026	-3.478
5% Crt. Val		-3.443	-2.882	6.490	4.710	5% Crt. Val		-3.443	-2.882

Notes: 1. k is the amount of lagged dependent variables in the ADF regression. TD3 is the three-month time-deposit interest rate. IB is the Interbank interest rate.
2. Column A and B give the t-statistics from ADF regression including constant and trend, and constant respectively. Column $F3$ and $F1$ are Dickey-Fuller F statistics; the critical values are from D-F (1981).
3. The superscripts * and ** denote rejection at 5% and 1% critical values.

sults which are carried out by including linear trend and constant terms, and only constant term respectively. The inclusion of a linear trend is indicated by a visual inspection of the time series, as well as formal statistical F-tests of Dickey and Fuller (1981). Based on the unit-root test results in Table 1, we conclude that all of the variables are I(1).

The implications of our unit-root test results for testing the long-run monetary model is that cointegration procedures should be used. In the next subsection, we thus test for cointegration between the nominal exchange rate and relative money, income, interest rates, inflation, and prices for Turkey.

4.2 Cointegration Test Results

The Johansen procedure is used to determine the rank r and to identify a long-run monetary model of exchange rate amongst the cointegrating vectors. The first stage of estimating VECM is to determine the proper lag length. Lag-length decision is based on the evidence provided by both the likelihood ratio test and AIC; however, in the case of serial correlation, a sufficient number of lags are introduced to eliminate the serial correlation of the residuals. The cointegration tests amongst s_t, m_t^d, y_t^d, i_t^d, π_t^d and P_t^{dTN} include 12 lags in the VECM. To capture the effects of seasonality on the va-

[4] Phillips and Perron (1988) test results are almost the same, therefore, we do not present those results here, but they are available from the author upon request.

riables, we introduced a set of monthly-centered seasonal dummy variables, a constant term, and also three impulse dummy variables: D91 is included to capture the effects of the 1990 Gulf War, D94 is included to capture the 1994 currency crises in Turkey and D00 to capture the 2000 stabilization program in Turkey. The diagnostics in the form of vector statistics and single-equation statistics indicate that our VAR model is a satisfactorily close approximation to an actual data-generating process, apart from some non-normality of residuals.[5] Gonzalo (1994) has shown that the performance of the maximum-likelihood estimator of the cointegrating vectors is little affected by nonnormal errors.

Table 2 reports the estimates of the Johansen procedure and the relevant statistics for cointegration analysis. In determining the number of cointegrating vectors, we used a degrees-of-freedom adjusted version of the maximum eigenvalue and trace statistics, since given small samples with many variables or lags the Johansen procedure tends to overestimate the number of cointegrating vectors (see (Cheung – Lai, 1993) and (Gonzalo – Pitarakis, 2000)). These test statistics strongly reject the null hypothesis of no cointegration in favor of one cointegration relationship.[6] Table 2 also reports standardized eigenvectors, β', and adjustment coefficients, α. The first row of β' is the estimated cointegration vector, and can be written as:

$$\underset{\text{(std. err.)}}{s_t} = \underset{(0.147)}{5.384} + \underset{(0.034)}{0.827}m_t^d - \underset{(0.337)}{0.887}y_t^d - \underset{(0.0003)}{0.002}\, i_t^d + \underset{(0.002)}{0.025}\, \pi_t^d + \underset{(1.191)}{3.309}\, P_t^{dTN}$$

All of the coefficients in this vector have anticipated signs and are statistically significant; the likelihood ratio test statistics reported in the table also confirms the significance of the variables in the long-run relationships. The magnitudes of money and income (proxied by industrial production) differential (proxied by industrial production) variables are consistent with the monetary model. The interest differential enters with a negative sign, which indicates that an increase in Turkish interest rates relative to the U.S. rate results in the appreciation of the Turkish lira. The estimated value of the interest-rate differential variable is 0.002, which is very small. These findings are consistent with the sticky-price monetary model of the exchange rate. Inflation differential enters with a positive sign, which indicates that an increase in domestic inflation relative to US inflation leads to the depreciation of the domestic currency. Finally, the relative price variable has a positive sign and is statistically significant.

Table 2 also report α, the estimated response of each of the variables to the error-correction terms. The exchange rate responds to the error-correction term by moving to reduce the disequilibrium. The rate of response is very slow. Adjustment to the conditional mean appears to be affected by in-

[5] These results are available upon request from the author.

[6] However, without the degrees of freedom the adjustment result did not alter.

TABLE 2 Cointegration Analysis of the Monetary Exchange Rate Model

Eigenvalues	0.451	0.223	0.179	0.089	0.067	0.016
Hypotheses	$r = 0$	$r <= 1$	$r <= 2$	$r <= 3$	$r <= 4$	$r <= 5$
Lmax [-Tlog(1- mu)]	100.7**	42.39**	33.16*	15.640	11.580	2.750
Lmax [using T-nm]	57.52**	24.220	18.950	8.939	6.616	1.572
95% critical values	40.300	34.400	28.100	22.000	15.700	9.200
Ltrace [-T Sum log(.)]	206.2**	105.5**	63.13**	29.970	14.330	2.750
Ltrace [using T-nm]	117.8**	60.290	36.070	17.130	8.188	1.572
95% critical values	102.100	76.100	53.100	34.900	20.000	9.200
Standardized eigenvectors (Beta')						
s	m^d (M2)	y^d	i^d (IB)	π^d (WPI)	P^{dTN}	Constant
1.000	-0.827	0.887	0.002	-0.025	-3.309	-5.384
-1.333	1.000	-1.382	-0.003	0.013	5.126	6.865
0.310	-0.206	1.000	0.000	-0.009	-2.851	-1.649
-1187.900	1248.000	653.360	1.000	2.182	-6633.200	5225.400
-250.640	242.920	-766.440	0.117	1.000	-409.190	1204.800
0.843	-0.844	1.615	0.001	-0.008	1.000	-3.694
Standardized alpha coefficients						
s	-0.014298	-0.010349	0.026242	-6.7462E-06	-0.000015902	0.015388
m^d (M2)	-0.030865	-0.023695	-0.033756	-0.000009792	0.00010958	0.0055514
y^d	-0.14534	0.057451	0.041537	0.000044793	0.000062975	0.011056
i^d (IB)	-108.14	13.92	290.19	-0.086413	0.089881	-95.966
π^d (WPI)	13.093	2.1651	3.6213	0.001266	0.0084116	1.9543
P^{dTN}	-0.014668	-0.015634	0.01746	0.000018424	-0.000024858	-0.002852
Significance test statistics (Chi-sqr [1])						
s	m^d (M2)	y^d	i^d (IB)	π^d (WPI)	P^{dTN}	Constant
12.293	10.394	4.4309	24.387	44.328	10.052	14.108
[0.0005] **	[0.0013] **	[0.0353] *	[0.0000] **	[0.0000] **	[0.0015] **	[0.0002] **
Multivariate Unit Root Tests (Chi-sqr [6])						
	s	m^d (M2)	y^d	i^d (IB)	π^d (WPI)	P^{dTN}
	63.822	82.911	88.895	68.583	40.027	84.818
	[0.0000] **	[0.0000] **	[0.0000] **	[0.0000] **	[0.0000] **	[0.0000] **
Weak Exogeneity Tests (Chi-sqr [1])						
	s	m^d (M2)	y^d	i^d (IB)	π^d (WPI)	P^{dTN}
	3.741	4.6209	25.874	2.6284	38.827	3.4831
	[0.0531]	[0.0316] *	[0.0000] **	[0.1050]	[0.0000] **	[0.0620]

Notes: 1. The estimation period is 1987:1–2000:12. VAR includes 12 lags on each variable, a constant term, centered seasonal monthly dummy variables, D91 dummy and D94 dummy and D00 variables.
2. The Lmax and Ltrace are maximum-eigenvalue and trace-test statistics, adjusted for degrees of freedom. The critical values are taken from Osterwald-Lenum (1992).
3. The multivariate-stationarity, weak-exogeneity and significance-tests statistics are evaluated by assuming a single cointegration vector.
4. The * and ** indicate rejection of likelihood ratio tests at 5% and 1% significance levels, respectively.

terest-rate changes. When the error-correction term is negative at time t (so that TL is weak), then between time period t and t+1 TL strengthens and the domestic interest rate rises relative to the US rate. Therefore, an increase in the interest rate is associated with the strengthening domestic currency. Further, to a lesser extent adjustment is affected by monetary-policy changes. The estimated α coefficient for the inflation differential has a positive sign, which indicates that inflation has a tendency to push the system into a state of disequilibrium. Therefore, in high-inflation economies it is essential to control inflation in order to keep the system in equilibrium.

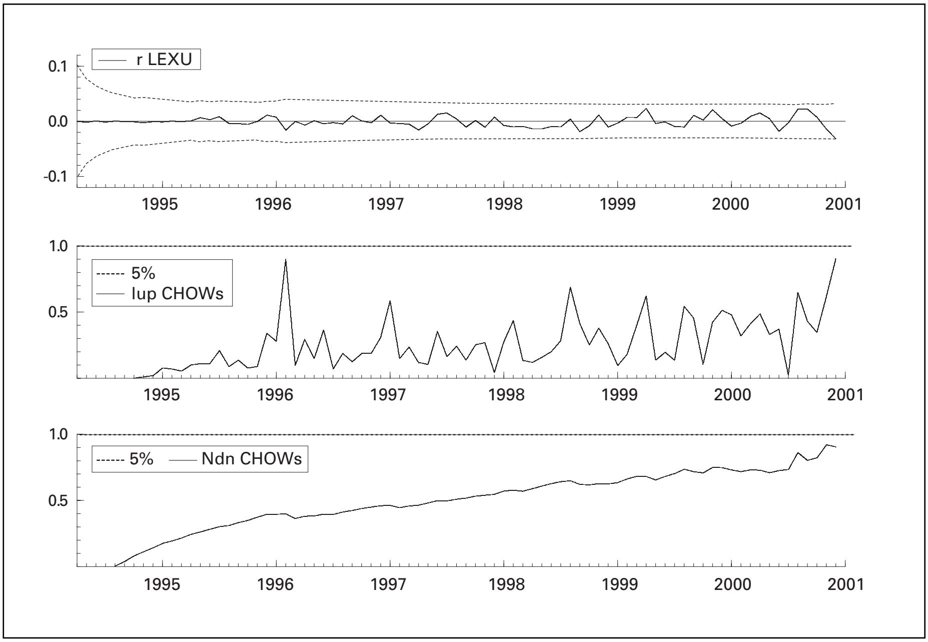
We can test various hypotheses on the parameters of the α matrix. An initial, interesting aspect is represented by the possibility of identifying the long-run weak exogeneity of the variable(s) with respect to the parameters of equilibrium relationships. A weak exogeneity test of a given variable for the cointegrating vector is presented in Table 2. The first term in α represents the speed at which the dependent variable in the first equation of the VECM moves toward restoring the long-run equilibrium, and the second term shows how fast the money differential responds to the short-run disequilibrium in the cointegration vector, and so forth. The test results show that the interest-rate differential and relative prices are weakly exogenous to long-run relationships at a 5% significance level, but we can not reject the weak exogeneity of relative prices at the 10% significance level. The weak exogeneity of the exchange-rate variable is on the border of the rejection area; however, given the small sample and the joint test statistics given below, we treat this variable as endogenous. The evidence found here is consistent with the fact that interest rates are mainly determined outside this system by the dynamics of the public-sector deficit in Turkey. The joint test of weak exogeneity shows that both of these variables are weakly exogenous at a 5% significance level; the likelihood ratio statistic is $\chi^2(2) = 5.864$ and the associated p-value of 0.0533 also confirms this result. The joint test for weak exogeneity including the exchange-rate, money, income, and inflation differentials rejects the null hypothesis of weak exogeneity; the corresponding likelihood ratio test statistics and the p-values are $\chi^2(4)= 70.999$ [0.000]. The weak exogeneity results also justifies a system approach to analyzing cointegration relationship and guides us in answering the question whether we have to model the exchange rate in a single equation or in a system context.

Finally, Table 2 also reports the multivariate stationarity of a given variable. The tests are based on the assumption that there is only one cointegrating vector. Here, the null hypothesis is the stationarity of the variable; furthermore, since it is multivariate and so includes a larger set of information, these statistics may have a stronger power against the univariate test (see (Johansen, 1995)). We reject the null hypothesis of stationarity for all the variables.

4.3 Constancy Test on the Long-run Equilibrium

Parameter constancy is an additional and crucial issue to ensure a well-specified equation. The potential for parameter instability increases sig-

FIGURE 1 Recursive Analysis of the Long-run Monetary Exchange Rate Model



nificantly during and after financial crisis, and the factors affecting the exchange rate may change. In order to evaluate the parameter stability, the cointegration analysis is redone by using the recursive estimation method. In this section, we report in *Figure 1* on a graphical instability test.[7] The first graph (a) shows one-step residuals from the monetary-exchange-rate model and the standard errors; the second graph (b) shows sequentially estimated one-step-ahead Chow statistics; and the third graphs (c) is the break--point Chow test.

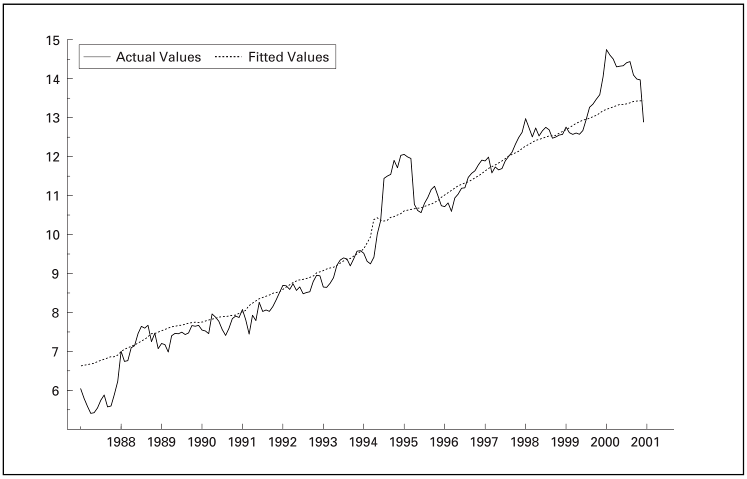
In the first graph, residuals lie inside the ± 2 s. e. bands, indicating that parameter constancy is not violated. The break-point Chow test for the sequence of (1994:2–2000:12, ...). None is statistically significant at the 5% level, indicating that parameter constancy cannot be rejected for the whole sequences of forecasts. Parameter constancy test indicates that, in general, the exchange-rate process in the long run remained unchanged over the sample period.

4.4 An Estimate of Equilibrium Exchange Rate

Recently, much literature has developed around testing the equilibrium exchange-rate relationship (see (Williamson, 1994), (MacDonald, 1995), (Hinkle – Monteil, 1999) (MacDonald, 2000)). Increasingly, both practitio-

[7] All the tests presented here employ the null hypothesis of parameter constancy.

FIGURE 2 Monetary Model Based Misalignment (in log)



ners and policy makers have been using such relationships to address issues of exchange-rate misalignment and also for assessment purposes. In the literature, the monetary model is widely used for testing the validity of the approach for exchange-rate determination and in terms of its out-of--sample forecasting properties. However, this model, or its variant, is not widely used for assessment purposes. Notably exceptions are (Chinn, 2000), (Husted – MacDonald, 1999), and (La Cour – MacDonald, 2000). These papers assess whether some currencies were overvalued or undervalued against the US dollar or the Japanese yen before the 1997 Asian crisis. In this paper, a cointegration vector is used to generate the equilibrium exchange rate, and misalignment is then calculated as a residual between the actual and generated equilibrium exchange rate. *Table 3* reports the implied misalignments for all of 2000. As of 2000:1, the Turkish lira was overvalued. Overvaluation increases leading up to the money-market crisis in February 2001.

Figure 2 presents a visual impression of a representative equilibrium and actual exchange rate. The figure clearly suggests that the TL was misaligned before the crisis. Of course, the derived equilibrium in this paper side

TABLE 3 Monetary Model Based Misalignment ($\hat{s}_t - s_t$)

Period	*2000:1*	*2000:2*	*2000:3*	*2000:4*	*2000:5*	*2000:6*
Misalignment	-0.15	-0.16	-0.17	-0.18	-0.2	-0.225
Period	*2000:7*	*2000:8*	*2000:9*	*2000:10*	*2000:11*	*2000:12*
Misalignment	-0.225	-0.225	-0.24	-0.26	-0.255	-0.25

Notes: 1. s is logarithm of TL per USD.
2. Estimated s is obtained from cointegrating vector. Misalignment is the residual between actual and estimated long-run exchange rate. Negative value indicates an overvaluation.

-steps the issue of the appropriateness of the underlying fundamentals, but nevertheless this kind of exercise is illustrative of the kind of equilibrium relationship our modeling strategy could be used to generate.

5. Summary and Conclusions

We have attempted in this paper to model the TL/USD exchange rate over 1987:1–2000:12 period using a variant of the monetary model of the exchange rate. We have conducted a test of the monetary-exchange-rate model augmented with relative prices for Turkey. A cointegration relationship between the exchange rate, monetary fundamentals and relative prices is evident. Thus, monetary fundamentals do affect the exchange rate in the long term.

Equilibrium relationships were also used to construct an equilibrium measure of the TL. The results in this paper argue that a sensible statement can be made about the equilibrium value of the TL/US dollar exchange rate. Is there any evidence in our analysis that the TL/US exchange rate was substantially overvalued on the eve of Turkish financial crisis of 2001? To answer this we compared actual and estimated values from the monetary model. The model suggests that the Turkish lira was overvalued vis-à-vis the US dollar before the crisis. This finding may be of interest to those concerned with exchange-rate assessment.

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SUMMARY

JEL classification: F31; F41; F47
keywords: exchange rates – monetary model – misalignment – Turkey

The Monetary Model of the Exchange Rate under High Inflation
The Case of the Turkish Lira/US Dollar

Irfan CIVCIR – Ankara University, Faculty of Political Sciences (civcir@politics.ankara.edu.tr)

This paper applies the Johansen cointegration technique to examine the validity of the monetary model of exchange-rate determination as an explanation of the Turkish lira/United States dollar relationship over the 1987:1–2000:12 period. A single cointegrating vector is identified whose coefficients conform in broad terms to the restrictions implied by the monetary model, thus lending support to the interpretation of the model as describing a long-run equilibrium relationship. This support is reinforced by the results derived from the adjustment coefficient, which identify a clear short-run tendency of the exchange rate to revert to the equilibrium value defined by the estimated long-run model. After finding support for the long-run monetary model, we calculate misalignment from the estimated long-run relationship to evaluate whether the lira was overvalued before the eve of the 2001 financial crisis in Turkey. Calculated misalignment shows a substantial overvaluation of the lira before the crisis.