Broad Money Demand and Currency Substitution in Turkey

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Abstract

The aim of this paper is to investigate the empirical relationship between broader definition of money, real income, interest rates, inflation and expected exchange rate, and to examine the constancy of this relationship, especially in the light of financial reform, deregulation of financial markets and financial crises. The estimation results show that long-run demand for real balances in Turkey depends upon real income, on the its own interest rate, interest rates on government securities, inflation and expected exchange rates. Significance of the expected exchange rate variable indicates existence of currency substitution in Turkey. The dynamics of money demand are important, the inflation and income affects are much smaller in the short-run than long-run. Our results also reveal that the demand for broad money in Turkey is stable, despite the economic reforms and financial crises.

1. Introduction

After serious balance of payment crises at the end 1970s, Turkey has changed its import substitution development strategy with outward-oriented ones.

In 1980s several reforms have been carried out including, liberalization of capital account and financial reform by the governments. After financial reform of 1980s, the Turkish financial system changed. The removal of foreign exchange controls and deregulation of financial markets have substantially changed the environment in which monetary policy operates. Government securities and interbank market have deepened, interest rate are determined freely, and reflects market sentiments, and new indirect instruments of monetary controls are developed.

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At the initial period of the 1980 reform program inflation had reduced to around 30 per cent, but between 1988 and 1993 it fluctuated around 70 per cent. With the surge of financial crises at the beginning of 1994 inflation reached three digit levels. A stand-by agreement with IMF and the re-functioning of the domestic debt market helped to reduce the strength of the crises, and inflation started to decline. After 1995 it fluctuated around 75 per cent per annum. The main factor behind the inflationary pressure in Turkey is the lack of budgetary discipline combined with monetary financing and/or domestic debt financing of budget deficit.

In response to these developments private sector has responded by requesting higher real interest rates on government securities and by increasing its foreign exchange holding of foreign currency. Turkish residents hold a considerable part of their financial wealth in German mark or US dollar (see Civcir, 2002). In this paper we provide empirical evidence on the currency substitution in Turkey.

The aim of this paper is to model the empirical relationship between broader definition of money, real income, interest rates, inflation and expected exchange rate. Our approach is novel in a number of ways. First, expected depreciation of the domestic currency is proxied with the exchange rate misalignment. Second, we use a multivariate cointegration technique to test for the existence of long-run relationship in contrast to previous works which uses two step Engle and Granger procedure. Third, our finding of cointegration and weak exogeneity test results facilitate an examination of the short-run money demand function using a dynamic equilibrium correction model. Finally, we examine the constancy of both short-run and long-run relationship, especially in the light of financial reform, deregulation of financial markets and financial crises.

The estimation results show that long-run demand for real balances in Turkey depends upon real income, on the its own interest rate, interest rates on government securities, inflation and expected exchange rates. The dynamics of money demand are important, the inflation and income affects are much smaller in the short-run than long-run.

This paper is organized as follows. Section 2 presents the model and the data. Section 3, analyses integration and cointegration properties of data, testing for cointegration and weak exogeneity of vector autoregression model. In section 4, a parsimonious error correction model of money demand estimated. Finally, the section 5 concludes.

2. The Model

Under currency substitution, the store of value services provided by a currency will determine its demand. Thus demand for low inflation currency is likely to increase. That is the store of value role of substitute currency increases

together with domestic inflation. The loss of purchasing power of domestic currency may motivate the demand for foreign currency. This will increase the role of money as a financial asset. Therefore, portfolio theory of money may be more appropriate to approach the demand for money balances in high inflation economies.

The portfolio balance approach was initially developed by McKinnon and Oates (1966). More recently, Girton and Roper (1981), Cuddington (1983), Branson and Henderson (1985) and Zervoyianni (1988, 1992) have extended the model to include the possibility of currency substitution.

These models assume that agents maximize the return to their wealth subject to given level of risk. Agents can hold four different assets and switch between them simultaneously. These assets are domestic money, domestic bonds, foreign money, and foreign bonds. In the empirical testing of currency substitution generally the demand for domestic money specification is used.

In general, empirical estimation of money demand functions based on the transactions and/or portfolio theories. The transaction theories view money as a medium of exchange and are demanded as an inventory for transaction purposes. Portfolio theories consider the demand for money in much broader terms as a part of problem of allocating wealth among a portfolio of assets, which includes money. The portfolio theories emphasize store of value function of money.

The demand for money as a financial asset is determined by the rate of return on the money itself, rate of returns on alternative assets, and by the total wealth (often proxied by income).

A long-run portfolio theory based money demand can be written as

$$m - p = m(y, R) \qquad (1)$$

where

m – nominal money,

p – the price level,

y – a scale variable,

R – the vector of rates of returns on various assets.

The money demand is increasing function of scale variable and those elements of R included in the m, and decreasing function of R those excluded from m (see Laidler, 1993, for the further details).

In an open economy, portfolio based money demand models, money is considered as part of portfolio, which consists of domestic financial and real assets, and foreign assets. The return on the domestic money is its own rate of interest. Regarding the assets alternative to money government securities (mainly Treasury bill) are important assets in the economic agents' portfolio, therefore the

rate of return on these assets will be included in the money demand function. The return on the real assets is usually represented by the expected rate of inflation, since the real value of money depreciates quickly under high inflation, while the real value of real assets maintained, the economic agents switches into real assets when the inflationary expectations are strong (for theoretical discussion see Friedman, 1956).

The currency substitution literature suggests that in an open economy with a liberalized capital account, foreign currency denominated assets forms an appropriate investment alternative. There will be the portfolio shifts between domestic and foreign currency, in which case expected exchange rate should be included in the money demand function. However, in the currency substitution literature estimating portfolio model with only expected exchange rate variable criticized by Cuddington and McKinnon on the ground that a model without expected exchange rate corrected foreign interest rate can not separate affect of currency substitution from capital flows. This factor focuses on the gradual elimination of administrative obstacles to free capital movements and the resulting opening of the Turkish financial system to the international markets that considerably broadened the residents' investment and financing opportunities, increasing range of alternative assets to those included in M2.

In accordance with foregoing, the set of variables in the money balance equation should be extended so as to add an interest rate on foreign assets. More specifically, exchange rate depreciation expectation adjusted foreign interest rate has been included in the money balance equation (see McKinnon, 1982; Cuddington, 1983; Giovannini and Turtelborn, 1993).

Following the discussion above a long-run money demand for Turkey can be written as

$$\frac{M}{P} = m(y, Rd, Rb, \pi, s^e, (s^e + Rf) \tag{2}$$

or in log linear form:

$$m - p = \beta_0 + \beta_1 y + \beta_2 Rd + \beta_3 Rb + \beta_4 \pi + \beta_5 s^e + \beta_6 (s^e + Rf) + \varepsilon \tag{3}$$

where

m – the logarithm of nominal money,
p – the logarithm of the price level,
y – the logarithm of real income,
Rd – interest rates on government securities,
Rb – the interest rates on Lira deposits,
π – the inflation rate,
s^e – the expected rate of change in the exchange rate
Rf – the foreign interest rates.

The expected signs and magnitudes of the coefficients are $\beta_1 > 0$, $\beta_2 > 0$, $\beta_3 < 0$, $\beta_4 < 0$, $\beta_5 < 0$ and $\beta_6 < 0$. In this model, β_5 measures the effect of currency substitution and β_6 measures the effect of capital movements. It will be tested whether the expected depreciation affects the demand for money directly or through $(s^e + Rf)$ or both.

3. The Data, Integration and Cointegration

3.1. The Data

The data used in the estimation of the money demand model are as follows:

M is broad money includes all Turkish Lira (TL) denominated currency in circulation plus demand and time deposits (foreign currency deposits are excluded) (TL billion);

P is the whole sale price index (1987 = 100), y is real GDP (TL billion, at 1987 factor costs), for GDP series only quarterly data is available, we converted quarterly data to monthly data by using industrial production index as an indicator;

Rb is the interest rates on government securities (monthly rate);

Rd is the interest rates on TL deposits (monthly rate- annual compounded rate converted to a monthly rate);

π is the monthly rate of inflation which is derived as monthly per centage change in WPI;

s^e is expected change in the exchange rate;

Rf is the three-month Eurodollar interest rate (monthly rate).

Expected exchange rate depreciation corrected foreign interest variable is highly collinear with expected exchange rate depreciation variable and the marginal contribution of foreign interest rate is very small. Therefore, we included foreign interest rate without correcting for expected exchange rate. The estimated coefficient of foreign interest rate had a wrong sign and statistically insignificant. Through out the period under consideration Turkey experienced high inflation, significant real appreciation of domestic currency and positive real interest rates on domestic securities. Therefore, only crises period (specifically 1994) matters for capital outflows. Also in a developing country home biasedness is highly significant phenomenon. Therefore, in the empirical specification this variable is excluded from the model.

Econometric studies on currency substitution have used different proxies for expected depreciation of the exchange rate. Ortiz (1983) used the difference between the official and real exchange rate for expected exchange rate depreciation. Cuddington (1983) used the ratio of the difference between forward and spot exchange rate to spot rate as a proxy for expected depreciation of the exchange

rate. Clements and Schwartz (1993) used inflation rate differential between U.S. and Bolivia. Selcuk (1994) used trade weighted real exchange rate index and TL per U.S. dollar nominal exchange rate as a proxy for expected depreciation of the exchange rate Turkey, and Akcay et al. (1997) used estimated exchange rate volatility based on the estimation of multi-currency version of Purchasing Power Parity (PPP) for Turkey. Boero and Tullio (1996) used interest rates differential and per centage deviation of the currencies from the PPP.

Recently, there has been renewed effort to understand the empirical determinants of currency crises. Most of the empirical studies shown that the real exchange rate deviation from trend or other form of calculating real exchange rate misalignments are important variables in predicting crises.

Kamisky et al. (1997) finds that the real exchange rate is the most reliable indicator in predicting future currency crises.[1] This result can be interpreted as a sign that this relative price is a key summary variable of several underlying fundamentals.

Goldfajn and Valdes (1996, 1998) analyzed large set of real exchange rate appreciation derived from an initial sample of monthly real exchange rates for 93 countries from 1960 to 1994 to evaluate whether real exchange rate misalignment lead to nominal devaluation. They found that probability of eliminating misalignment without nominal devaluation for different degree of misalignment is extremely low. That means real overvaluation usually corrected through nominal devaluation rather than inflation differentials.

The discussion above shows that real exchange rate misalignment can be taken as expected depreciation of exchange rate. Therefore, in this paper we will use real exchange rate misalignment based on PPP. In constructing real exchange rate whole sale price index is used. The misalignment is calculated as the deviation of the actual series from predicted series based on a regression of the real exchange rate on constant and trend. Increase in the RER reflects appreciation of domestic currency.

All the series are monthly and seasonally unadjusted and estimation sample extends from 1987 : 1 to 1999 : 12. Some evidence in the literature suggest that the length of the period is more important than the frequency of the data, we used monthly data for two reasons, some of the monetary variables are available as of 1986 : 1 and in a high inflation economy economic agents tend to make their decisions quite frequently in a small intervals.

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[1] Kamisky et al. (1997) also shows that interest rate differentials are not useful predicting crises. Interest differentials do not adequately reflects expected depreciation possibly due to the fact that changes in the interest differentials may reflect short run monetary policies that increase domestic interest rates or changes in the risk premium. These factors may produce enough noise that prevents extraction of reliable expected depreciation measures.

3.2. Integration and Cointegration

This section presents unit root test for the variables of interest. Then Johansen's (1988, 1991) maximum likelihood procedure is applied to test for cointegration among real money, real output, interest rates on government securities, interest rates on time deposits, inflation and expected exchange.[2]

Integration

Before modeling money demand the univariate unit roots were tested to determine the order of integration of the variables. Table 1 reports augmented Dickey-Fuller (1981) test statistics.[3] The unit root tests are given for the level and first differences (changes) of the data. The lag lengths for the ADF tests are determined by the AIC and likelihood ratio tests. When the serial correlation problem detected further lag(s) added to eliminate the serial correlation. Following the suggestion of Dickey and Pantula (1987) the unit root tests are first performed for two roots, and if two roots rejected then single unit root tested for. Tests are carried out with constant term included in the ADF regression.

Table 1

ADF(k) Unit Root Test Results

Levels				*First differences*		
Variables	k	A		Variables	k	A
M	12	1.591		ΔM	10	–8.394**
$(m-p)$	12	0.506		$\Delta(m-p)$	10	–4.301**
Rb	6	–1.791		ΔRb	7	–5.134**
Rd	6	–1.656		ΔRd	7	–5.406**
Rf	3	–1.607		ΔRf	4	–3.646**
Y	12	0.198		Δy	10	–21.633**
π	12	–2.503		$\Delta\pi$	10	–6.353**
S^e	3	–2.231		Δs^e	2	–5.724**
1 % Critical Value		–2.882				
5 % Critical Value		–3.478				

Notes:

[1] k is the number of lagged dependent variables in the ADF regression.

[2] Column A gives the t-statistics from ADF regression with constant.

[3] For the real GDP series centred seasonal dummy variables included as additional regressors in the ADF regression.

[4] The critical values are from MacKinnon (1991). The superscripts * and ** denotes rejection at 5 per cent and 1 per cent critical values.

[2] The expected exchange rate corrected foreign interest rate has been included in the model, however due to strong multicollinearity of this variable with expected exchange rate, foreign interest rate without correcting for expected depreciation included in the model. Estimation result showed that this variable has wrong sign and statistically insignificant. Therefore, we did not include this variable into our model.

[3] Phillips and Perron (1988) unit root tests are also performed and the results are very similar to the ADF tests. Therefore, they are not reported here but available from the author upon request.

Test results show that none of the variable seems to show evidence of two unit roots, and all the variables are not able to reject the null hypothesis of single unit root. That is, all the variables are non-stationary in levels but stationary after fist differencing.

Cointegration

The cointegration tests in this paper are conducted by using reduced rank procedure developed by Johansen (1988) and Johansen and Juselius (1990). This method should produce asymptotically optimal estimates because it incorporates a parametric correction for serial correlation and the system nature of the estimator means that the estimates are robust to simultaneity bias, and it is robust to departure from normality (Cheung and Lai, 1993) and Johansen, 1995). Johansen method detects number of cointegrating vectors in non-stationary time series and allows for hypothesis testing regarding the elements of cointegrating vectors and loading matrix.

Basically, the cointegration analysis takes place in the following unrestricted vector autoregressive (VAR) framework,

$$\Delta x_t = \sum_{i=1}^{k} \Gamma_i \Delta x_{t-i} + \Phi x_{t-1} + \Psi d + \varepsilon_t \qquad (4)$$

where x_t is vector of non-stationary (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 $p \times r$ full rank matrices, and contains adjustment coefficients and the cointegrating vectors respectively. D is the deterministic variables which may include constant term, the linear trend, seasonal dummies and impulse dummies. Finally, the error term is normal process.

In order to test for the number of cointegration relationships amongst the variables Johansen (1988) and Johansen and Juselius (1990) provides 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 r cointegrating vector is tested against $r+1$ cointegrating vectors.

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

One problem with Johansen and Juselius procedure is that it is not able to identify exactly the parameters in α and β matrices. Only if there is just one cointegrating vector found than we can make truly concrete conclusions about any unique long-run relationship between the variables, otherwise we can not.

In the Johansen's cointegration analysis deterministic part of the system is also important. Doornik et al. (1998) statistically analyzed over-specified trend

in the cointegration space and suggest that adopting a model that includes a trend in the cointegration space have low cost even when Gross Domestic Product (GDP) does not display tend. They found that including an unrestricted trend was problematic.

However, a restricted trend in the cointegration space with an unrestricted constant produced a good power and reasonable size (for further details see Doornik et al. (1998). Franses (1999) also suggests that exclusion of deterministic trend from cointegration space is not safe.

Concerning the treatment of impulse indicator variables they strongly recommended that these to be entered unrestrictedly if they are used to establish an estimate of the innovation variance. They certainly advise against their restriction to the cointegrating space (see Hendry and Doornik, 1994).

Following these suggestions, we have included dummy variables unrestrictedly into the cointegration space.

Cointegration Results

Johansen procedure is used to determine the rank r and to identify a long-run money demand amongst the cointegrating vectors.

The number of lags used in the VAR is based on the evidence provided by both likelihood ratio test and AIC, however, in the case of serial correlation sufficient number of lags introduced to eliminate the serial correlation of the residuals. The cointegration tests amongst $m\text{-}p$, y, Rb, Rd, π and s^e include seven lags in the VAR. As real GDP is affected by seasonality, we introduced a set of monthly centred seasonal dummy variables (see Johansen, 1995), a constant term and further, the estimates of unrestricted VAR include also two impulse dummy variables: D8889 is included to capture the interest rate intervention during 1988 : 10 – 1989 : 3 period, and D94 is included to capture currency crises in 1994.

Table 2

Full System Diagnostics

Statistics	Values	p-values
Vector portmanteau 12 lags	331.88	
Vector AR 1 – 7 F(252,205)	1.1955	[0.0916]
Vector normality $\chi^2(12)$	74.698	[0.0000] **
Vector $\chi^2(1764)$	1742.1	[0.6401]

Notes:

[1] The results are obtained using PcFiml version 9.2 (see Doornik and Hendry, 1997).

The diagnostics in the form of vector statistics reported in Table 2 indicates that our VAR model is satisfactorily a close approximation to actual data generating process, apart from some non-normality of residuals.

Table 3

Cointegration Tests

Cointegration Tests						
Eigenvalues	0.381	0.266	0.205	0.115	0.042	0.010
Hypotheses	r = 0	r ≤ 1	r ≤ 2	R ≤ 3	r ≤ 4	r ≤ 5
λmax	45.57**	29.440	21.770	11.590	4.122	0.983
95 % crt. Values	39.400	33.500	27.100	21.000	14.100	3.800
λtrace	113.5**	67.910	38.470	16.700	5.105	0.983
95 % crt. Values	94.200	68.500	47.200	29.700	15.400	3.800
Standardized eigenvectors (β')						
(m − p)	*y*	*Rb*	*Rd*	*π*	s^e	
1.000	**−1.009**	**13.393**	**−17.613**	**31.333**	**1.461**	
−0.314	1.000	3.056	−10.298	−0.278	−0.169	
0.056	−0.107	1.000	0.115	−0.447	0.027	
−0.001	0.008	−0.544	1.000	−0.175	0.008	
0.424	−0.538	1.944	−0.056	1.000	0.549	
−4.890	3.942	−33.597	24.873	2.859	1.000	

Notes:

[1] VAR includes seven lags on each variable, a constant term, centred seasonal monthly dummy variables, d8889 dummy, d9192 dummy and d94 dummy variables. The estimation period is 1987 : 1 – 1999 : 12. None of the deterministic variable is restricted to the cointegration space.

[2] The λ_{max} and λ_{trace} are maximum eigenvalue and trace test statistics, adjusted for degrees of freedom. The critical values are taken from Osterwald-Lenum (1992). The * and ** indicate rejection of likelihood ratio tests at 5 per cent and 1 per cent significance levels, respectively.

Table 3 reports the estimates of Johansen procedure and standard statistics. In determining the number of cointegrating vectors we used degrees of freedom adjusted version of the maximum eigenvalue and trace statistics, since in the existence of small samples with too many variables or lags Johansen procedure tends to over estimates the number of cointegrating vectors (see Cheung and Lai, 1993; and Gonzalo and Pitarakis, 1994). These test statistics strongly rejects the null hypothesis of no cointegration in favor of one cointegration relationship.[4] Table 3 also reports standardized eigenvectors, β', and adjustment coefficients, α. The first row of β' is the estimated cointegration vector, can be written as

$$(m\text{-}p)_t = \mathbf{1.009}y_t - \mathbf{13.393}Rb_t + \mathbf{17.613}Rd_t - \mathbf{31.333}\pi_t - \mathbf{1.461}s_t^e$$

(t-stat) (2.049) (2.418) (1.988) (6.062) (2.805)

All of the coefficients in this vector have anticipated signs and statistically significant. The unit income homogeneity restriction is not rejected, the associated likelihood ratio statistic and asymptotic *p*-values are $\chi^2(1) = 0.0002$ and [0.99] respectively. Furthermore, the semi elasticity of own rates are approximately the equal magnitude and opposite sign to the Treasury bill rate has been tested[5] and we cannot reject this hypothesis $\chi^2(1) = 0.176$ [0.675].

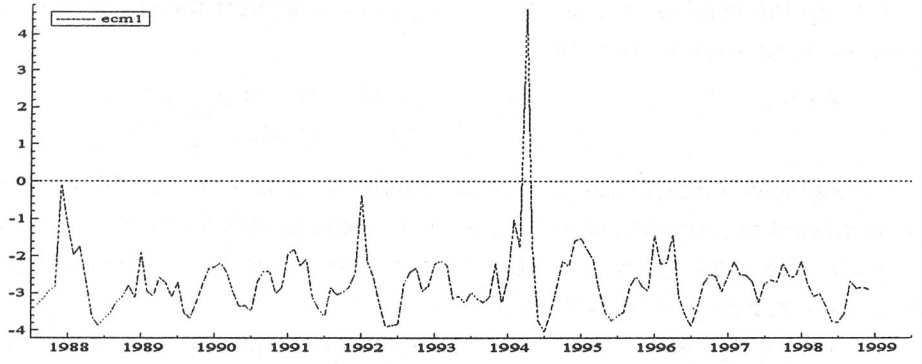
[4] However, without degrees of freedom adjustment result did not alter.

[5] We also experimented with a spread between its own rates and Tbill rates, this measure is negatively related to the real money balances.

The inflation has semi elasticity of 31.33 (when the valid restrictions imposed on the cointegrating vector coefficient on the inflation reduces to 26.37 further, this magnitude is sensitive to number of lags in the VAR, while it is apparently high, this elasticity is similar to those obtained in studies of broad money demand for Turkey, Metin (1999) reports 22.24 and on the other countries (see Ericsson, 1998). Like in many developing countries the impact of the expected inflation on real money balances has been much more pronounced in Turkey. This is in accord with the view that agents sway away from money holding to real assets when the inflation expected to rise.

Figure 1

Cointegration Relationship



The coefficient on the expected depreciation variable has correct sign and statistically significant, this indicates an existence of currency substitution in Turkey. However, the estimated coefficient of this variable is quite low. Given the historical developments in period under consideration, this is not surprising, since in the most of this period Turkey has experienced positive interest rates on the domestic financial assets, and domestic currency appreciated significantly. Further more, the real rates of return on domestic financial assets exceeded that of foreign currency deposits, yet the ratio of foreign currency deposits to the broad money, (M2) is stayed significantly high in this period. Therefore, the small coefficient on the expected depreciation variable can be attributed to the existence of dollarization hysteresis in Turkey.

We can test various hypotheses on the parameters of α matrix. A first interesting aspect is represented by the possibility of identifying long-run weak exogeneity of the variable(s) with respect to the parameters of equilibrium relationships. If the cointegration vector do not have any influence on a particular variable, in which case, all the weights will be equals to zero, then that variable is said to be long-run weakly exogenous for the long-run parameters. The test

results show that output, both interest rates on Treasury bill and domestic currency deposits are weakly exogenous for real money balances. The joint test statistics $\chi^2(3) = 1.447$ [0.695] also confirms this result. However, the joint test including the inflation and the expected depreciation of exchange rate variables weak exogeneity cannot be rejected, the corresponding likelihood ratio test statistics and the p-values are $\chi^2(5) = 39.914$ [0.00]. We also tested jointly weak exogeneity of output, both interest rates on Treasury bill and domestic currency deposits, plus long-run unit income elasticity and equal but opposite signs on interest rates, LR statistics is $\chi^2(5) = 1.939$ [0.858], we cannot reject these restrictions on the long-run money demand relationship. The evidence found here consistent with the fact that interest rates are determined outside the system by the dynamics of the public sector deficit.

With all the valid restrictions (except equality of interest rates) imposed, the money demand equation becomes:

$$(m\text{-}p)_t = 1y_t \underset{(3.290)}{- 16.324\, Rb_t} + \underset{(3.230)}{18.17\, Rd_t} - \underset{(5.694)}{26.374\, \pi_t} - \underset{(2.969)}{1.339\, s_t^e}$$

The restricted cointegration regression result also shows that all the variables are significant at 1 per cent significance level. The restricted feedback coefficient is –0.046. The restricted cointegration vector is economically meaningful representation of a money demand function.

The weak exogeneity results justifies a system approach to analyzing cointegration relationship and guides us in answering the question whether we have to model the money demand in a single equation or in a system context. Further, in the following subsection, by using the above results we will analyze the money demand relationship within a three equation system. Specifically, we will open the system and condition on the weakly exogenous variables.

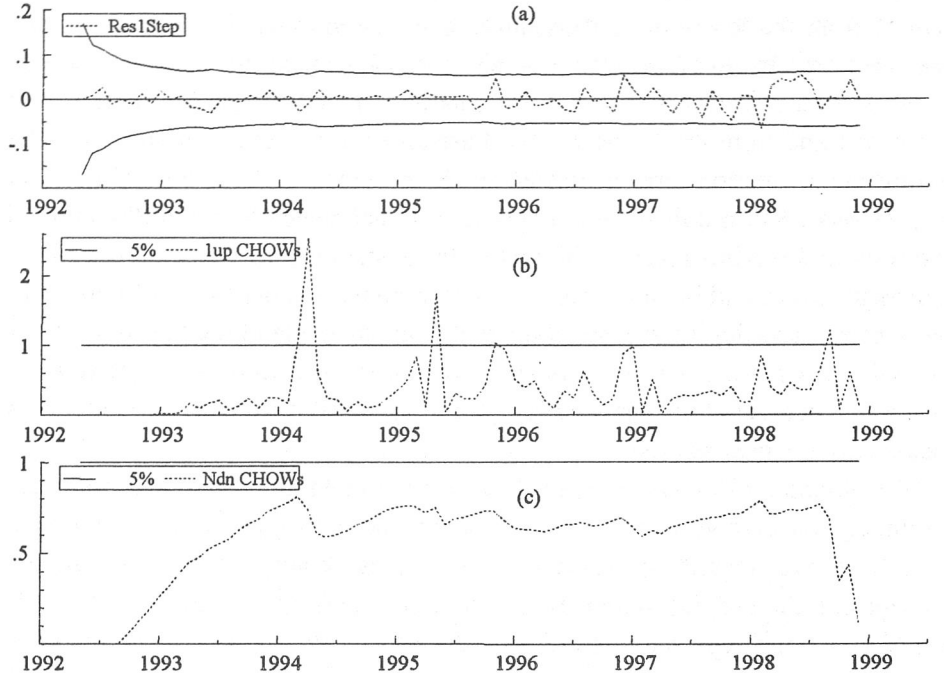
3.3. Constancy Test on the Long-run Equilibrium

Parameter constancy is an additional and crucial issue to ensure well specified equation. The potential for parameter instability increases significantly during and after financial crises, and the factors affecting money demand 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 graphical instability test[6] in Figure 2. The first graph (a) shows one step residuals and the money demand equation standard errors, second graph (b) shows sequentially estimated 1-step ahead Chow statistics, third graphs (c) break point chow test.

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

Figure 2

Recursive Analysis of the Long-run Money Demand Relationship



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 (1992 : 6 – 1999 : 12, ...) None is statistically significant at 5 per cent level indicating that constancy of the parameters can not be rejected for the whole sequences of forecasts. In the one step Chow test, only 3 points are above the 5 per cent significance level, but these are numerically small.

Constancy of the parameters indicates that, in general, the money demand process in the long run remained unchanged over the sample period.

4. Equilibrium Correction Model

Estimated cointegration relationship reveals factors affecting long-run real currency demand. In the short-run, deviation from this relationship could occur this reflects shocks to any of relevant variables. Furthermore, the dynamics governing the short-run behavior of real currency demand are different from those in the long run. Engle and Granger (1987) showed that if there is a cointegration relationship between nonstationary variables, there must be an equilibrium correction

representation of the data. In this section, based on the estimation of the cointegration relationship between *(m-p)*, *y*, *Rb*, *Rd*, π and s^e we proceed with the estimation of the equilibrium correction representation, taking into account the deviations from the long-run relationship and short-run dynamics of the real currency demand. In this representation, short-run dynamics are modeled by estimating in first differences. The adjustments in response to the deviation of real money demand from the long-term level are taken into account by including the equilibrium correction term estimated in the previous section. Furthermore, in the previous section unit income hypothesis is not rejected, and the likelihood ratio tests on the adjustment coefficient of the system we have not rejected weak exogeneity of the real income, interest rates on bonds and deposits independently and jointly. Accordingly, we will analyze the money demand relationship further within three equation system, namely real money, inflation and expected exchange rate depreciation. The stability of the estimated error correction model is discussed in the next section.

The estimated VAR system is reparameterized in equilibrium correction form and through successive steps, reduced to a parsimonious representation (PVAR). The test of over identifying restrictions allows us to judge whether the model encompasses the general system from which it is derived and can then be considered as a valid representation of the data generating process of the modeled series.

The equilibrium correction models were estimated with full information maximum likelihood (FIML) for period 1987 : I – 1999 : XII minus the included lags. The system was initially estimated with seven lags. The final lag structure is determined based on the significance of the each variable in each equation. Before coming to the interpretation of the results, we need to look at the diagnostic statistics of the system, given in Table 4.

Table 4

Parsimonious System Diagnostics

Statistics	Values	p-values
Vector portmanteau 12 lags	77.250	
Vector AR 1 – 12 F(108, 228)	0.707	[0.9787]
Vector normality $\chi^2(6)$	31.288	[0.0000] **
Likelihood ratio test for over identifying restrictions $\chi^2(120)$	119.227	[0.503]

The system diagnostic statistics do not indicate any misspecification in the model, except non-normality of the residuals. Since the normal distribution is only of limited importance for our inference we do not consider these results as problematic. However, single equation statistics shows that residual of the real

money equation does not violate normality assumption. We still consider it as a valid representation of the underlying process that describes the series and a good balance between the need to have parsimony in the number of parameters and to have a congruent model. The likelihood ratio test for over identifying restriction is not rejected, suggesting that the restricted dynamic model parsimoniously encompasses the VAR.

Table 5

Parsimonious FIML Conditional ECM

Dependent Variable : $\Delta(m-p)$				
Variable	Coefficient	Std.Error	t-prob	HCSE
$\Delta(m-p)_{t-2}$	0.2336	0.0696	0.0011	0.0653
$\Delta(m-p)_{t-3}$	0.2378	0.0589	0.0001	0.0506
$\Delta(m-p)_{t-6}$	0.1293	0.0552	0.0210	0.0504
$\Delta\pi_t$	–1.6611	0.2175	0.0000	0.2316
$\Delta\pi_{t-7}$	0.3266	0.0760	0.0000	0.0636
Δs^e_{t-1}	–0.2249	0.0740	0.0029	0.0792
Δs^e_{t-2}	0.2246	0.0780	0.0047	0.0870
Δs^e_{t-6}	–0.1587	0.0711	0.0276	0.0684
Δy_t	0.0512	0.0202	0.0127	0.0191
Δy_{t-1}	0.0822	0.0182	0.0000	0.0184
Δy_{t-2}	–0.1165	0.0204	0.0000	0.0198
Δy_{t-3}	–0.0871	0.0238	0.0004	0.0190
Δy_{t-5}	0.0675	0.0197	0.0009	0.0191
Δy_{t-7}	–0.0901	0.0260	0.0008	0.0246
ΔRb_t	–4.8530	0.7207	0.0000	0.6964
ΔRb_{t-4}	–0.7033	0.4185	0.0956	0.3587
ΔRb_{t-5}	–0.7397	0.3584	0.0413	0.3351
ΔRd_t	3.9229	1.0200	0.0002	0.9653
ΔRd_{t-1}	4.2655	0.6118	0.0000	0.5325
ΔRd_{t-2}	2.5886	0.5401	0.0000	0.5797
ΔRd_{t-4}	2.1051	0.6012	0.0007	0.5678
SD	–0.0892	0.0123	0.0000	0.0118
SD2	–0.0377	0.0139	0.0076	0.0139
SD4	–0.0367	0.0124	0.0038	0.0088
SD7	–0.0772	0.0118	0.0000	0.0115
Constant	–0.1128	0.0201	0.0000	0.0218
D8889	–0.0289	0.0134	0.0330	0.0138
D94	0.0356	0.0094	0.0002	0.0084
ECM_{t-1}	–0.0408	0.0071	0.0000	0.0077
σ= 0.0280111				

Notes:

[1] HCSE is White's robust standard error. Sdi is centred seasonal dummy variables for January through November.

In Table 5, we report the estimates of the $\Delta(m - p)$ equation which can be interpreted as a dynamic money demand function.[7] An important feature to notice is the significance of the equilibria in the money demand equation. In the

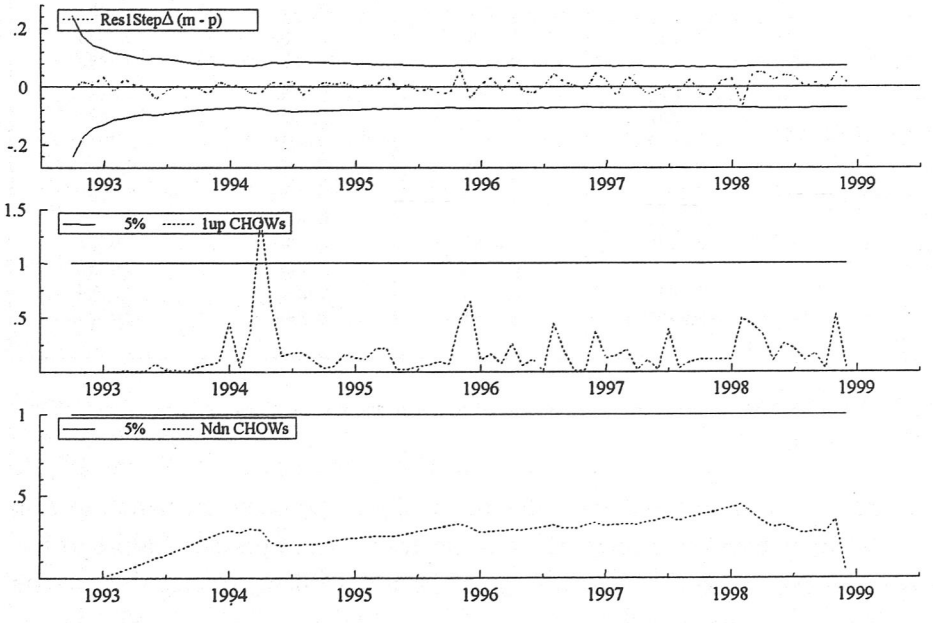
[7] The inflation and expected depreciation equation results are not presented here, but available upon request from the author.

short run agents adjust their holdings by 4 per cent of the past month's deviation from equilibrium. Negative sign on EC term indicates that agents have corrected in the current period a proportion of the previous period's disequilibrium in money balances. Specifically, a fall in excess money in the last period will result in higher level of desired money balances in the current period. That is, it is essential for maintaining long-run equilibrium to reduce the existing disequilibrium over time. In addition to this disequilibrium effect, agents respond to interest rate changes, changes in inflation, and changes in the expected depreciation of the exchange rate which have theoretical signs. The expected depreciation of the exchange rates is highly significant. It is noticeable that the cumulative sum of the transaction level does not seem to affect short-run real money demand. This however, could be due to the measurement error on the monthly data.

Recursive estimation of the system allows the detection of possible instability in the parameters. In the graphical test shown in Figure 3, there is no sign of major parameter instability as in the case of cointegration relationship. However, in the 1 step Chow test only one point is above the 5 per cent significance level in 1994. Constancy of the parameters indicates that, in general, the money demand process in the short-run also remained unchanged over the sample period.

Figure 3

Recursive Analysis of Short-run Model



Conclusion

The paper models the money demand relationship for Turkey for the period of 1987 : I – 1999 : XII. The period characterized by the high PSBR/GNP ratio, high inflation, financial liberalization, capital account liberalization, and financial innovation driven mainly by an increasing government debt.

Empirical analysis carried out by means of Johansen multivariate cointegration analysis and constrained error correction models. Cointegration analysis reveals that there is a stationary long-run relationship between broad money balances, real income, interest rates on demand deposits and Treasury bill rates, inflation and expected exchange rate depreciation. Significance of the expected exchange rate variable in the money demand function indicates currency substitution in Turkey.

The paper finds evidence that real money demand both in the long-run and short-run in Turkey remained stable throughout the period under investigation. This might provide justification for the monetary authority to target broad money. However, currency substitution may complicate the choice of intermediate targets of the monetary policy, which introduces foreign currency components into money supply. The currency substitution do not seem to undermine the stability of the money demand in Turkey. The suitability of the targets that exclude or include foreign currency depends on the target's relationship with inflation and output. Various measures of monetary aggregate's relationship with the end-objective, such as inflation rates, are not addressed in this paper.

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DOPYT PO PENIAZOCH V ŠIRŠOM POŇATÍ A MENOVÁ SUBSTITÚCIA V TURECKU

Irfan CIVCIR

Cieľom predloženého článku je preskúmať empirický vzťah medzi širšie definovanými peniazmi, reálnym príjmom, úrokovými mierami, infláciou a očakávaným devízovým kurzom, ako aj otestovať stálosť tohto vzťahu, osobitne v súvislosti s finančnou reformou, deregulóciou finančných trhov a finančných kríz.

Predbežné výsledky poukazujú na to, že dlhodobé úsilie o reálnu bilanciu v Turecku závisí od reálnych dôchodkov, od všeobecnej úrokovej miery, od úrokovej miery vládnych cenných papierov, inflácie a očakávaných výmenných kurzov. Význam premennej *očakávaný devízový kurz* svedčí o existencii menovej substitúcie v Turecku. Dôležitá je dynamika dopytu po peniazoch, vplyvy inflácie a dôchodkov sú omnoho menšie v krátkodobom než dlhodobom aspekte. Výsledky štúdie taktiež ukazujú, že v širšom chápaní dopyt po peniazoch je v Turecku stabilný, a to aj napriek ekonomickým reformám a finančným krízam.