THE TURKISH BROAD MONEY DEMAND

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ABSTRACT
In this paper, we try to examine the Turkish M2Y broad money demand and its determinants for the period 1987.1-2004.2 with quarterly data. For this purpose, we first specify the construction of a money demand model, and give a literature review of international evidence for empirical studies carried out. By using modern econometric techniques, then, we construct an empirical broad money demand model for Turkish economy, and compare the estimated results with the findings of some other empirical money demand studies carried out on Turkish economy. The main findings of our study indicate that the broad money demand is insensitive to real income, and we attribute this case to the highly unstable growth performance of the economy and the rapid financial innovation process which decreases the correlation between monetary and income agregates. Money demand function indicates instabilities within estimation period, probably because of domestic economic crises conditions and political uncertanties. Also, the main determinant of our money demand model is estimated as inflationary expectations.

Keywords: ***Broad Money Demand, Turkish Economy, Cointegration.***

TÜRKİYE EKONOMİSİ İÇİN GENİŞ TANIMLI PARA TALEBİ

ÖZET
Çalışmamızda 1987.01-2004.02 dönemi için üçer aylık veriler kullanılarak M2Y para talebinin belirleyicileri incelenmeye çalışılmaktadır. Bu amaçla öncelikle para talebi modelinin oluşumu incelenmekte ve konuyla ilgili yazın taraması özet bilgiler sunularak gerçekleştirilmeye çalışılmaktadır. Daha sonra Türkiye ekonomisi üzerine bir para talebi uygulaması gerçekleştirilmekte ve tahmin edilen sonuçlar diğer bazı uygulamalı çalışmalar ile karşılaştırılmaktadır. Çalışmamız sonucu elde edilen başlıca bulgular M2Y para talebinin reel gelire karşı duyarlılık göstermediği ve tahmin edilen para talebi ilişkisinin inceleme dönemi içerisinde önemli istikrarsızlıklar gösterdiği şeklindedir. Bu sonuçlar ise ekonominin göstermiş olduğu hayli istikrarsız büyüme sürecine ve finansal piyasaların yaşadığı hızlı değişime atfedilmektedir. Ayrıca incelenen para talebi ilişkisinin başlıca belirleyici unsuru enflasyon olgusu olarak tahmin edilmektedir.

Anahtar kelimeler: ***Geniş Para Talebi, Türkiye Ekonomisi, Eşbütünleşim.***

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1. INTRODUCTION

The phenomenon money demand deals with for what motives the people wish to hold the money balances. By deducing from the estimations of the money demand equations, the monetary authority can decide about which monetary policies is better to implement under the current economic conditions. A stable demand function for money has long been perceived as a prerequisite for the use of monetary aggregates in the conduct of policy (Goldfeld and Sichel, 1990: 300). For instance in a situation where the demand for real money balances which should be under the control of monetary authority is perceived with an endogeneous charasteristics to the other economic aggregates, the monetary authority cannot probably follow an independent monetary policy to attain the ex-ante specified targets. Also if an unstable characteristic for these money balances in the time period under investigation is estimated, this case can indicate the invalidity of the operations of the monetary authority based on these ex-ante money demand estimation results, that is, the policies based on these results can take the monetary authority to implement the wrong policies for the specified targets. As Kontolemis (2002: 3) expresses, stability of long run money demand function is an important factor of long run growth rates of monetary variables. Otherwise, disorderly or repeated velocity shocks are likely to lead to persistent deviations of growth of monetary aggregates from estimated values, which leads to errors in the formulation of monetary policy.

For the empirical estimation purposes, we can distinguish the motives of demand for money into mainly two behavioral assumptions; the transactions and the asset or portfolio balance approaches. The approaches emphasizing the importance of the transactions motive specify the money's role as a medium of exchange. Especially the well-known studies of Baumol (1952: 545-556) and Tobin (1956: 241-247) develop the underpinnings of this approach. For this approach, money is viewed essentially an inventory held for the transaction purposes. Transaction costs of going between money and other liquid financial assets justify holding such inventories, even though other assets offer higher yields (Judd and Scadding, 1982: 994). In this approach, the demand for money balances increases proportionally with the volume of transactions in the economy, while decreases with the increase of returns in the alternative costs of holding money. For the portfolio balance approach, we mean that people hold money as a store of value, and money is only one of the assets among which people distribute their wealth. People give more importance to the expected rate of return for the assets held relative to the transaction necessities, also considering a longer time period, and should take into account the risk factor for these assets because of the probable changing ratio of returns against each other. Thus we can say that the basic contribution of the portfolio balance approach is to enter the risk considerations explicitly into the determination of the demand for money (Branson, 1989: 328). Friedman (1956: 3-21) and also Friedman (1959: 327-351) with an influential empirical study which highligts the new quantity theory and Tobin (1958: 65-86) can mainly be considered as the studies emphasizing the importance of the risk factor and the portfolio decision for the demand for money.

In this paper, we try to specify the determinants of the broad money demand for the Turkish economy by constructing an empirical model, and to test it by using modern econometric estimation techniques. Thus, our focus inclines on the portfolio balance theory of the demand for money. The next section gives a literature review with international evidence of demand for broad money balances. The third section interests with data issues and model specification, and also estimates an empirical model for the Turkish economy. And the final section concludes.

2. LITERATURE SURVEY

This section includes the literature review of the empirical studies which gives international evidence of the demand for broad money balances using cointegration and vector error correction techniques.

Choudry (1995: 77-91) attempts to determine whether there exists a stationary long run money demand function for M1 and M2 aggregates in Argentina, Israel and Mexico. He finds a strong support for a stationary money demand function in the long run in all three countries. But the results estimated only hold with considering the effect of currency substitution in the money demand function. Since currency substitution reduces the domestic monetary control that is ensured by a flexible exchange rate, and also reduces the base of the inflation tax and the financing of deficit by means of seignorage, he suggests policies that reduce currency substitution.

Triechel (1997: 1-27) examines the stability of M2 and M4 broad money functions for Tunisia, with respect to conducting of monetary policy. Based on the money demand estimations, he suggests a base regime in which after the prediction of a multiplier the monetary base is manipulated so as to achieve a certain growth of the money supply with exogeneous interest rates for monetary authority, instead of a price regime in which short term interest rates are targeted so as to be consistent with the growth rate of money supply with the endogeneous base money supply.

Dekle and Pradhan (1997: 1-38) examines the impact of financial market development and liberalization on money demand behavior in Indonesia, Malaysia, Singapore and Thailand by using both narrowly defined and broad monetary aggregates. They find instability characteristics of various money demand equations estimated, and relate this case to the rapid growth and ongoing changes in financial markets.

Eitrheim (1998: 339-354) investigates the long run relationship between money, prices and wages in Norway. He finds broad money as endogenously determined, and that monetary balances were exposed to large shocks during the period of financial deregulation, while in the long run these shocks are absorbed in a way causing a long run money demand relationship.

Vega (1998: 387-400) analyses the stability of broad money demand for Spain. His results indicate that shifts affecting the broad money demand alter its long run properties, and as causes of this structural break he finds that increasing openness of the Spanish financial system to international markets increases the instability of the money demand.

Ericsson and Sharma (1998: 417-436) develops an equilibrium correction model of M3 broad money for Greece. Their results indicate that the estimated model is constant in spite of large fluctuations in the inflation rate, the introduction of new financial instruments and liberalization of the financial system. They estimate that the long run demand for money depends upon real income with a unit elasticity and the own interest rate, and that assets outside money affect money demand through a spread between their rate of return and the rate of return on broad money. Also in a short run dynamic specification, it is found that dynamics of money demand are important with price and income elasticities being much smaller in the short run than in the long run.

Nachega (2001: 1-39) investigates the behavior of M2 broad money demand for Cameroon. In an open economy modeling framework, he estimates unitary income elasticity which is consonant with the quantity theory of money. Also the results estimated indicate positive sensitivity of broad money to own rate of return and negative sensitivity to the rate of inflation, currency substitution and foreign interest rate. Besides, the estimation process of money demand relationship reveals the source of inflation as imported.

Kontolemis (2002: 1-30) reviews the stability of long run M3 money demand in the Euro area, and finds a stable long run money demand with a slow speed of adjustment back to equilibrium than the European Central Bank estimates, and relates this case to the velocity shocks.

For Turkish economy; in a study comparing backward and forward looking appoaches to modeling money demand, Yavan (1993: 381-416) estimates M2 broad money demand for the period 1980.1-1991.2 with quarterly data. By using different estimation techniques, he finds inflationary expectations as the most dominant factor affecting money demand. He explains this result in a such way that the expectations of economic agents catch up the inflation rate extensively, and this case enables them to get rid of inflation tax by reducing their monetary holdings.

Metin (1994: 231-256) estimates M1 narrow money demand for the period 1948.1-1987.4 with quarterly data. The results estimated confirm the existence of a long run money demand relationship with a quitely high positive income elasticity and also a negative inflation elasticity as opportunity cost for the money demand equation.

Koğar (1995: 1-18) tries to test whether there exists a stable long run money demand function for Turkey and Israel, which experience high inflation during the analyzed period. For the Turkish case, using quarterly data in the period 1978.1-1990.4, it is found that there exists a long run relationship between real money (M1 and M2) demand, real income, inflation and exchange rate with an elasticity of income quitely lower than unity and also an elasticity of exchange rate highly low.

Civcir (2000: 1-31) models the empirical relationship between M2 broad money, real income, interest rates and expected exchange rate. He thus examines the constancy of this relationship in the light of financial reforms, deregulation of financial markets and financial crises. The results obtained indicate the existence of a stable real broad money demand relationship with a positive unitary income elasticity confirming the quantity theory of money and negative opportunity cost variables. He expresses that this case might provide justification for the monetary authority to target broad money, together with considering the effect of dolarization.

Mutluer and Barlas (2002: 55-75) analyzes the Turkish broad money demand of deposits denominated in foreign currency for the period 1987-2001 with quarterly data. Their results also indicate the existence of a long run relationship for real broad money in Turkey, with a unitary income elasticity estimated, as was in Civcir (2000: 1-31). The dominant factors affecting the broad money demand in their model are the inflation rate and the CPI based real effective exchange rate established by CBRT, as returns of alternative assets.

Akıncı (2003: 1-25) models the demand for real cash balances in Turkey for the period 1987.1-2003.3 with quarterly data. The estimated results indicate that there exists a long run relationship between real currency issued, private consumption expenditure as scale variable, interest rates on government securities and the exchange rate. In the long run, the income elasticity is found to be close to unity, and the opprtunity cost variables have the expected negative magnitudes.

Altınkemer (2004) investigates the base money demand function for Turkey under an assumption of rational expectations, and she succeeds in estimating a stable long run base money demand function. Also a stable long run M2Y function is estimated. The empirical findings indicate the joint endogeneity of inflation and real base money, which does not support the possibility of monetary targeting for Turkey and also give an indirect support for the alternative targeting regimes, specifically for inflation targeting. For policy purposes, however, it is expressed that it is better to target and also keep an eye on the developments of base money till the conditions for inflation targeting mature and even after that, in the view that money can play at least informational role for an inflation targeting framework.

3. DATA and MODEL SPECIFICATION

While investigating the money demand function, a critical point to consider is the identification problem. By this notion, we mean the non-observability of the money demand. We can only measure the quantity of money supplied. And in this point, we have to make an important supposition that the quantity of money supplied and demanded equal each other, thus assuming equilibrium in the money market (Laidler, 1973: 85). For the transaction purposes, we can suppose that the narrowly defined monetary variables are better to consider, while the broadly defined monetary variables are better for the portfolio balance approaches of the money demand equations.

After defining the money demand variable, narrowly or broadly for our purpose, we should choose the explanatory factors affecting this variable. First, we should choose the scale-income variable which specifies the maximum limit of money balances we can hold (Keyder, 1998: 283). This choice can vary for the motive the demand for money is considered. For example if we mainly interest with the volume of transactions in the economy, the current real national income or a scale variable representing the expenditure-sided approaches would be suitable for our aims. Thus, the current real gross national (domestic) product or private consumption expenditures in the national income accounts can be used for this variable. But if our aim is to investigate the portfolio balance approach, the expexted or permanent income variable considering the weighted averages of subsequent income periods or a wealth variable representing the values of all the tangible assets in the economy would be better to consider for our demand for money function. But in the economics literature, this variable is also represented by the current real gross national product because of ease of use and calculation. The expected sign for this variable is positive.

Since the money demand function interests with for what motives people hold this balances on their hand, we should as a next step determine what alternative costs are current in the economy, thus discouraging people to hold this balances. These alternative costs may be the interest rate on bonds in home and foreign country, returns of equities, changes in the exchange rate and also the inflation rate representing the increase of prices of intangible assets under the assumption of substitution between commodities and domestic money. An expectation of an inrease of prices for all these assets would probably decrease the demand for money, thus we expect a negative coefficient for these variables. Besides, in a money demand equation we can add a variable representing the own return of the monetary variable considered. This return would be expected with a positive relationship with the demand for this variable.

We now construct a model of money demand for our empirical purposes by using quarterly data. We use a variety of econometric procedures available in the program Eviews 4.1. All the data we use are from the CBRT electronic data delivery system and indicate seasonally unadjusted values except the real income variable. The sample period for all the time series is 1987.1 – 2004.2.

The monetary variable we consider is the broad money supply in logarithmic form, that is M2Y with the end of period values, which is the sum of currency in circulation, demand deposits and time deposits in domestic currency in the banking system and also deposits denominated in foreign currencies. This choice can reflect the responsiveness of a broader range of financial system to changes in explanatory variables than a narrowly defined monetary aggregate. But as Mutluer and Barlas (2002: 55-75) says, a narrower definition would also be more flexible and reactive to market operations and interest rate policies of the monetary authority.

For the scale-income variable, we use gross national product at constant 1987 prices. The aggregates representing national income can normally be expected to indicate seasonality, thus for estimation purposes they are used in a de-seasonalized form. In our analysis, we also use this variable in a de-seasonalized form by using US Census Bureau's X-12 seasonal adjustment program within Eviews 4.1.

The variables representing alternative costs to hold money are the maximum rate of interest on the Treasury bills whose maturity are at most twelve months or less, with the exception for the period 2000.1 for which we use the interest rate on the government bond whose maturity is thirteen months, and the annualized quarterly inflation rate based on consumer price index. Following the modern literature on this issue, we use these variables in the level form, not in a logarithmic scale. Akıncı (2003: 1-25) argues that nominal interest rates are alone sufficient in the money demand models. The justification is that when there is a moderate inflation in the economy variations in the nominal interest rate can capture the variations in the expected rate of inflation. But as also she accepts, the inflation variable can have an impact on money demand through channels other than the nominal interest rates. So we try to determine the effects of both inflation and interest rate on real money balances demanded. For additionally alternative cost variables, the stock exchange index values of ISE (İMKB) and a proxy variable for foreign interest rates in an open economy framework can be added to our model. These might be for instance U.S. treasury bond or London interbank offer rate (LIBOR) which can additionally be used in a cointegration analysis with a larger time period or a high frequency data. But in this paper we do not use these variable specifications.

As a last variable for our money demand equation, we consider the nominal exchange rate defined as TL / \$US in logarithmic form, and accept that this variable indicates the own rate of return for the broad money balances because of high positive correlation between M2Y aggregate and the price of US. dollar. The estimated correlation between these variables is 0.972193. Two dummies which take on values of unity concerning financial crises occured in 1994 and 2001 are considered as exogeneous variables. By considering the effect of these economic crises, especially for the post-1994 period, we expect a positive coefficient for dummy variables in the VECM specification (see Figure 1). Under the assumption of no money illusion which is quitely reasonable for a cronic-high inflationary country, we can suppose that the demand for money is a demand for real money balances. In our case, we use the consumer price index (CPI) to deflate the broad money supply. So below is the our demand for money relationship as a functional form of endogeneous variables normalized on real broad money demand. The expected signs are indicated under the variables,

$$(LNRM2Y) = f\,(\underset{+}{LNRY},\ \underset{-}{ENFLASYON},\ \underset{-}{BONOFA\dot{I}Z},\ \underset{+}{LNDOLAR}) \qquad (1)$$

where,

LNRM2Y = real money balances in natural logarithm for the M2Y aggregate deflated by CPI

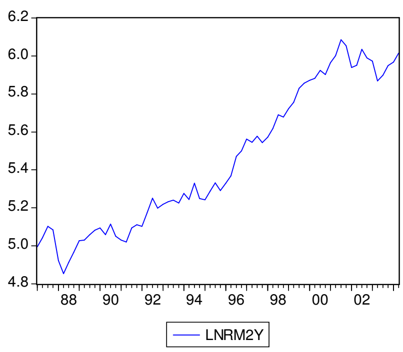
LNRY = gross national product in 1987 prices in natural logarithm in a de-seasonalized form

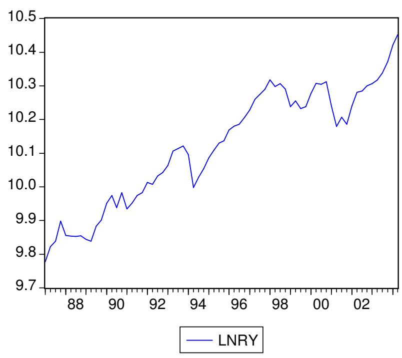
ENFLASYON = annualized quarterly inflation rate based on CPI

BONOFAİZ = the maximum rate of interest on Treasury bills whose maturity is at most twelve months.

LNDOLAR = nominal exchange rate defined as TL / $US in natural logarithm

Below is shown the graphical representation of the time series used in the analyses,





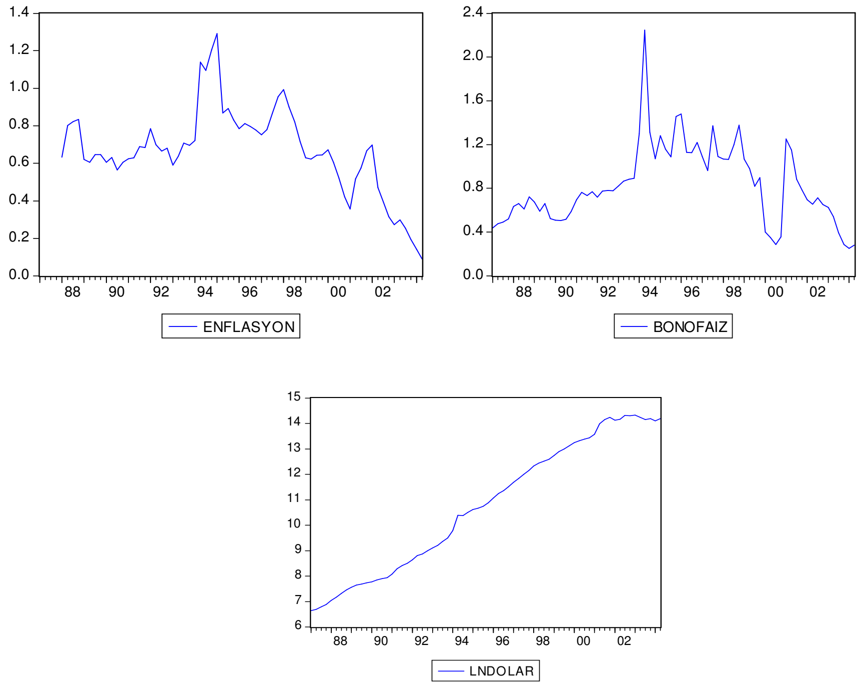


Figure 1

As a next step for our econometric analysis, we investigate the time series properties of the variables used. Granger and Newbold (1974: 111-120) indicates the occurance of the spurious regression problem in the case of using non-stationary time series, causing unreliable correlations within the regression analysis. At first, by using the augmented Dickey-Fuller (ADF) unit root test (Dickey and Fuller, 1979: 427-431) and the Phillips-Perron unit root test (Phillips and Perron, 1988: 335-346), we check for the stationarity condition of our variables by comparing ADF statistics obtained, with the MacKinnon (1996: 601-618) critical values, also possible in Eviews 4.1. For the case of stationarity, we expect that the ADF statistic is larger than the MacKinnon critical values in absolute value and that it has a minus sign. Although differencing eliminates trend, we also report the results of unit root tests for the first differences of variables with a linear time trend in the test regression. The results are shown in Table 1 below,

Table 1[a,b,c]
Unit Root Tests

	ADF Test		Phillips-Perron Test	
	Contant	Constant&Trend	Constant	Constant&Trend
Variable				
LNRM2Y	-0.2025(0)	-2.5655(0)	-0.0010(14)	-2.4713(9)
LNRY	-0.6573(4)	-1.9498(4)	-0.8588(2)	-2.7662(3)
ENFLASYON	-0.1112(4)	-0.9198(4)	-1.2248(3)	-1.9331(3)
BONOFAİZ	-2.8787(0)***	-2.8051(0)	-2.8787(0)***	-2.8051(0)
LNDOLAR	-1.2770(1)	-0.5384(1)	-1.2308(3)	-0.2497(3)
DLNRM2Y	-7.7571(0)*	-7.7350(0)*	-7.8775(14)*	-7.9007(14)*
DLNRY	-5.8230(3)*	-5.7729(3)*	-7.5581(2)*	-7.4962(2)*
DENFLASYON	-5.7966(3)*	-6.2105(3)*	-7.8103(3)*	-7.8455(2)*
DBONOFAİZ	-8.1789(1)*	-8.3023(1)*	-10.781(9)*	-12.338(11)*
DLNDOLAR	-5.8489(0)*	-5.9959(0)*	-5.8481(1)*	-5.9959(0)*

MacKinnon (1996) critical values

	Constant	Constant&Trend
%1 level	-3.53	-4.10
%5 level	-2.90	-3.48
%10 level	-2.59	-3.17

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[a] For the MacKinnon critical values, we consider %1, %5 and %10 level critical values for the null hypothesis of a unit root for the both unit root tests.
[b] The letter 'D' beginning of a variable indicates the first difference operator.
[c] The numbers in paranthesis are the lags used for the ADF stationarity test augmented up to a maximum of 12 lags, and the automatic bandwidth using Newey-West bandwidth selection method. The choice of the optimum lag for the ADF test was decided on the basis of minimizing the Schwarz Information Criterion (SIC). The test statistics and the critical values are from the ADF or UNITROOT procedures in Eviews 4.1. ADF is the augmented Dickey-Fuller test with critical values based on MacKinnon (1996, pp. 601-618). A significant test statistic rejects the null hypothesis in favor of stationarity. '*' , '**' and '***' indicate the rejection of the null hypothesis of a unit root for the %1, %5 and %10 level respectively.

When we examine the results of the unit root tests, we see that the null hypothesis that there is a unit root cannot be rejected for all the variables with both constant and constant & trend terms in the test equation in the level form. But inversely, for the first differences of all the five variables, the null hypothesis of a unit root is strongly rejected at 1% level. As a result, we accept that all the five variables contain a unit root, that is, non-stationary in their level forms, but stationary in their first differenced forms, thus enable us testing for cointegration.

We now examine whether the variables used are cointegrated or not. Engle and Granger (1987: 251-276) indicates that even though economic time series may be non-stationary in their level forms, there may exist some linear combination of these variables that converges to a long run relationship over time. If the series are individually stationary after differencing, but a linear combination of their levels is stationary, then the series are said to be cointegrated. That is, they cannot move too far away from each other in a theoretical sense (Dickey, Jansen and Thornton, 1991: 58). For this purpose, we estimate a VAR-based cointegration relationship using the methodology developed in Johansen (1991: 1551-1580) and Johansen (1995) to specify the long run relationship between the variables. Let us assume a VAR of order p

$$y_t=A_1y_{t-1}+\ldots+A_py_{t-p}+Bx_t+\varepsilon_t \qquad (2)$$

where y_t is a k-vector of non-stationarity I(1) variables, x_t is a d-vector of deterministic variables as constant term, linear trend and seasonal dummies, and ε_t is a vector of innovations. We can rewrite this VAR as

$$\Delta y_t=\Pi y_{t-1}+\sum_{i=1}^{p-1} \Gamma_i\Delta y_{t-i} + Bx_t + \varepsilon_t \qquad (3)$$

where

$$\Pi=\sum_{i=1}^{p} A_i-I \qquad \Gamma_i = -\sum_{j=i+1}^{p} A_j \qquad (4)$$

Granger representation theorem asserts that if the coefficient matrix Π has reduced rank $r<k$, then there exists kxr matrices α and β each with rank r such that $\Pi=\alpha\beta'$ and $\beta' y_t$ is I(0). r is the number of cointegrating relations (the rank), and each column of β is the cointegrating vector. The elements of α are known as the adjustment parameters in the VEC model, and measure the speed of adjustment of particular variables with respect to a disturbance in the equilibrium relationship. Johansen's method is to estimate the Π matrix from an unrestricted VAR and to test whether we can reject the restrictions implied by the reduced rank of Π. Also we can express that this method performs better than other estimation methods even when the errors are nonnormal distributed, or when the dynamics are unknown, and the model is overparametrized by including additional lags in the error correction model (ECM) (Gonzalo, 1994: 225). Thus, we first determine the lag length of our unrestricted VAR model, for the maximum lag number selected is 8, by using five lag order selection criterions, that is, sequential modified LR test statistic (LR), final predicton error criterion (FPE), Akaike information criterion (FPE), Schwarz information criterion (SC) and Hannan-Quinn information criterion (HQ). We think that the maximum lag number selected as 4 can restrict the estimation capability of a VAR-based cointegration process with the quarterly data. So we consider the maximum lag number as 8 with the quarterly data. As the lag order to be selected, LR test statistic suggests 5, FPE, AIC and HQ suggest 8, and also SC suggests 1 lag orders. We choose the lag order selected by sequential LR statistic, that is 5, to check our econometric model for the cointegration specification. Below other lag criterions will be briefly investigated for our cointegration specification.

As a next step, we estimate the long run cointegrating relationship(s) between the variables by using two likelihood test statistics offered by Johansen and Juselius (1990: 169-210) known as maximum eigenvalue for the null hypothesis of r versus the alternative of r+1 cointegrating relationships and trace for the null hypothesis of r cointegrating relations against the alternative of k cointegrating relations, for r = 0,1,...,k-1 where k is the number of endogeneous variables, to find the number of cointegration relationships. For the trace test, the alternative of k cointegrating relationships corresponds to the case where none of the series has a unit root, and a stationary VAR may be specified in terms of the levels of all of the variables. Table 2 reports the results of max-eigen and trace tests with a linear deterministic trend.

Table 2
Cointegration Rank Test

Sample (adjusted) : 1989.3 2004.2
Included observations : 60 after adjusting endpoints
Trend assumption : Linear deterministic trend
Series : LNRM2Y LNRY ENFLASYON BONOFAİZ LNDOLAR
Exogeneous series : DUMMY1 DUMMY2
Lags interval (in first differences) : 1 to 5

Unrestricted Cointegration Rank Test

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	5 Percent Critical Value	1 Percent Critical Value
None**	0.473724	80.58238	68.52	76.07
Atmost 1	0.333665	42.06667	47.21	54.06
Atmost 2	0.156065	17.70888	29.68	35.65
Atmost 3	0.080835	7.528081	15.41	20.04
Atmost 4	0.040342	2.470716	3.76	6.65

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	5 Percent Critical Value	1 Percent Critical Value
None*	0.473724	38.51572	33.46	38.77
At most 1	0.333635	24.35779	27.07	32.24
At most 2	0.156065	10.18080	20.97	25.52
At most 3	0.080835	5.057365	14.07	18.63
At most 4	0.040342	2.470716	3.76	6.65

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*(**) denotes rejection of hypothesis at the 5%(1%) level
Trace test indicates 1 cointegrating equation at both 5% and 1% levels
Max-eigenvalue test indicates 1 cointegrating equation at the 5% level
The critical values are taken from Osterwald-Lenum (1992: 461-472), also available from the VAR and COINT procedures in Eviews 4.1.

From the Table 2, for both trace and max-eigen test statistics, the null hypothesis of no cointegration is rejected at the 5% level in favor of one cointegrating relationship. Our results support the cointegration rank of 1, thus indicating a long run equilibrium, co-movement relationship between the variables used. Below is shown the cointegrating vector after normalizing on the variable LNRM2Y by dividing each variable by the negative of the variable LNRM2Y to obtain economically meaningful results,

LNRM2Y	LNRY	ENFLASYON	BONOFAİZ	LNDOLAR	(5)
-1.00000	+0.130360	-0.292862	-0.116956	+0.131091	
	(0.15928)	(0.11915)	(0.07059)	(0.01103)	

Asymptotic standard errors are reported in parantheses.
Adjustment Coefficients (std. err. in parantheses)

D(LNRM2Y)	D(LNRY)	D(ENFLASYON)	D(BONOFAİZ)	D(LNDOLAR)
-0.584748	-0.146946	-0.375815	-0.718149	-0.882169
(0.19262)	(0.11799)	(0.34847)	(1.02203)	(0.40097)

We should express that the normalization of the system by restricting a trend effect in the long run space produces the same results with a statistically insignificant trend effect. When we examine the results of our cointegrating equation, we see that all the variables have the expected signs with our a priori expectations and are statistically significant except the variable LNRY. An increase in the opportunity cost of holding broad money balances, especially the inflation rate with a semi-elasticity, reduces demand for these balances. Our estimations also indicate that the broad money demand is sensitive to the own-rate of return with a positive relationship as we have expected.

Besides, the variable representing real income, LNRY, has a quite low positive elasticity for a broadly defined monetary aggregate, and also it is statistically insignificant in a way not in accordance with the quantity theory of money. This case can indicate the non–or low- monetization of the economy by the monetary authority, and also the endogeneous characteristics of the monetary variables which should be considered in an economic policy perspective of monetary authority, or the rapid financial innovation process which decreases the correlation between the real monetary and the real income aggregates. The financial innovation period of Turkish economy between 1987 and 2003 does not coincide with a steady growth period of real national income. Under this consideration, the growth trend of the monetary variables do not have to follow the growth path of the real income variables. Especially for the narrowly defined monetary aggregates which can be used by the monetary authority for monetary targeting purposes, should have not been surprised to estimate a zero income elasticity for the Turkish economy as is in the case of our broad money demand equation, due to a negative real growth trend of these variables against the real national income (Keyder 1998: 306).

Also, in the money merket, it is important to ascertain whether the disequilibrium are due to exogeneous money supply shocks that may trigger inflation in the medium-or long-run, or some unanticipated velocity shift that should in principle be accommodated by the monetary authorities (Kontolemis, 2002: 4) in an endogeneous way. Civcir (2000: 1-31), Mutluer and Barlas (2002: 55-75), and Altınkemer (2004) estimate positive unitary income elasticity for Turkish real broad money balances. But our estimation results contrast with their findings, in a way which implies that the cointegrating vectors reject any stationary linear combination of income velocity of money, the inflation rate, the domestic interest rate and the exchange rate similar to the findings of Choudry (1995: 77-91) for the cases of Israeli, Argentina and Mexico. The studies which are mentioned below dealing with the inflationary process in Turkey confirm the result of non-or low- monetization process, especially for the post-1994 period. And with an average and of more consequence unstable annual real growth rate of about 4% in Turkish economy, our results might be an indicator of a stagflationist economic environment. In this framework, an increasing or a sticky trend in inflation may not be necessiated an increase in monetary aggregates, even a decrease because of stagnationist conditions in the real sector of economy with a sticky price-setting framework reflecting construction of a cost-push inflation, that is stagflation, in a way not in accordance with the quantity theory of money, coinciding with some potential structural breaks. In turn, high inflationary environment would deteriorate the real balances in a low monetization process in the economy, thus enable us to estimate a lower than unity maybe negative real income elasticity for money demand relationship, as was in our case.

We additionally test this result by restricting the real income variable to –1, in other words for our demonstration case b(1,2) = -1, or with the same result b(1,1) = -b(1,2) or b(1,1) + b(1,2) = 0 as was in Hoffman and Rasche (1991: 665-674), Choudry (1995: 77-91), Civcir (2000: 1-31) and Mutluer and Barlas (2002: 55-75). That is, we hypothesize that the real money demand function is homogeneous of degree one with respect to the real national income, and estimate an LR statistic with $\chi^2(1) = 6.386075$ (prob. 0.011502) under the null of unitary elasticity. This result confirms the non-positive unitary income elasticity estimation above.

This conditions can lead the monetary variables be endogeneous out of control of monetary authority. In our opinion, if the demand for money had been strongly sensitive to the real national income with an income elasticity of 1 or larger, namely if the money supply in real terms considering the effect of inflation increases proportional to real income in equilibrium money market conditions as was a priori hypothesized in quasi-quantity theoretical approaches, then monetary variables could be considered as one of the main reasons for chronic high inflationary process in Turkey, which also supports the demand-pull effects for the source of inflationary process. But the empirical findings does not support this case (Agénor and Hoffmaister, 1997: 1-38; Erol, 1997: 363-382; Neyaptı, 1998: 25-34; Alper and Üçer, 1998: 7-38; Özmen, 1998: 543-553; Akyürek, 1999: 31-53 ; CBRT, 2002: 68-70; Leigh and Rossi, 2002: 1-18; Koru and Özmen, 2003: 591-597 for different perspectives of Turkish inflation, but also supporting our findings).

Especially, CBRT (2002: 1-79) emphasizes the changing viewpoint of the monetary authority to the inflation phenomenon in Turkey in an endogeneous money creation framework for the post-2001 period. For the empirical studies emphasized above, the evidence in a general way suggests that inflationary inertia, the exchange rate shocks, and the public sector pricing behaviour be the main causes of inflationary process in Turkey supporting the cost-push and public sector inflationary framework resulted from public sector pricing behavior, not the monetary - aggregate demand pull factors. Besides, in the case of an economy with currency substitution as the Turkish economy, monetary volatility increases through the shocks in demand for domestic currency relative to foreign currency in an endogeneous way to the monetary authority. Also our estimation results indicate that the larger the price level of foreign exchange the more demand for broad money including also the foreign exchange based assets.

We can also deduce from these results that the insensitivity characteristics of broad money demand to real income and estimated inflation variable being the most important –alternative- explanatory factor on broad money demand can indicate that the possible reductions in inflation rate can reverse the dolarization process in economy independently than real income. Bahmani and Domaç (2003: 1-26) approaches this case with respect to an inflationary targeting framework such that in an inflationary targeting framework an increasing volatility of exchange rates against the price level may restrict the currency substitution process in Turkish economy.

By using the adjustment coefficients in equation (1), we measure the speed of the short run response to disequilibrium occuring in various equations of endogeneous variables entered in the system, that is, the feedback effects of the lagged disequilibrium in the cointegrating vector onto the variables in the vector autoregression (VAR) system (Sriram, 1999: 20). When we examine the adjustment coefficients in equation (1) above, we see that all the adjustment coefficients have the minus sign indicating an adjustment process of the short run disequilibrium in the cointegration system towards the long run equilibrium, but the only statistically meaningful variables are DLNRM2Y and DLNDOLAR with the coefficients –0.584748 and –0.882169 respectively. In specific for our analysis of broad money demand, the minus coefficient –0.584747 means that lagged real excess money balances induce smaller holdings of current real monetary balances, which can be considered as a quitely speed adjustment process.

Our cointegration analysis may be sensitive to the lag specification specified by the other lag criterions. Thus now we briefly present the normalized cointegrating equations on the variable LNRM2Y with respect to the different lag specifications. By using the criterion SC (namely lag length selected is one), we cannot estimate any cointegrating relationship between the endogeneous variables for both max-eigen and trace tests. By using the criterions AIC, FPE and HQ (namely lag length is eight), we have estimated four cointegrating equations shown below,

LNRM2Y	LNRY	ENFLASYON	BONOFAİZ	LNDOLAR	(6)
-1.00000	+0.075182	-0.205225	-0.098984	+0.137622	
-1.00000	-1.274212	-0.847561	+0.162216	+0.198765	
-1.00000	-0.427314	+4.749669	-4.252381	+0.400279	
-1.000000	-4.507635	+2.064236	-0.107097	+0.492920	

As can be seen above, no income coefficient which has a positive unitary elasticity is estimated by the cointegrating system. Also the first cointegrating vector is reminiscent to which we have estimated above. Additionally we estimate the cointegrating system with 4 lags which is the lag order suggested by criterions LR and AIC if we choose the maximum lag number as 4, and found one cointegrating relationship. In this case, other lag criterions suggest to use the lag order 1. Below is shown the estimation results which are consistent with our findings above,

LNRM2Y	LNRY	ENFLASYON	BONOFAİZ	LNDOLAR	(7)
-1.00000	-0.379258	-0.398505	-0.041372	+0.160047	

So, we conclude that the equation 1 above estimates the unrestricted long run relationship that we are interested in with the lag length five and the rank one, that is, one cointegration relationship.

Now we carry out the long and short run linear restrictions on our cointegrating equation (1) to determine whether all the variables belong to the cointegrated vector by using the LR test for the exclusion of each variable. First, we use the restrictions on the short run adjustment coefficients by using weak exogeneity tests, then on the long run cointegrating variables of our cointegrating relationship. Hendry and Ericsson (1991: 21) expresses that in no case is it legitimate to make variables exogeneous simply by not modeling them. So, here, we try to examine this issue in our cointegration modeling approach of money demand. For the weak exogeneity test, we examine whether the *i-th* row of the short run adjustment matrix is all zero for the null hypothesis of being weak exogeneity. In this case, *i-th* endogeneous variable is said to be exogeneous with respect to the cointegrating vector parameters. If the null hypothesis is not rejected, cointegrating relationship does not feed back onto that variable. Also for the long run cointegrating vector, we apply the restrictions for the long run parameters. Since we have found one cointegrating relationship, the tests are carried out under the assumption of rank 1. Below is the results from these tests for equation (1).

Table 3
Likelihood Ratio Tests

	LR Statistic	Degrees of Freedom	Probability
a(1,1) = 0	7.556151	1	0.005981
a(2,1) = 0	1.694707	1	0.192982
a(3,1) = 0	1.831790	1	0.175916
a(4,1) = 0	0.681698	1	0.409003
a(5,1) = 0	5.662610	1	0.017330
a(1,1)=a(5,1)=0	7.942221	2	0.018852
a(2,1)=a(3,1)=a(4,1) = 0	4.269169	3	0.233827
b(1,1) = 0	14.11310	1	0.000172
b(1,2) = 0	0.544302	1	0.460655
b(1,3) = 0	4.489408	1	0.034105
b(1,4) = 0	1.370716	1	0.241689
b(1,5) = 0	10.69044	1	0.001077
b(1,1)=b(1,3)=b(1,5)=0	28.98214	3	0.000002
b(1,4) = b(1,4) = 0	1.592380	2	0.451044

The phrase a(i,j) = 0 means the i-th endogeneous variable's adjustment coefficient (error correction term) in the j-th cointegrating relation equals zero. And the phrase b(i,j) = 0 means the j-th endogeneous variable in the i-th cointegrating relation is zero. The symbols used here are based on Eviews 4.1. For the restriction tests, we consider the 10% significance level.

Our results obtained above indicate that the variables LNRY, ENFLASYON and BONOFAİZ are weakly exogeneous. We also tested the exclusion of these variables' adjustment coefficients together and found the same result. The exclusion of the adjustment coefficients of the variables LNRM2Y and LNDOLAR from the vector error correction specification are rejected for both single variables and together. Thus, we determine that a short run vector error correction model should be estimated on the variables LNRM2Y and LNDOLAR by considering the other variables as weakly exogeneous. For the long run cointegrating equation restrictions, we estimate that there is a co-movement between the variables LNRM2Y, ENFLASYON and LNDOLAR, thus considering this variables in a co-movement within the long run analysis. Both the single variables and together are not rejected for being in the long run restricted cointegration vector. Inversely, the variables LNRY and BONOFAİZ seem not to be in the long run relationship. We now have restricted short and long run equations, after applying the tests suggested by contemporaneous econometric theory, and can go on the way of our main economically interest area of the demand for broad money balances. The estimation results are below,

Restricted cointegrating coefficients (std.err. in parantheses) (8)

LNRM2Y	LNRY	ENFLASYON	BONOFAİZ	LNDOLAR
-1.00000	0.000000	-0.410321	0.000000	+0.134197
(0.00000)	(0.00000)	(0.09571)	(0.00000)	(0.00427)

Adjustement coefficients (std.err. in parantheses)

D(LNRM2Y)	D(LNRY)	D(ENFLASYON)	D(BONOFAİZ)	D(LNDOLAR)
-0.652838	0.000000	0.000000	0.000000	-0.593314
(0.14651)	(0.00000)	(0.00000)	(0.00000)	(0.26751)

As is seen in equation (8), in the long run the inflation rate and the exchange rate are the main factors affecting the real broad money demand. As in accordance with our a priori expectations, any increase in inflation rate would decrease the demand for money in the long run, and an increase in exchange rate means also an increase of the demand for real M2Y. Below we present the graph of cointegrating relationship which can be seen as stationary except the period post-2001 which may mean a possible breakpoint through indicating a drift.

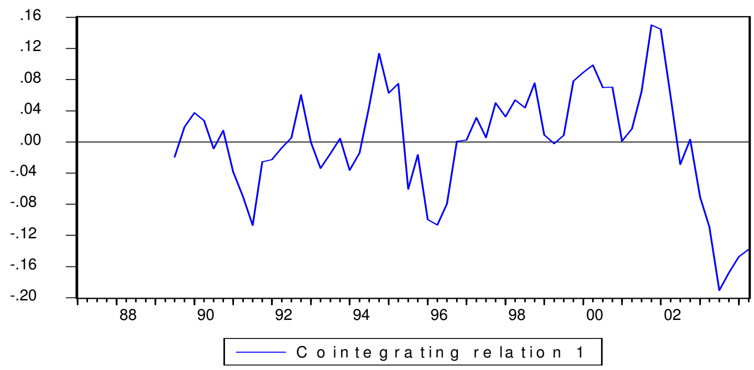

Figure 2

Having established the long run cointegration equilibrium model, we now estimate the vector error correction (VEC) model in light of the weak exogeneity test results on the variable DLNRM2Y by using a reduced form model with the econometrically meaningful variables shown below and the estimated error correction term produced in the cointegration relationship (5). We neglect the VEC specification on the variable DLNDOLAR to save space. *, ** and *** indicate significance at 1%, %5 and 10% respectively.

Table 4
Error Correction Model On Money Demand

Dependent Variable : DLNRM2Y
Method : Least Squares
Sample (adjusted) : 1989.4 – 2004.2
Included observations : 59 after adjusting endpoints
White Heteroskedasticity-Consistent St.Err. & Covar.

Variable*	Coefficient (St.Err.)
C	+0.029875 (0.013141)**
EC(-1)	-0.218210 (0.087569)**
DLNRM2Y(-1)	-0.214191 (0.097162)**
DLNRM2Y(-5)	-0.288679 (0.119436)**
DENFLASYON(-5)	-0.206250 (0.054996)*
DLNDOLAR(-1)	+0.168794(0.031361)*
DLNDOLAR(-3)	-0.097908 (0.056512)***
DLNDOLAR(-5)	+0.382058 (0.072205)*
DUMMY1	-0.043483 (0.038463)
DUMMY2	+0.041780 (0.017810)**

R^2 = 0.535527
Adj. R^2 = 0.450216
S.E. of regression = 0.035209
Durbin-Watson stat. = 2.996783
F-statistic = 6.277323 (0.000007)

For the error correction model, the error correction coefficient EC(-1) is statistically significant and has a minus sign indicating an adjustment process of the short run disequilibrium in the model towards the long run equilibrium process, consistent with the adjustment coefficient of the variable LNRM2Y in equation 5 above. For DLNRM2Y model, the main determinant of real broad money demand in the short run is the exchange rate as expected. The net effect of the exchange rate on broad money demand is positive, while the inflation variable has a negative effect on money demand as was in the long run, even though only one lag of inflation variable is included in the model with statistically significant coefficient. The error correction term representing excess money and the autoregressive terms indicate the portfolio adjustment process in our money demand equation in a such way that a 22% deviation from the long run behaviour of the demand for money is corrected within one period. In this way, as Sriram (1999: 1-49) and Civcir (2000: 1-31) express, a fall in excess money balances in the last period would 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. Additionally, the dummy variable DUMMY2 confirms an increasing trend for the demand for real M2Y in the period of 2000-2001 economic crises.

We now present the statistical diagnostic test results for our estimation process in Table 4, by using Breusch-Godfrey serial correlation LM test, ARCH (autoregressive conditional heteroskedasticity) and White heteroskedasticity test, Jarque-Bera normality test, Chow breakpoint and Chow forecast tests, the CUSUM (cumulative sum) of squares, recursive residuals, one-step forecast and recursive coefficients tests.

Table 5
Diagnostic Tests

Model : DLNRM2Y

Breusch-Godfrey Serial Correlation LM Test (Probs. in parantheses)

Lag : 4	F-statistic	0.57 (0.68)		
	Obs*R-squared	2.86 (0.58)		
ARCH Test				
Lag : 4	F-statistic	0.81 (0.53)		
	Obs*R-squared	3.33 (0.50)		
White Heteroskedasticity Test (with no cross terms)				
	F-statistic	1.12 (0.37)		
	Obs*R-squared	17.61(0.35)		
Normality Test				
Jarque-Bera statistic		2.03 (0.36)		
Chow Breakpoint Test (dummies are excluded) : 1994.2				
	F-statistic	0.696169	Prob.	0.692813
	Log lihelihood ratio	7.185771	Prob.	0.516729
Chow Breakpoint Test (dummies are excluded) : 2000.1				
	F-statistic	0.894682	Prob.	0.529294
	Log likelihood ratio	9.084063	Prob.	0.335254
Chow Breakpoint Test (dummies are excluded) : 2001.2				
	F-statistic	1.108374	Prob.	0.376686
	Log likelihood ratio	11.061147	Prob.	0.198240
Chow Forecast Test (dummies are excluded) : from 1994.2 to 2004.2				
	F-statistic	2.318640	Prob.	0.077375
	Log likelihood ratio	138.7672	Prob.	0.000000
Chow Forecast Test (dummies are excluded) : from 2000.1 to 2004.2				
	F-statistic	1.484431	Prob.	0.158638
	Log likelihood ratio	34.99617	Prob.	0.009463
Chow Forecast Test (dummies are excluded) : from 2001.2 to 2004.2				
	F-statistic	1.754718	Prob.	0.088502
	Log likelihood ratio	27.74122	Prob.	0.009832

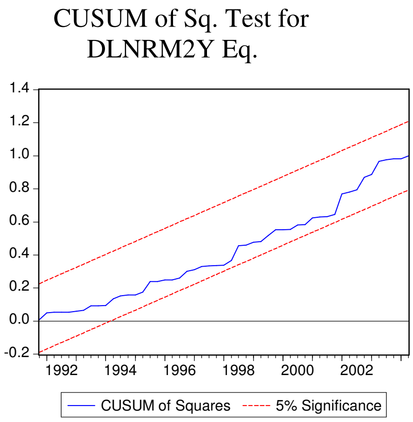

Figure 3

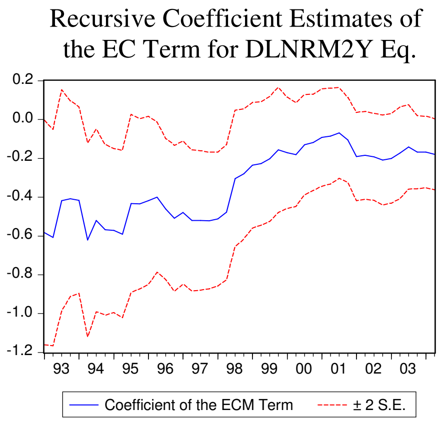

Figure 4

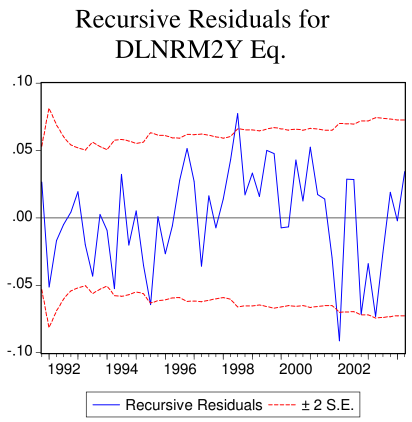

Figure 5

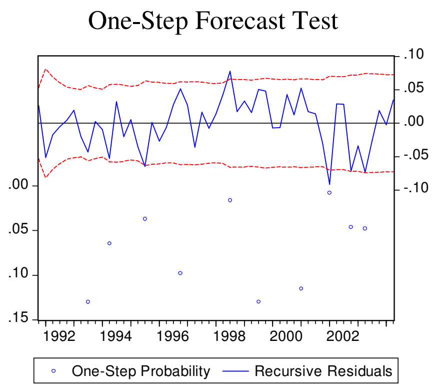

Figure 6

Vector Tests

VEC Residual Serial Correlation LM Test (Probs from chi-square with 25 df.)
Ho : no serial correlation at lag order h

Lags	LM –Stat.	Prob.
4	32.48563	0.1444

VEC Residual Normality Test
Orthogonalization : Cholesky (Lutkepohl)
Ho : residuals are multivariate normal
Jarque-Bera χ^2 (10) = 60.72267 (prob. 0.0000)

VEC Residual Heteroskedasticity Test : No Cross Terms
Ho : No heteroskedasticity or (no misspecification) in the model
Joint test

Chi-sq	819.3605
Degrees of freedom	810
Prob.	0.4020

The diagnostic test results indicate no predicament with aurocorrelation, heteroskedasticity and non-normality problem in a significant way for DLNRM2Y error correction model. Chow breakpoint test results reveal that under the null hypothesis of no structural shift we accept the null hypothesis for all the breakpoints specified. But Chow Forecast tests indicate breakpoints for the post-1994 and post-2001 periods, by considering both F-test and log-likelihood statistics together, contrasting with the findings of Defne and Mutluer (2002: 55-75) and Akıncı (2003: 1-25). We exclude the dummies from equations to be able to keep the periods before 1994 and 2001 in eye. We have also applied the CUSUM (cumulative sum) of squares test to examine the parameter consistency of our EC model. The test gives a plot of the cumulative sum of square residuals together with two critical lines. If the cumulative sum moves outside the region defined by the two critical lines, then the test suggests parameter instability. We have found that the cumulative sum of squares is within the 5% significance lines, suggesting that the residual variance is stable. But inversely the recursive coefficient estimates of the error correction term for DLNRM2Y model indicates instability within the estimation period.

Besides, for recursive residuals, plus and minus two standard errors are plotted about the zero line. Residuals outside the standard error bands suggest instability in the parameters of the equation. For DLNRM2Y error correction model, we can see that major parameter instabilities occur in the second half of 1995 and 1998, in the first half of 1999, in the beginning and at the end of of 2002 in a way coinciding for all these cases with the periods of general elections and increasing political uncertainty. Similarly the first half of 2003 indicates a similar effect, that might reflect the period in which controversies for the role of Turkey with respect to Iraq war increase. The one-step forecast test indicates the probability values for those sample points where the hypothesis of parameter constancy would be rejected at the 5, 10, or 15 percent levels. The points with p-values less than 0.05 correspond to those points where the recursive residuals fall outside the two standard error bands. Also the one-step forecast test results confirm the recursive residual estimates.

Thus, for our broad money demand equation we attribute potential instabilities due to possible breakpoint for M2Y aggregate after 1994, which Mutluer and Barlas emphasize a similar point for the post-1997 period, and a possible breakpoint after 2001 because of a policy regime shift in the Turkish economy, in the manner explained in Özmen (1996: 271-292). Also the political uncertainty conditions are found as reasons creating instability of money demand equation for our case, as was in Akıncı (2003: 1-25).

Finally, we have applied the vector diagnostic tests for our analysis. The results do not indicate any problem of autocorrelated residuals or any heteroskedasticity problem, but non-normality for residuals, no problem in our model through Gonzalo (1994: 203-233).

4. CONCLUDING REMARKS

The broad money consisting of M2Y aggregate is not a policy target for CBRT, but its growing trend can be used to investigate financial liberalization and financial deepening process of Turkish economy, by relating it to other macroeconomic indicators.

In our empirical study, as the main conclusions obtained in a long run perspective, we can express that income elasticity of broad money demand is under unity and maybe it is statistically insignificant in a way not consistent with quantity theory of money. And we attribute this case to the highly unstable growth trend of Turkey. We have estimated inflation phenomenon as the main determinant of broad money for Turkey. The own return of money is estimated in a positive relationship with broad money demand as expected.

For the short run error correction models, we have found higher adjustment coefficient towards long run equilibrium than the adjustment coefficients found in other studies for Turkish money demand. This result can reflect the financial development period of Turkey with an increasing trend in time.

After a stabilization effort against inflation, we can expect for Turkish economy an increase in the domestic money demand which is consistent with the international evidence of money demand for pre-and-post stabilization periods in Stone (1998: 1-41), and a decline in dolarization process in the light of the findings in our study.

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