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DERGİSİ

MULTIRANK COINTEGRATION ANALYSIS OF TURKISH M1 MONEY DEMAND (1987Q1-2006Q3)

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Abstrac

In our paper, we employ multivariate cointegration analysis to the Turkish M1 narrow money demand. The ex-post estimation results reveal that it is possible to identify a money demand vector in the cointegrating space as a priori hypothesized through economics theory. But some structural break points and parameter instabilities coincided with post-1994 economic crisis period and 2000-stabilization program cast some doubt upon whether the estimated model can represent all the period under investigation. Besides, a second potential vector found in the long-run variable space has been decomposed to reconcile it with excess aggregate demand reacting to the domestic inflation.

Keywords:: Money Demand, Aggregate Demand, Turkish Economy, Cointegration, Identification, Super Exogeneity, Structural Breaks, Economic Policy
Jel Classification: E30, E40, E41, E44, E52

Özet

Çalışmamızda Türkiye Ekonomisi koşullarında dar kapsamlı M1 para talebi çok değişkenli eşbütünleşim çözümlemesi kullanılarak incelenmeye çalışılmıştır. Elde edilen tahmin sonuçları iktisat kuramı doğrultusunda oluşturulacak bir para talebi vektörünün deng e eşbütünleşim uzayında tanımlanabileceğini göstermektedir. Ancak 1994 kriz sonrası ve 2000 istikrar programı sonrası tahmin edilen yapısal kırılma ve katsayı istikrarsızlıkları tahmin edilen modelin bütün bir inceleme dönemini temsil edip edemeyeceği ile ilgili olarak bazı kuşkuların oluşmasına yol açmıştır. Ayrıca, uzun dönem değişken uzayı içerisinde tahmin edilen ikinci bir vektör de yurtiçi enflasyona karşı duyarlılık gösteren aşırı toplam talep olgusu ile ilişkilendirilmiştir.

Anahtar Kelimeler: Para Talebi, Derneşik Talep, Türkiye Ekonomisi, Eşbütünleşim, Tanımlama, Süper Dışsallık, Yapısal Kırılmalar, İktisat Politikası
Jel Sınıflaması: E30, E40, E41, E44, E52

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I. INTRODUCTION

A stable demand function for money has long been perceived as a prerequisite for the use of monetary aggregates in 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 characteristics to the other economic aggregates, the monetary authority cannot probably follow an independent monetary policy in order to attain the *ex-ante* specified policy targets. Also if an unstable characteristics for these money balances are 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) 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 lead to errors in the formulation of monetary policy. Beginning by the time of well-known missing money arguments and the stability controversies of the demand for money function of Goldfeld (1973: 577-638) and Goldfeld (1976: 683-730), a great importance has been attributed to this subject in economics literature.[1]

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 as an inventory held for the transactions purposes. Transactions 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

[1] Hafer and Hein (1979: 3-14), Laumas and Spencer (1980: 455-459), Judd and Scadding (1982: 993-1023), Gordon (1984: 403-434) and Hendry and Ericsson (1991: 8-38) touch on the instability and misspecification problems of the theory-based accepted money demand functions.

assets among which people distribute their wealth. People give more importance to the expected rate of return for the assets held relative to the transactions 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. We can thus 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 highlights the new quantity theory and Tobin (1958: 65-86) can mainly be considered as the pioneer 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 narrow money demand for the Turkish economy by constructing an empirical model including a set of explanatory aggregates relating to the money demand function and aim to test it by using modern econometric estimation techniques. Our focus thus inclines more on the transactions motive of the demand for money. Below we first give some literature review for the case of Turkey. The third section interests in data issues and model specification and also estimates an empirical model for the Turkish economy. And the final section concludes.

II. SOME LITERATURE REVIEW ON TURKISH ECONOMY

In a study comparing backward and forward looking appoaches to modelling 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 the sense that the expectations of economic agents catch up the inflation rate extensively and that this case enables them to get rid of inflation tax by reducing their monetary holdings.

Metin (1994: 231-256) and Metin (1995) estimate 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 with a negative inflation elasticity as opportunity cost for the money demand equation.

Koğar (1995a) 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 with an elasticity of exchange rate highly low.

Civcir (2000) 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 point out the existence of a stable real broad money demand relationship with a positive unitary income elasticity confirming the quantity theory of money and with negative opportunity cost variables. He expresses that this case might provide justification for the monetary authority to target broad money by considering the effect of dollarization.

Mutluer and Barlas (2002: 55-75) analyze the Turkish broad money demand including deposits denominated in foreign currency for the period 1987-2001 with quarterly data. Their results also point out the existence of a long run relationship for real broad money in Turkey with a unitary income elasticity estimated as was in Civcir (2000). 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) models the demand for real cash balances in Turkey for the period 1987.1-2003.3 with quarterly data and estimates that there exists a long run relationship between real currency issued, private consumption expenditure as scale-income 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 opportunity cost variables have the expected negative magnitudes.

Altınkemer (2004) investigates the base money demand for Turkey under an assumption of rational expectations and succeeds in estimating a stable long run base money demand function. Also a stable long run M2Y function is estimated. The empirical findings give support to the joint endogeneity of inflation and real base money which do not support the possibility of monetary targeting for Turkey and also which give an indirect support for the alternative targeting regimes specifically for inflation targeting. For policy purposes, however, 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.

III. DATA AND MODEL SPECIFICATION

While investigating the money demand function, a critical point to be considered 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 and critical assumption that the quantity of money supplied and demanded equal each other thus assuming equilibrium in the money market (Laidler, 1973: 85). For transactions purposes, we can suppose that narrowly defined monetary variables are better to be considered, while broadly defined monetary variables are better for the portfolio balance approaches in the money demand equation.

After defining the money demand variable, narrowly or broadly for our purpose, we should choose the explanatory factors affecting this variable. We should first 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 instance, if we mainly interest in the volume of transactions in the economy, the current real national income or a scale variable representing the expenditure-pulled approaches would be appropriate for our aims. The current real gross domestic product or private consumption expenditures in the national income accounts can thus be used for this variable. If our aim is to investigate the portfolio balance approach, the expexted or permanent income variable considering the weighted averages of the subsequent income periods or a wealth variable representing the values of all the tangible assets in the economy would be better off to be considered 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 in for what motives people hold these balances in their hand, we should as a next step determine what alternative costs are current in the economy thus discouraging people to hold these balances. These alternative costs may be the interest rates on bonds, returns of equities, changes in the exchange rate representing currency substitution and also the inflation rate representing the increase of prices of intangible assets under the assumption of substitution between commodities and domestic money. More condensed on portfolio approach, more intruments would be necessary for economic agents to hold in hand. An expectation of an increase of the prices for all these assets would probably decrease the demand for money. We thus expect a negative coefficient for these variables.

We now construct a model of money demand for our empirical purposes for the investigation period of 1987Q1-2006Q3 using quarterly observations. We use a variety of econometric procedures available in the program EViews 5.1. All the data we use are taken from the electronic data delivery system of Central Bank of Republic of Turkey (CBRT) and indicate seasonally unadjusted values except the real income variable. The monetary variable we consider (LNRM1) is the narrowly defined monetary balances in natural logarithms which is the sum of currency in circulation and demand deposits in the banking system. As Mutluer and Barlas (2002: 55-75) express, a narrow definition would be more flexible and reactive to market operations and to interest rate policies of the monetary authority.

For the scale-income variable, we use real gross domestic product (GDP) data in natural logarithms (LNREALGDPSA) 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. We use U.S. Census Bureau's X12 seasonal adjustment program also available within EViews 5.1 to adjust real income variable against seasonality.

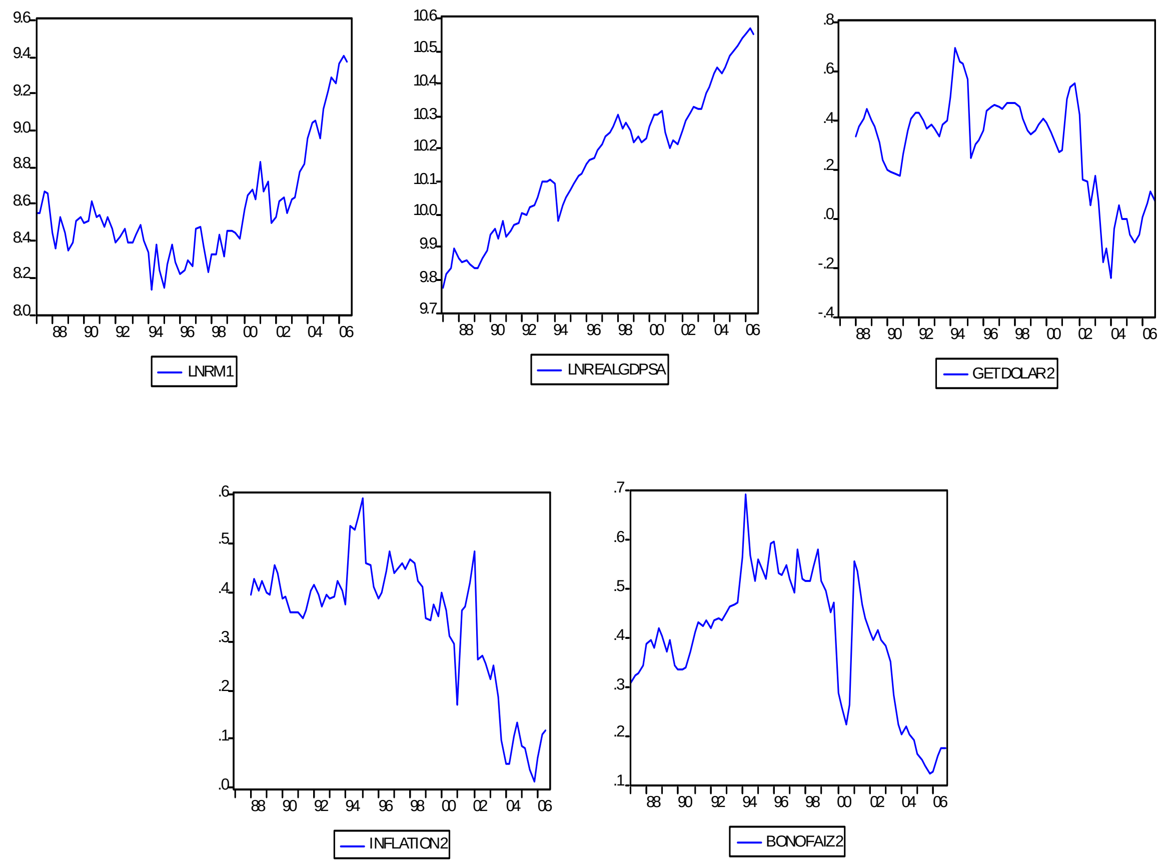
The variables representing alternative cost to hold money in our paper are the maximum rate of interest on the Treasury bills (BONOFAIZ2) whose maturity are at most twelve months or less, the annualized quarterly domestic inflation rate (INFLATION2) based on GDP deflator (DEFL) which is calculated as (DEFL – DEFL(-4)) / DEFL(-4), and the annualized quarterly change in TL / US$ exchange rate (GETDOLAR2).[2] Choudry (1995: 86) expresses that a significant presence of the rate of change of exchange rate in the demand function for real money balances may provide evidence of currency substitution in high inflation countries, which reduces domestic monetary control by also reducing the financing of deficit by means of seigniorage and the base of the inflation tax.[3] He indicates that for three high inflation countries, i.e., Argentia, Israel and Mexico, stationary long run money demand relationship only holds with the inclusion of currency depreciation in the money demand function. Akıncı (2003) also argues that nominal interest rates alone are sufficient in the money demand models for Turkish economy for that 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. We will thus take account of all these cases in our modelling approach of the Turkish money demand. We additionally assume that own rate of return for narrowly defined monetary balances is zero for simplicity. Two impulse-dummy variables which take on values of unity from 1994Q1 till 1994Q4 and from 2001Q1 till

[2] Following Calvo and Leiderman (1992, 1179-194), and Easterly, Mauro and Schmidt-Hebbel (1995, 583-603), if the sequential values are not very close to each other the cost of holding money can be considered such as [ENFLASYON / (1+ENFLASYON)], which fits well to the Turkish case.

[3] Selçuk (1994: 509-518), Akçay, Alper and Karasulu (1997: 827-835), Selçuk (1997: 225-227), Kural (1997: 45-57), Şıklar (1998: 3-14), Selçuk (2001: 41-50), Soydan (2003), and Özdemir and Turner (2004) deal with either the issue of currency substitution or its effect on the degree of inflation tax for the case of Turkish economy.

2001Q4 concerning the financial crises occured in 1994 and 2001 are considered as exogeneous variables. Under the assumption of no money illusion which is quitely reasonable for a cronic-high inflation country, we can suppose that the demand for money is a demand for real money balances. In our case, we use the GDP deflator to deflate the narrow money supply for the 1987:100 based price indices ended by the time of 2004Q4 and since then only the 2003:100 based prices indices can be considered. Below we give in Figure 1 the time series representation of the variables used in this paper,

FIGURE 1: TIME SERIES USED IN THE PAPER



We should specify that, following the modern literature on this issue, we use the variables INFLATION2, GETDOLAR2 and BONOFAIZ2 in the linear form, not in natural logarithms.[4]

[4] However, Hoffman and Rasche (1996: 105-110) criticize using such form of variables expressed in ratios in the level form. Considering the natural log of M1 velocity and the natural log of commercial paper rate, they criticize Friedman and Kuttner (1992: 472-492) using semi-log functional form of interest rates which are used in difference form, and a spread variable between commercial paper and treasury bill rate in the level form, and allege that this form are not robust to the estimated results in Friedman and Kuttner (1992: 472-492), and are completely reversed when the early 1980s are excluded from the sample for the case of U.S. economy for that the large elasticity of velocity in a high interest rate regime implied by the semi-log functional form does not

As a next step for our econometric analysis, we investigate the time series properties of the variables. Granger and Newbold (1974: 111-120) indicate 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) and Phillips-Perron (PP) unit root tests, we check for the stationarity condition of our variables assuming constant and trend terms in the regressions. Thus for the ADF and PP tests, we compare the ADF and PP statistics obtained with the MacKinnon (1996: 601-618) critical values also possible in EViews 5.1, and for the case of stationarity we expect that these statistics are larger than the MacKinnon critical values in absolute value and that they have a minus sign. Although differencing eliminates trend, we report the results of the unit root tests for the first differences of the variables with a linear time trend in the test regression. The results are shown in Table 1 below,[5]

TABLE 1: UNIT ROOT TESTS (assuming constant&trend)

	ADF test	PP test	ADF test	PP test
Variable	(in levels)		(in first differences)	
LNRM1	-1.476514(0)	-1.071925(17)	-10.94212(0)*	-17.15090(26)*
LNREALGDPSA	-2.775424(0)	-3.063836(3)	-8.913043(0)*	-8.913345(1)*
INFLATION2	-1.128250(4)	-2.525271(3)	-7.874700(3)*	-9.853141(3)*
GETDOLAR2	-1.954039(5)	-2.331833(0)	-4.045848(4)**	-7.346675(2)*
BONOFAIZ2	-2.131476(0)	-2.102758(3)	-8.404662(0)*	-8.776833(6)*
Test Critical Values	ADF and PP			
%1 level	-4.081666			
%5 level	-3.469235			

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 using constant&trend terms in the

capture the interest rate trend that dictates the aggregate of agent's long run preferences for liquidity. According to Hoffman and Rasche (1996: 105-110), Friedman and Kuttner's result does not reflect a change in aggregate structure in the 1980s, but the inadequacy of the semi-log functional form to deal with the range of interest rates that were observed in the early 1980s.

[5] For the MacKinnon critical values, we consider %1 and %5 level critical values for the null hypothesis of a unit root. The numbers in parantheses are the lags used for the ADF stationary test and augmented up to a maximum of 10 lags, and we add a number of lags sufficient to remove serial correlation in the residuals, while the Newey-West bandwidths are used for the PP test. The choice of the optimum lag for the ADF test was decided on the basis of minimizing the Schwarz Information Criterion (SC). The test statistics and the critical values are from the ADF or UNITROOT procedures in EViews 5.1. '*' and '**' indicate the rejection of a unit root for the %1 and %5 levels, respectively.

test equation in the level form. But inversely, for the first differences of all the variables the null hypothesis of a unit root is strongly rejected. So we accept that all the 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 with each other. Engle and Granger (1987: 251-276) indicate that even though economic time series may be non-stationary in their level forms, there may exist some linear combination of these variables that converge to a long run relationship over time. 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 (1995) in order to specify the long run relationships between the variables considered making use of EViews 5.1 User's Guide by QMS (2004: 735-748). Let us assume a VAR of order p,

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

where y_t is a k-vector of non-stationarity I(1) variables, x_t is a d-vector of deterministic variables such as constant term, linear trend, seasonal dummies, and crisis variables and ε_t is a vector of innovations, i.e. independent Gaussian variables with mean zero and variance Ω. Such kind of exogeneous variables are often included to take account of short-run shocks to the system, such as policy interventions and shocks or crises which have an important effect on macroeconomic conditions. It is worth noting that including any other dummy or dummy-type variable will affect the underlying distribution of test statistics so that the critical values for these tests are different depending on the number of dummies included (Harris, 1995: 81). We can rewrite this VAR as,

$$\Delta y_t=\Pi y_{t-1}+\sum_{i=1}^{p-1}\Gamma_i\Delta y_{t-i}+Bx_t+\varepsilon_t \qquad (2)$$

where

$$\Pi=\sum_{i=1}^{p}A_i-I \qquad \Gamma_i=-\sum_{j=i+1}^{p}A_j \qquad (3)$$

Granger representation theorem asserts that if the coefficient matrix Π has reduced rank $r<k$, then there exist 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 vector error correction (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 non-normal distributed or when the dynamics are unknown and the model is over-parameterized by including additional lags in the error correction model (Gonzalo, 1994: 225). We thus first determine the lag length of our unrestricted VAR model for which the maximum lag number selected is 5 due to using quarterly frequency data considering five lag order selection criterions, that is, sequential modified LR statistics employing Sims' (1980: 1-48) small sample modification, final predicton error criterion (FPE), Akaike information criterion (AIC), Schwarz information criterion (SC) and Hannan-Quinn information criterion (HQ). As the lag order selected, LR, FPE and AIC statistics suggest 5, and SC and HQ statistics suggest 1 lag orders. We choose the lag order selected by minimized AIC statistics, that is, lag order 5.[6] We also add eleven centered seasonal dummies which sum to zero over a year as exogeneous variable. In this way, the linear term from the dummies disappears and is taken over completely by the constant term, and only the seasonally varying means remain (Johansen, 1995: 84).

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. Following Harris (1995: 87-88) briefly to say, to test the null hypothesis that there are at most r cointegrating vectors and thus k-r unit roots amounts to,

$$H_0: \lambda_i = 0, \qquad i = r+1, \ldots, k \tag{4}$$

[6] For the appropriate lag length to ensure that the residuals are Gaussian, i.e.. they do not suffer from autocorrelation, non-normality, etc., considering the presence of cointegrating relationships, Cheung and Lai (1993: 513-528) find that Monte Carlo experience carried out using data generating processes (DGPs) suggests that tests of cointegration rank are relatively robust to over-parametrizing, while setting too small a value of lag length –such as lag length one or two generally suggested by SC statistics also producing serial correlation problem- severely distorts the size of the maximum likelihood tests (Cheung and Lai, 1993: 319-322: Harris, 1995: 121 footnote 12). Gonzalo (1994: 220-221) also reveals that the cost of over-parametrizing by including more lags in the maximum likelihood based error correction model (ECM) is small in terms of efficiency lost, but this is not the case if the ECM is underparametrized.

where only the first r eigenvalues are non-zero. This restriction can be imposed for different values of r and then the log of the maximised likelihood function for the restricted model is compared to the log of the maximised likelihood function of the unrestricted model and a standard likelihood ratio test computed. That is, it is possible to test the null hypothesis using the trace statistic,

$$\lambda_{trace} = -2\log(Q) = -T\sum_{i=r+1}^{k}\log(1-\hat{\lambda}_i), \qquad r = 0, 1, 2, \ldots, k-2, k-1 \tag{5}$$

where Q = (restricted maximised likelihood / unrestricted maximised likelihood), T is the sample size. Asymptotic critical values are provided in Osterwald-Lenum (1992: 461-472). Another test of the significance of the largest λ_i is the maximal-eigenvalue statistic,

$$\lambda_{max} = -T\log(1-\hat{\lambda}_{r+1}), \qquad r = 0, 1, 2, \ldots, k-2, k-1 \tag{6}$$

which tests that there are r cointegration vectors against the alternative that r+1 exist. Table 2 below reports the results of Johansen Cointegration Test using max-eigen and trace tests based on critical values taken from Osterwald-Lenum (1992: 461-472) and on newer p-values for the rank test statistics from MacKinnon-Haug-Michelis (1999: 563-577) also available from the VAR and COINT procedures in EViews 5.1. For the cointegration test, we restrict intercept and trend factor into our long run variable space, so assume that the trend factor can include the effects of other factors which are not considered in our cointegrating analysis.[7] From the Table 2, both trace and max-eigen statistics indicate jointly 2 potential cointegrating vectors in the long run variable space considering %5 level critical values,

TABLE 2: COINTEGRATION RANK TEST ASSUMING LINEAR DETERMINISTIC

[7] We follow here the *so-called* Pantula principle. Johansen (1992: 383-397) and Harris (1995: 96-97) suggest the need to test the joint hypothesis of both the rank order and the deterministic components, and the former tries to demonstrate how to use the tables in Johansen and Juselius (1990: 169-210) for conducting inference about the cointegration rank. The reason that inference is difficult is that the asymptotic distribution under the null of the test statistic depends on which parameter value is considered under the null. In the case of a cointegration analysis, the limit distribution depends on the actual (true) number of the cointegrating relations and also on the presence of a linear trend. Following Pantula (1989: 256-271), they propose to identify the sub-hypotheses, which give different limit distributions, and construct a test statistic and a critical region for each of these sub-hypotheses. The critical region for the test of the original null hypothesis is then the intersection of the critical regions constructed for each of the sub-hypotheses or, in other words, the hypothesis in question is only rejected if all sub-hypotheses are rejected. Following Harris (1995: 97), the test procedure is to move through from the most restrictive model and at each stage to compare the trace or max-eigen test statistics to its critical value and only stop the first time the null hypothesis is not rejected. However, a critical point to be considered here may be that assuming quadratic deterministic trends in cointegrating space allowing for also linear trends in the short run VEC model may be economically difficult to justify especially if the variables are entered in log-linear form or in linear growth rates, since this would imply an implausible ever-increasing or decreasing rate of change (Harris, 1995: 96).

TREND RESTRICTED IN THE COINTEGRATION EQUATION

Null hypothesis	r=0	r≤1	r≤2	r≤3	r≤4
Eigenvalue	0.476351	0.413529	0.196401	0.148834	0.053516
λ trace	111.4606*	66.82212*	30.00153	14.91431	3.795063
%5 Crirical Value	88.80380	63.87610	42.91525	25.87211	12.51798
Prob.	0.0005	0.0277	0.5021	0.5817	0.7716
λ max	44.63848*	36.82060*	15.08722	11.11924	3.795063
%5 Critical value	38.33101	32.11832	25.82321	19.38704	12.51798
Prob.**	0.0083	0.0123	0.6266	0.5011	0.7716

* denotes rejection of the hypothesis at the 0.05 level.

** MacKinnon-Haug-Michelis (1999) p-values

Unrestricted Cointegrating Coefficients

LNRM1	LNREALGDPSA	GETDOLAR2	INFLATION2	BONOFAIZ2	TREND
19.71932	-19.96111	19.62741	4.799142	15.60260	0.091202
12.53251	31.10716	-11.63061	7.394423	22.53284	-0.372688
-3.829214	-6.988123	-11.09615	68.40042	-40.72605	0.123075
-5.750845	22.44747	18.31567	-34.64220	-4.395738	-0.195232
4.895027	-0.733574	-5.416343	-22.55628	41.40765	-0.083932

Unrestricted Adjustment Coefficients (alpha)

D(LNRM1)	-0.031119	0.006074	-0.004506	-0.001761	0.001222
D(LNREALGDPSA)	-0.002182	-0.005447	-0.005072	-0.004418	0.000491
D(GETDOLAR2)	0.011858	0.015117	-0.000632	0.001190	0.007377
D(INFLATION2)	0.008862	-0.000268	-0.005636	0.005759	0.001334
D(BONOFAIZ2)	0.012900	0.017910	-0.001218	-0.001459	-0.004120

It is not uncommon to find more than one cointegrating relationship in a system with more than two variables using Johansen procedure. Some researchers in this situation revert back to a system with one cointegrating vector by choosing the vector corresponding to the largest eigenvalue or by choosing the most theoretically plausible cointegrating relationship (Baharumshah, 2001: 301). Following here Dickey, Jansen and Thornton (1991: 61-65), the objective of cointegration analysis is to find a k by k matrix β', of rank k, such that $\beta' y_t$ decomposes y_t into its stationary and non-stationary components. This is accomplished by obtaining an r by k sub-matrix of β', $\boldsymbol{\beta}'$, of rank r such that the transformed series $\beta' y_t$ is stationary. The r rows of β' associated with these stationary series are called cointegrating vectors. The remaining k-r unit root combinations are termed "common trends" and are theoretically uncorrelated with the stationary elements in Equation (2) above (Harris, 1995: 87). Let us also consider a model with no common trends, so the system is stationary and variable vector never wanders "too far" from its steady-state equilibrium value. If there is one common trend and k-1 cointegrating vectors, there are k-1 directions where the variance is finite and one direction in which it is infinite. On the other hand, if there is only one cointegrating vector, the system can wander off in k-1 independent directions and it is stable in only one direction. The more cointegrating vectors there are, the more stable would be the system. Hence, all other things the same, it is desirable for an economic system to be stationary in as many directions as possible.

We should specify that in such a situation of multi-rank cointegrating relationship, when interpreting the cointegrating vectors obtained from the Johansen approach, it needs to be stressed that what the reduced rank regression procedure provides is information on how many unique cointegrating vectors span the cointegration space, while any linear combination of the stationary vectors is itself a stationary vector and thus the estimates produced for any particular column in β are not necessarily unique. This can easily be seen by nothing that $\alpha\beta' = \alpha\xi^{-1}\xi\beta' = \alpha^{*}\beta'^{*}$ where ξ is any rxr non-singular matrix. Thus if we can find a ξ matrix that transforms β into β^{*}, we still have the same unique number of cointegration vectors, but the vectors themselves are not unique. This would be a major limitation if we could not determine unique structural relationships for each cointegrating vector (Harris, 1995: 95). Therefore, since the Johansen approach only provides information on the uniqueness of the cointegration space, it will be necessary to impose restrictions motivated by economic arguments (e.g., that some of the β_{ij} are zero, or that homogeneity restrictions are needed such as β_{1j}=-β_{2j}) to obtain unique vectors lying within that space and then test whether the columns of β are identified (Harris, 1995: 104). By taking linear combinations of the unrestricted β vectors, it is always possible to impose r-1 just identifying restrictions and one normalization on each vector without changing the likelihood function (Johansen and Juselius, 1994: 20). If k_i>r-1 restrictions were imposed, we would be applied to the case of over-identifying retrictions, which is the case in our paper as well. EViews 5.1 used in this paper for empirical purposes would report asymptotic standard errors for the estimated cointegrating parameters only if the restrictions identify the whole cointegrating vectors. However, restrictions can be binding even if they are not identifying (QMS, 2004: 731). Thus, identification of the vectors would be achieved if when applying the restrictions of the first vector to the other r-1 vectors, the result is a matrix of rank r-1, that is, a matrix with r-1 independent columns.

For this purpose, we first apply to a unitary income homogeneity restriction and a zero restriction to the inflation coefficient for the first money demand vector normalized on real money balances.[8] Following Harris (1995: 125-138) and Nachega (2001), we decompose the second potential cointegrating vector normalized on real income variable in order to reconcile it with excess aggregate demand reacting to the domestic inflation, restricting other variables to the zero so as to yield an independent economic relationship. Under the assumption of two cointegrating vectors in the variable space, these restrictions would over-identify the whole system. Below is shown the estimation results,[9]

TABLE 3: VECTOR ERROR CORRECTION ESTIMATES

Sample (adjusted): 1989Q3 2006Q3

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

Restrictions identify all cointegrating vectors

LR test for binding restrictions (rank = 2):

Chi-square(3)	6.296002	
Probability	0.098064	
Cointegrating Eq.	CointEq1	CointEq2
LNRM1(-1)	1.000000	0.000000
LNREALGDPSA(-1)	-1.000000	1.000000
GETDOLAR2(-1)	1.123850	0.000000
	(0.25383)	
	[4.42762]	
INFLATION2(-1)	0.000000	-0.287057
		(0.11316)
		[-2.53679]

[8] Otherwise, for the latter restriction, our *ex-post* estimation results reveal that however the coefficients of both inflation and interets rate on bonds would have negative normalized signs on money demand as expected, the coefficient on domestic inflation is statistically insignificant and the VEC system over-identification test rejects the applied null hypothesis of system restrictions using likelihood ratio tests at the %5 significance level. Thus following Akıncı (2003) expressed above, we here take account of only interest rates. These estimation results are also possible upon request.

[9] D(LNRM1) and D(LNREALGDPSA) are the adjustment coefficient of real money balances for the first cointegrating vector and that of real income variable considered for the second cointegrating vector, respectively.

BONOFAIZ2(-1)	0.828772	0.000000
	(0.28833)	
	[2.87434]	
TREND(87Q1)	0.003942	-0.010718
	(0.00072)	(0.00058)
	[5.47443]	[-18.6184]
C	0.785362	-9.626186

We must specify that our restrictions assuming two cointegrating vectors lying in the long run variable space and imposed under the null hypothesis are not rejected using $\chi^2(3)$= 6.296002 with prob. value of 0.098064 against $\chi^2(3)$-table = 7.81473 assuming %5 prob. values.[10] We also succeed in identification of the system in the sense that no equation can be a linear combination of other equation(s) (Harris, 1995: 136). As can be seen above, the adjustment coefficient of real money balances in the first vector and that of domestic real income in the second vector have statistical significance with a negative error correcting coefficient carrying the long run knowledge of cointegrating vectors into the short run dynamic error correcting models. In Table 3, we find that both alternative cost variables, i.e., GETDOLAR2 and BONOFAIZ2, have the *a priori* hypothesized signs and are statistically significant. Besides, removing homogeneity restriction of income elasticity yields a positive unitary coefficient but also gives rise to the rejection of null of over-identifying restrictions in our VEC system using $\chi^2(2)$ = 6.212654 with a prob. value of 0.044765 against $\chi^2(2)$-table = 5.99147. But, this finding should be appreciated cautiously, since the depreciation of narrowly defined monetary aggregates due to chronic high inflationary framework especially for the pre-1998 period coincided with a positive long run growth trend (see Figure 1 above) leads us easily to the conclusion of velocity instability and breaks in the money demand specification inside the examination period. Thus, possible break points inside the period should be searched for, otherwise not only dynamic misspecifications but also an invalid conditioning and a change in the relevant variable space due to a policy regime change and/or financial innovation should be taken as potentially complementary explanations of a money demand instability (Özmen, 1996: 271-292).[11]

[10] Following Harris (1995: 113), we calculate the degrees of freedom using the formula $v = \Sigma_i (n - r + 1 - s_i)$, where n is the number of endogenous variables considered in the cointegrating analysis, r the number of cointegrating vectors and s_i the number of unrestricted parameters in each vector. In our case, (n - r + 1) = (5 - 2 + 1) = 4 and for the first vector v_1 = (4 - 3) = 1, and for the second vector v_2 = (4 – 2) = 2. Thus degrees of freedom for VEC system over-identification test can be estimated such that $v_1 + v_2 = 3$.

[11] See Ardıç (1997) for an overview of well-known Lucas' critic upon this issue.

The second cointegrating vector indicates the long-run inflationary impact of excess aggregate demand and would give us to what extent the trend-adjusted real income is responsive to the domestic inflation structure. If we follow here Harris (1995: 125-138) and Nachega (2001) and give below the normalized relationship on real income,

LNREALGDPSA = 0.287057 INFLATION2 + 0.010718 TREND(87Q1) + 9.626186 (7)

and normalize this vector on domestic inflation rate in our VEC system estimation structure,

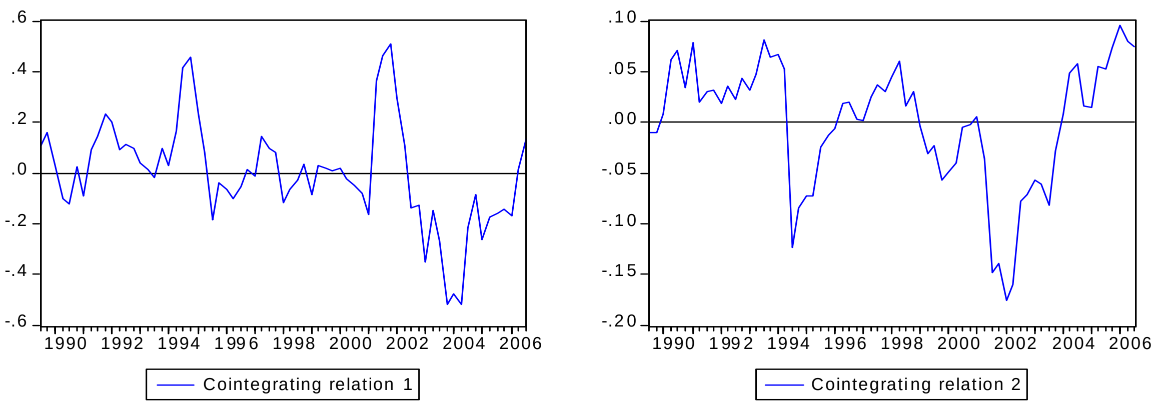
INFLATION2 = 3.483053 LNREALGDPSA - 0.037332 TREND(87Q1) - 33.52846 (8)

and rearrange equation (8),

INFLATION2 = 3.483053(LNREALGDPSA - @TREND(87:01)) - 33.52846 (9)

where the expression 0.010718@TREND(87Q1) is the potential output. Following Nachega (2001), in equation (9) we find that a %1 deviation of actual output from potential output induces on average a %3 increase in the rate of inflation. Below, we present the graphs of cointegrating relations estimated in this paper, which can be seen as stationary,

FIGURE 2: COINTEGRATING RELATIONS



Having established the long run cointegrating equilibrium models, we now estimate the vector error correction (VEC) models on the variables DLNRM1 and DLNREALGDPSA by using a reduced form model with the econometrically meaningful variables shown and the estimated error correction term(s) produced in the cointegrating relationships, which carries the long run knowledge into the short run disequilibrium conditions.[12] Since all the variables

[12] Harris (1995: 134) reports it makes no difference whether y_t in Equation 2 above enters the error correction term with a lag of t-1 or t-i. We estimate that in our money demand parsimonious error correction model, a lag of t-1, t-2, t-3 or t-4 yield negative and significant error correction coefficients, but the larger the lags in error

in the model are now I(0) statistical inference using standard t and F tests is valid (Harris, 1995: 134). We have calculated t-statistics of each variable by dividing relevant coefficient by its standard error. If we consider %0.10 significance level for the relevant coefficient, the probability value lower than 0.10 indicates the significance of that variable as the one which can be accepted econometrically. We also include both an F- and an LR-test for the reduction of insignificant variables in our model. Using QMS (2004: 563), the test is for whether a subset of variables in an equation all have zero coefficients and might thus be deleted from the equation. The F-statistic has an exact finite sample F-distribution under H_0 if the errors are independent and identically distributed normal random variables and the model is linear. Below, at first we give the parsimonious VEC model on money demand,

TABLE 4: PARSIMONIOUS VEC MODEL ON MONEY DEMAND

Redundant Variables: COINTEQ02(-1) D_Q2 D_Q3 DLNRM1(-4) DLNREALGDPSA(-1) DLNREALGDPSA(-3) DLNREALGDPSA(-4) DLNREALGDPSA(-5) DGETDOLAR2(-1) DGETDOLAR2(-2) DGETDOLAR2(-3) DGETDOLAR2(-4) DGETDOLAR2(-5) DINFLATION2(-5) DBONOFAIZ2(-2) DBONOFAIZ2(-4) DBONOFAIZ2(-5)

F-statistic	0.824128	Prob. F(17,35)	0.656939
Log likelihood ratio	22.89422	Probability	0.152726

Test Equation:

Dependent Variable: DLNRM1

Method: Least Squares

Sample: 1989Q4 2006Q3

Included observations: 68

White Heteroskedasticity-Consistent Standard Errors &Covariance

Variable	Coefficient	Std. Error	t-Statistic	Prob.
C	0.052466	0.010882	4.821502	0.0000
COINTEQ01(-1)	-0.504966	0.063862	-7.907150	0.0000
D_Q4	-0.046786	0.017571	-2.662653	0.0103
DUMMY1	-0.094800	0.052768	-1.796561	0.0782
DUMMY2	0.112070	0.023737	4.721338	0.0000
DLNRM1(-1)	-0.432906	0.102493	-4.223750	0.0001
DLNRM1(-2)	-0.331484	0.075795	-4.373420	0.0001

correction term, the lower the speed of adjustment to long run equilibrium. This estimation results not reported here are also available upon request.

DLNRM1(-3)	-0.479137	0.090782	-5.277903	0.0000
DLNRM1(-5)	-0.363432	0.127158	-2.858110	0.0061
DLNREALGDPSA(-2)	-1.045694	0.314307	-3.326978	0.0016
DINFLATION2(-1)	0.838849	0.241690	3.470767	0.0011
DINFLATION2(-2)	0.620710	0.129008	4.811426	0.0000
DINFLATION2(-3)	0.501811	0.166416	3.015408	0.0040
DINFLATION2(-4)	0.637216	0.118772	5.365011	0.0000
DBONOFAIZ2(-1)	-0.397998	0.132772	-2.997601	0.0042
DBONOFAIZ2(-3)	-0.624178	0.101340	-6.159265	0.0000

R-squared	0.786515	Mean dependent var	0.012702
Adjusted R-squared	0.724933	S.D. dependent var	0.098261
S.E. of regression	0.051535	Akaike info criterion	-2.890806
Sum squared resid	0.138102	Schwarz criterion	-2.368569
Log likelihood	114.2874	F-statistic	12.77183
Durbin-Watson stat	2.233750	Prob(F-statistic)	0.000000

Breusch-Godfrey Serial Correlation LM Test

Lag 1	F-statistic	2.069673	Prob. F(1,51)	0.156363
	Obs*R-squared	2.651943	Prob. Chi-Square(1)	0.103423
Lag 4	F-statistic	1.196312	Prob. F(4,48)	0.324586
	Obs*R-squared	6.164543	Prob. Chi-Square(4)	0.187192
Jarque-Bera		1.825944	Prob.	0.401330

Chow Breakpoint Test: 1994Q2 (dummies are excluded)

F-statistic	0.767655	Prob. F(14,40)	0.695669
Log likelihood rat.	16.18240	Prob. Chi-Square(14)	0.302363

Chow Breakpoint Test: 2000Q1 (dummies are excluded)

F-statistic	2.181907	Prob. F(14,40)	0.027120
Log likelihood rat.	28.58289	Prob. Chi-Square(14)	0.000423

Chow Breakpoint Test: 2001Q1 (dummies are excluded)

F-statistic	1.313541	Prob. F(14,40)	0.242467
Log likelihood rat.	25.72154	Prob. Chi-Square(14)	0.028078

Chow Forecast Test: 1994Q2 to 2006Q3 (dummies are excluded)

F-statistic	5.153480	Prob. F(50,4)	0.059611
Log likelihood rat.	284.2948	Prob. Chi-Square(50)	0.000000

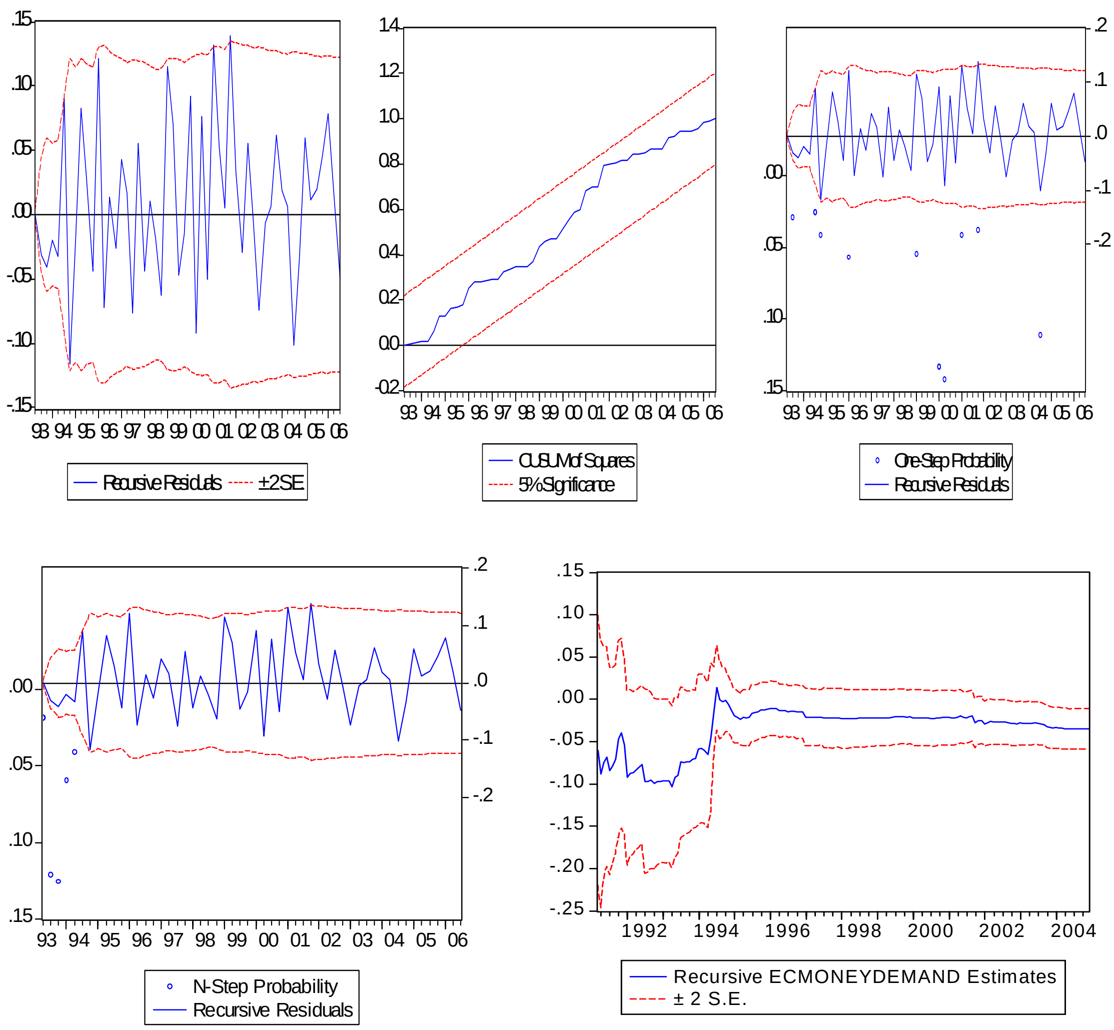
Chow Forecat Test: 2000Q1 to 2006Q3 (dummies are excluded)

F-statistic	1.117216	Prob. F(27,27)	0.387729
Log likelihood rat.	51.00693	Prob. Chi-Square(27)	0.003475

Chow Forecast Test: 2001Q1 to 2006Q3 (dummies are excluded)

F-statistic	0.907962	Prob. F(23,31)	0.589331
Log likelihood rat.	35.02045	Prob. Chi-Square(23)	0.051773

FIGURE 3: RESURSIVE ESTIMATES OF THE VEC MODEL UPON MONEY DEMAND

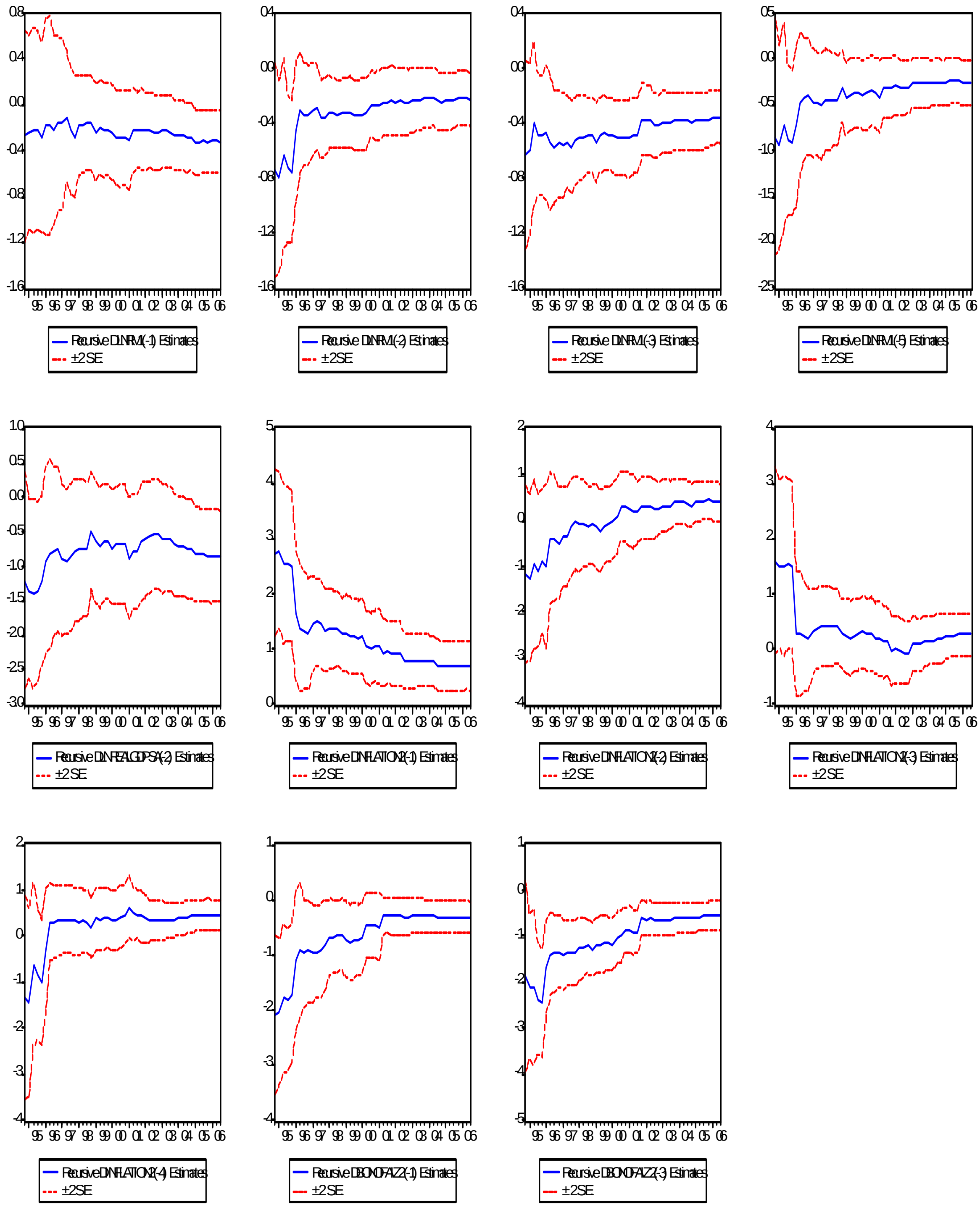


Above, COINTEQ01 and COINTEQ02 are the estimated error correction coefficients upon money demand equation and trend adjusted real income equation conditional on domestic inflation, respectively. Supporting the findings of Harris (1995: 134-136), the money demand error-correction term (cointegration relationsip) is only significant in the first equation, while the excess demand error correction-term is only significant in the second equation. About %50 deviation from the long run path of the real balances is corrected within one period indicating a highly quick adjustment process to long run equilibrium relationship. As Sriram (1999) and Civcir (2000) express, in the case of negative significant error correction term of the money demand equation, 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. Here we focus only upon how properties our VEC coefficient has due to the fact that economics theory rarely interests in the short run characteristics of economic variables but generally pays attention to the long run behavior of those aggregates.

Dealing with the diagnostics, no residual correlation of the 1^{st} or 4^{th} degree or nonnormality problem have been revealed through our error correction model, thus our model seems to have white-noise normally distributed errors. But the Chow tests and the parameter instability tests as a whole point out that some parameter instabilities and possible break points have been occured by the time of economic crisis of 1994 and for the period of 2000 stabilization program and subsequent crisis period. Recursive error correction estimate considering ±2 standard error bands verifies this conclusion such that the major parameter instability for the money demand error correction term occurs, as can easily be noticed, by the time of 1994 economic crisis. We can also attribute such kind of instabilities to the financial innovation period of the Turkish economy inside the period considered as well, since the structural changes may easily raise some questions about the stability of the whole financial system (Koğar, 1995b).

Engle, Hendry and Richard (1983: 284) point out that a conditional model is structurally invariant if all its parameters are invariant for any change in the distribution of the conditioning variables, which guarantee the appropriateness of policy simulations or other control exercises since any change in the distribution of the conditioning variables has no effect on the conditional submodel and therefore on the conditional forecasts of the endogeneous variables. That is, following Metin (1995: 118) and Özmen (1996: 283), super exogeneity implies that the parameters of the conditional model remain constant when the parameters of the marginal process change, i.e., the Lucas' critique does not hold. Examining the issue of parameter instability in the money demand error correction equation in Figure 4 below for all the variables included except the error correction term and the exogeneous centered seasonal dummies and the crisis dummies reveals how structural properties those variables considered have. We can see that for the lags of real money balances, parameter stability seems to be somewhat implemented, maybe except the post-1994 crisis period. But when considering the autoregressive terms of domestic inflation in our parsimonious vector error correction equation, major parameter instabilities thus possible regime changes are revealed. For the lagged domestic interest rate, similar parameter instabilities occur for the pre-2000 period.

FIGURE 4: FIGURE 4: RECURSIVE ANALYSIS OF THE COEFFICIENTS OF MONEY DEMAND EQUATION



We now deal with in Table 5 below the parsimonious error correction model from the second cointegrating model,

TABLE 5: PARSIMONIOUS VEC MODEL ON DLNREALGDPSA

Redundant Variables: COINTEQ01(-1) D_Q2 D_Q3 D_Q4 DLNRM1(-1) DLNRM1(-2) DLNRM1(-3) DLNRM1(-4) DLNRM1(-5) DLNREALGDPSA(-2) DLNREALGDPSA(-3)

DLNREALGDPSA(-5) DGETDOLAR(-1) DGETDOLAR(-2) DGETDOLAR(-3) DGETDOLAR(-4) DGETDOLAR(-5) DINFLATION2(-1) DINFLATION2(-3) DINFLATION2(-4) DINFLATION2(-5) DBONOFAIZ2(-3) DBONOFAIZ2(-4) DBONOFAIZ2(-5)

F-statistic	0.622054	Prob. F(24,35)	0.886648
Log likelihood ratio	24.15766	Prob. Chi-Square(24)	0.452612

Test Equation:

Dependent Variable: DLNREALGDPSA

Method: Least Squares

Sample: 1989Q4 2006Q3

Included observations: 68

White Heteroskedasticity-Consistent Standard Errors &Covariance

Variable	Coefficient	Std. Error	t-Statistic	Prob.
C	0.022560	0.003729	6.050427	0.0000
COINTEQ02(-1)	-0.082537	0.040038	-2.061471	0.0437
DUMMY1	-0.042746	0.024203	-1.766111	0.0826
DUMMY2	-0.053772	0.008569	-6.275115	0.0000
DLNREALGDPSA(-1)	-0.334806	0.155244	-2.156642	0.0351
DLNREALGDPSA(-4)	-0.281297	0.099561	-2.825371	0.0064
DINFLATION2(-2)	0.079859	0.041450	1.926615	0.0588
DBONOFAIZ2(-1)	-0.130709	0.053084	-2.462313	0.0167
DBONOFAIZ2(-2)	0.133006	0.037221	3.573395	0.0007

R-squared	0.504057	Mean dependent var	0.010136
Adjusted R-squared	0.436810	S.D. dependent var	0.029075
S.E. of regression	0.021820	Akaike info criterion	-4.689282
Sum squared resid	0.028090	Schwarz criterion	-4.395523
Log likelihood	168.4356	F-statistic	7.495654
Durbin-Watson stat	1.912630	Prob(F-statistic)	0.000001

Breusch-Godfrey Serial Correlation LM Test

Lag 1	F-statistic	0.190860	Prob. F(1,58)	0.663824
	Obs*R-squared	0.223033	Prob. Chi-Square(1)	0.636738
Lag 4	F-statistic	0.829434	Prob. F(4,55)	0.512221
	Obs*R-squared	3.868568	Prob. Chi-Square(4)	0.424085
Jarque-Bera		10.41549	Prob.	0.005474

Chow Breakpoint Test: 1994Q2 (dummies are excluded)

F-statistic	0.882794	Prob. F(7,54)	0.526311
Log likelihood rat.	7.367716	Prob. Chi-Square(7)	0.391623

Chow Breakpoint Test: 2000Q1 (dummies are excluded)

F-statistic	0.408304	Prob. F(7,54)	0.893046
Log likelihood rat.	3.507118	Prob. Chi-Square(7)	0.834471

Chow Breakpoint Test: 2001Q1 (dummies are excluded)

F-statistic	0.532058	Prob. F(7,54)	0.806391
Log likelihood rat.	4.535336	Prob. Chi-Square(7)	0.716458

Chow Forecast Test: 1994Q2 to 2006Q3 (dummies are excluded)

F-statistic	1.774924	Prob. F(50,11)	0.151170
Log likelihood rat.	284.2948	Prob. Chi-Square(50)	0.000000

Chow Forecat Test: 2000Q1 to 2006Q3 (dummies are excluded)

F-statistic	0.478351	Prob. F(27,34)	0.974045
Log likelihood rat.	21.89511	Prob. Chi-Square(27)	0.742688

Chow Forecast Test: 2001Q1 to 2006Q3 (dummies are excluded)

F-statistic	0.600621	Prob. F(23,38)	0.901206
Log likelihood rat.	21.08541	Prob. Chi-Square(23)	0.575871

In Table 5, we see that estimated error correction coefficient for the second vector is statistically significant with a negative error correction coefficient. About %8 deviation from the long run path of the real income is corrected within one period. As for diagnostics, no serial correlation but nonnormality problem can be noticed. Chow tests reveal that no breakpoint can be found in the model as well. But we do not here go further away, since our main interest area in this paper is the long run structural cointegrating modelling of the same order integrated I(1) variables in a standard money demand variable space.

IV.CONCLUDING REMARKS

In our paper, we have investigated a standard money demand function of M1 real money balances for the Turkish economy. Considering same order integrated I(1) variables, we have constructed a money demand model employing Johansen multivariate cointegration analysis enabling researchers to identify multirank structural economic hypotheses. The *ex-post* estimation results reveal that it is possible to identify a money demand vector as *a priori* hypothesized through economics theory. But some structural break points and parameter instabilities coincided with post-1994 economic crisis period and post-2000 stabilization program cast some doubt upon whether the estimated model can represent all the period under investigation. Besides, a second potential vector found in the long-run variable space has been decomposed in order to reconcile it with excess aggregate demand reacting to the domestic inflation following Harris (1995: 125-138) and Nachega (2001). Using trend-adjusted real income yielded in the second cointegrating vector, we have estimated that a %1 deviation of actual output from potential output induces on average a %3 increase in the rate of inflation.

As a contemporaneous and ever-evolving estimation technique, cointegration analysis can provide many useful insights both in constructing economics theory and in appreciating the *ex-post* outcomes of economic policies implemented by policy makers. In line with such a proposal and considering the methodology applied in this paper, future researchs should be inclusive of more detailed restrictions on dynamic short-run VEC disequilibrium processes in addition to the restrictions on long-run cointegrating relationships. Besides, macroeconometric modelling employing long-run identified structural hypotheses suggested by economic theory, e.g., Garrat et al. (2000: 94-131), Garrat et al. (2003: 412-455), Pesaran, Schuermann and Weiner (2004: 129-181) and also Pesaran and Smith (2006: 24-49), should be of great concern to the researchers interested in such economic issues of estimation in the future papers when constructing economic theories and related policy conclusions.

REFERENCES

Akçay, O. Cevdet, Alper, C. Emre and Karasulu, Meral (1997), "Currency Substitution and Exchange Rate Instability: The Turkish Case", **European Economic Review**, 41, 827-835.

Akıncı, Özge (2003), "Modeling the Demand for Currency Issued in Turkey", **Central Bank Review 1**, 1-25.

Altınkemer, Melike (2004), "Importance of Base Money Even When Inflation Targeting", **CBRT Research Department Working Paper**, No. 04/04.

Ardıç, Kaya (1997), "Lucas Eleştirisi", in Merih Paya (ed.), **Makroiktisat**, Filiz Kitabevi, İstanbul, 330-333.

Baharumshah, Ahmad Zubaidi (2001), "The Effect of Exchange Rate on Bilateral Trade Balance: New Evidence from Malaysia and Thailand", **Asian Economic Journal**, 15/3, 291-312.

Baumol, William J. (1952), "The Transactions Demand for Cash: An Inventory Theoretic Approach", **The Quarterly Journal of Economics**, 66, 1952, 545-556.

Branson, William H. (1989), **Macroeconomic Theory and Policy**, Third Ed., Harper&RowPublishers, New York.

Calvo, Guillermo A. and Leiderman, Leonardo (1992), "Optimal Inflation Tax under Precommitment: Theory and Evidence", **The American Economic Review**, 82/1, 179-194.

Choudhry, Taufiq (1995), "High Inflation Rates and the Long-Run Money Demand Function: Evidence from Cointegration Tests", **Journal of Macroeconomics**, 17/1, Winter, 77-91.

Cheung, Yin-Wong and Lai, Kon S. (1993), "Finite-sample Sizes of Johansen's Likelihood Ratio Tests for Cointegration", **Oxford Bulletin of Economics and Statistics**, 55, 313-28.

Civcir, İrfan (2000), "Broad Money Demand, Financial Liberalization and Currency Substitution in Turkey", **ERF Seventh Annual Conference Proceedings**.

Dickey, David A., Jansen, Dennis W. and Thornton, Daniel L. (1991), "A Primer on Cointegration with An Application to Money and Income", **Federal Reserve Bank of St. Louis Review**, March/April, 58-78.

Easterly, William R., Mauro, Paolo and Schmidt-Hebbel, Klaus (1995), "Money Demand and Seigniorage-Maximizing Inflation", **Journal of Money, Credit and Banking**, 27/2, 583-603.

Engle, Robert F. and Granger, C. W. J. (1987). "Co-integration and Error Correction: Representation, Estimation, and Testing", **Econometrica**, 55, 251-276.

Engle, Robert F., Hendry, David F. and Richard, Jean-Francois (1983), "Exogeneity", **Econometrica**, 51/2, March, 277-304.

Friedman, Milton (1956), "The Quantity Theory of Money - A Restatement", in Milton Friedman (ed.), **Studies in The Quantity Theory of Money**, The University of Chicago Press, 3-21.

Friedman, Milton (1959), "The Demand for Money: Some Theoretical and Empirical Results", **The Journal of Political Economy**, 67/4, August, 327-351.

Friedman, Benjamin M. and Kuttner, Kenneth N. (1992), "Money, Income, Prices and Interest Rates", **The American Economic Review**, 82/3, June, 472-492.

Garrat, Anthony, Lee, Kevin, Pesaran, M. Hashem and Shin, Yongcheol (2000), "A Structural Cointegrating VAR Approach to Macroeconometric Modelling", in Sean Holly and Martin Weale (eds.), **Econometric Modelling: Techniques and Applications**, Cambridge: Cambridge University Press, 94-131.

Garrat, Anthony, Lee Kevin, Pesaran, M. Hashem and Shin, Yongcheol (2003), "A Long Run Structural Macroeconometric Model of the UK", **The Economic Journal**, 113, April, 412-455.

Goldfeld, Stephen M. (1973), "The Demand for Money Revisited", **Brookings Papers on Economic Activity**, 3, 577-638.

Goldfeld, Stephen M. (1976), "The Case of Missing Money", **Brookings Papers on Economic Activity**, 3, 683-730.

Goldfeld, Stephen M. and Sichel, Daniel E. (1990), "The Demand for Money", in B. M. Friedman and F.H. Hahn (eds.), **Handbook of Monetary Economics**, 1, 300-356.

Gonzalo, Jesus (1994), "Five Alternative Methods of Estimating Long-Run Equilibrium Relationships", **Journal of Econometrics**, 60, 203-233.

Gordon, Robert (1984), "The Short-Run Demand for Money: A Reconsideration", **Journal of Money, Credit and Banking**, 16/4, 403-434.

Granger, C.W.J. and Newbold, Paul (1974), "Spurious Regressions in Economics", **Journal of Econometrics**, 2/2, 111-120.

Hafer, R. W. and Hein, Scott (1979), "Evidence on the Temporal Stability of the Demand for Money Relationship in the United States", **FRB of St. Louis Review**, 61/12, 3-14.

Harris, R.I.D. (1995), **Using Cointegration Analysis in Econometric Modelling**, Prentice Hall.

Hendry, David and Ericsson, Neil R. (1991), "An Econometric Analysis of U.K. Money Demand in *Monetary Trends in the United States and United Kingdom* by Milton Friedman and Anna J. Schwartz", **The American Economic Review**, 81/1, 8-38.

Hoffman, Dennis L. and Rasche, Robert H. (1996), **Aggregate Money Demand Functions**, Kluwer Academic

Publishers.

Johansen, Søren (1992), "Determination of Cointegration Rank in the Presence of a Linear Trend", Oxford Bulletin of Cointegration Rank in the Presence of a Linear Trend", **Oxford Bulletin of Economics and Statistics**, 54/3, 383-397.

Johansen, Søren (1995), **Likelihood-based Inference in Cointegrated Vector Autoregressive Models**, Oxford University Press.

Johansen, Søren and Juselius, Katarina (1990), "Maximum Likelihood Estimation and Inference on Cointegration-with applications to the demand for money", **Qxford Bulletin of Economics and Statistics**, 52, 169-210.

Johansen, Søren and Juselius, Katarina (1994), "Identification of the Long-run and the Short-run Structure An Application to the ISLM Model", **Journal of Econometrics**, 63, 7-36.

Judd, John P. and Scadding, John L. (1982), "The Search for a Stable Money Demand Function: A Survey of the Post-1973 Literature", **Journal of Economic Literature**, 20/3, September, 993-1023.

Keyder, Nur (1998), **Para, Teori, Politika, Uygulama**, Gel. 6. Baskı, Beta Dağıtım.

Koğar, Çiğdem İzgi (1995a), "Cointegration Test for Money Demand The Case for Turkey and Israel", **CBRT Research Department Discussion Paper**, No: 9514, May.

Koğar, Çiğdem İzgi (1995b), "Financial Innovation and Monetary Control", **CBRT Research Department Discussion Paper**, No. 9515, May.

Kontolemis, Zenon G. (2002), "Money Demand in The Euro Area: Where Do We Stand (Today)?", **IMF Working Paper 02/185**.

Kural, Vural (1997), "Para İkamesi Altında Enflasyonist Finansman", **Hazine Dergisi**, 5, 45-57.

Laidler, David E.W. (1973), **The Demand for Money Theories and Evidence**, International Textbook Company.

Laumas, G.S. and Spencer, David E. (1980), "The Stability of the Demand for Money: Evidence from the Post-1973 Period", **The Review of Economics and Statistics**, 62/3, 455-459.

MacKinnon, J.G. (1996), "Numerical Distribution Functions for Unit Root and Cointegration Tests", **Journal of Applied Econometrics**, 11, 601-618.

MacKinnon, James G., Alfred A. Haug, and Leo Michelis (1999), "Numerical Distribution Functions of Likelihood Ratio Tests For Cointegration," **Journal of Applied Econometrics**, 14, 563-577.

Metin, Kıvılcım (1994), "Modelling The Demand for Narrow Money in Turkey", **METU Studies in Development**, 21/2, 231-256.

Metin, Kıvılcım (1995), **The Analysis of Inflation: The Case of Turkey (1948-1988)**, Capital Markets Board Publications, No. 20.

Mutluer, Defne and Yasemin Barlas (2002), "Modeling the Turkish Broad Money Demand", **Central Bank Review 2**, 55-75.

Nachega, Jean-Claude (2001), "A Cointegration Analysis of Broad Money Demand in Cameroon", **IMF Working Paper**, No. 01/26.

Osterwald-Lenum, Michael (1992), "A note with Quantiles of the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistics", **Oxford Bulletin of Economics and Statistics**, 54, 461-472.

Özdemir, K. Azim and Turner, Paul (2004), "The Demand for Base Money in Turkey: Implications for Inflation and Seigniorage", **CBRT Research Department Working Paper**, No. 04/12.

Özmen, Erdal (1996), "The Demand for Money Instability", **METU Studies in Development**, 23/2, 271-292.

Pantula, S.G. (1989), "Testing for Unit Roots in Time Series Data", **Econometric Theory**, 5, 256-271.

Pesaran, M. Hashem, Schuermann, Til, Weiner, Scott M. (2004), "Modelling Regional Interdependencies Using a Global Error-Correcting Macroeconometric Model", **Journal of Business and Economic Statistics**, 22/2, April, 129-181.

Pesaran, M. Hashem and Smith, Ron (2006), "Macroeconometric Modelling with a Global Perspective", **The Manchester School**, Supplement, 24-49.

QMS (2004), **EViews 5 User's Guide**, April.

Selçuk, Faruk (1994), "Currency Substitution in Turkey", **Applied Economics**, 26, 509-518.

Selçuk, Faruk (1997), "GMM Estimation of Currency Substitution in a High-Inflation Economy: Evidence from Turkey", **Applied Economics Letters**, 4, 225-227.

Selçuk, Faruk (2001), "Seigniorage, Currency Substitution and Inflation in Turkey", **Russian and East European Finance and Trade**, 37/6, 41-50.

Sims, Christopher A. (1980), "Macroeconomics and Reality," **Econometrica**, 48/1, January, 1-48.

Soydan, Aylin (2003), "Financial Liberalization, Currency Substitution and Seigniorage Evidence from Turkey", **Paper Presented to the Conference on Policy Modeling**, July 3-5, Istanbul.

Sriram, Subramanian S. (1999), "Demand for M2 in an Emerging – Market Economy: An Error Correction Model for Malaysia", **IMF Working Paper 99/17**.

Şıklar, İlyas (1998), "Currency Substitution and Seigniorage Revenue in a Developing Country: The Turkish Case", **Yapı Kredi Economic Review**, 9/1, 3-14.

Tobin, James (1956), "The Interest-Elasticiy of Transactions Demand for Cash", **The Review of Economics and Statistics**, 38/3, August, 241-247.

Tobin, James (1958), "Liquidity Preference as Behavior Towards Risk", **The Review of Economic Studies**, 25/2, February, 65-86.

Yavan, Zafer A. (1993), "Geriye Dönük Modelleme / Çoklu Koentegrasyon ve İleriye Dönük Modelleme Yaklaşımları Çerçevesinde Türkiye'de Para Talebi", **ODTÜ Gelişme Dergisi**, 20/3, 381-416.