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ABSTRACT

In this paper, money demand models using narrowly- and broadly-defined monetary aggregates have been tried to be constructed for the Turkish economy. Using some contemporaneous co-integration estimation techniques for the 1987-2007 period with quarterly data, our findings indicate that for the narrowly-defined monetary aggregates the unit real income elasticity assumption cannot be rejected, but no such a finding can be obtained for the unit price elasticity assumption. For the broadly-defined monetary aggregates the reverse is true, that is, the unit price elasticity assumption cannot be rejected, but we are unable to give support to the unit real income elasticity. Furthermore, we find that interest rate as an alternative cost to holding money is only statistically significant for the broad money demand equation.

Key words: Money Demand; Prices; Real Income; Homogeneity; Turkish Economy;

JEL Classification: C32; E30; E40; E41; E52;

BİR PARA TALEBİ EŞİTLİĞİNDE REEL GELİR VE FİYATLAR İÇİN BAĞDAŞIKLIĞIN SINANMASI: TÜRKİYE ÖRNEĞİ

ÖZET

Bu çalışmada, dar ve geniş tanımlı parasal büyüklükleri kullanan para talebi modellerinin Türkiye ekonomisi için oluşturulmasına çalışılmıştır. Bazı çağdaş eş-bütünleşim tahmin

yöntemlerinin 1987-2007 dönemi için üçer aylık verilerle kullanılması sonucu elde ettiğimiz bulgular dar-tanımlı parasal büyüklükler için birim reel gelir esnekliği varsayımının reddedilemeyeceğini, fakat böyle bir bulgunun birim fiyat esnekliği varsayımı için elde edilemediğini göstermektedir. Geniş-tanımlı parasal büyüklükler için bu durumun tersinin geçerli olduğu, yani birim fiyat esnekliği varsayımının reddedilemediği, fakat birim reel gelir esnekliğine destek verilemediği gözlenmiştir. Ayrıca, faiz oranı parasal büyüklük tutumuna almaşık bir maliyet şeklinde yalnızca istatistiksel olarak geniş para talebi eşitliği için anlamlı bulunmuştur.

Anahtar Kelimeler: Para Talebi; Fiyatlar; Reel Gelir; Bağdaşıklık; Türkiye Ekonomisi;

JEL Sınıflaması: C32; E30; E40; E41; E52;

INTRODUCTION

The concept of money demand provides researchers with the knowledge of motives that determine holding of monetary balances as well as with the course of the expectations shaped by people under the long-term money market equilibrium conditions. Searchs based on the theoretical underpinnings of the behavioral hypotheses that lead to the demand for monetary balances would serve as a bridge relating empirical regularities extracted from the actual data to the identification of the correct design of the monetary policies so that planned characteristics of the monetary policy can be fitted with the end-of-period properties of the policy implementations. On this point, it is convenient to assume that policy makers and especially the monetary authorities are likely to be reached to the stabilization purposes provided that they could estimate both the long- and the short-term course of the monetary aggregates under their control, at least to some extent, and only when they succeed in achieving this task will the policy outcomes reflect the desired consequences as for the stabilization purposes. Thus, complementary to this assumption must be to estimate the true data generating process of the aggregate money demand relationships consistent with *a priori* economic fundamentals.

To start with, investigating the money demand function requires a critical point to be considered as the identification problem which means non-observability of the demand for monetary balances (Laidler, 1993). As is generally hypothesized, researchers make on this point an important assumption that the quantity of money supplied and demanded equal each other which means that money market tends to be converged to the steady-state in the long-term. Then, it is crucial to determine the motives that can be attributed to the behavioral assumptions leading economic agents to demand for money. Considering also applicability for empirical purposes, in this sense, two main approaches can be taken account to explain why economic agents hold these balances, that is to say, the *transactions* and the *assets* or *portfolio balance* approaches. For the *transactions* motives, money is mainly emphasized as a medium of exchange and the demand for monetary balances is assumed to increase proportionally with the volume of transactions in the economy. In this approach, the narrower the definition of money the more likely for the monetary aggregate to be under the control of the monetary authorities. Thus we can additionally assume here that such a variable choice is appropriate for especially policy purposes and would be a measure of the extent to which monetary and interest rate policies respond to changes in general economic outlook. Based on the seminal papers of Baumol (1952) and Tobin (1956), we can state that money is viewed essentially as an inventory held for the transactions purposes, and the 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). On the other side, the *assets* approaches to demand for money imply that economic agents hold money as a store of value and are most likely to consider relative expected returns of assets they hold and thus would take account of the risk factor for these assets because of the probable changing ratio of returns against each other, together with a longer time period in constructing their expectations. Tobin (1958) and Friedman (1959) can be considered some main pioneering papers for the theoretical bases and the empirical applications of this approach in the economics literature.[1]

Following the identification of the main motives of demand for money, it is essential to determine appropriate scale-income and alternative cost variables for the money demand relationship. The scale-income variable can be viewed as representing the extent of the maximum amount of money balances to be held in hand. As Metin (1995) states, the disequilibria between real income and money balances would be able to affect the current demand through the inverse of the monetary velocity. If real income elasticity is found equal to *unity* in a long-term stationary relationship, this case will give support to the *quantity theortical approaches* that assume a strong proportional relationship between real income and monetary balances to provide a stationary income velocity of monetary aggregates. If real income elasticity takes values between *one-half* and *unity*, such a finding will be consistent with the economies of scale argument put forward in the context of the *inventory-theoretic transactions models* of Baumol (1952) and Tobin (1956).[2] On the other hand, if real income elasticity is found significantly above unity, which will indicate an increasing ongoing monetization process in the economy, demand for money balances can be considered like a demand for luxury goods, which will be expected to result in declining monetary velocity.

Having specified the role of scale-income variable in a money demand functional relationship, we have to choose the appropriate alternative cost variable which works as a factor that discourages people from holding money balances in hand. We assume here that in line with the general acceptance in most empirical papers constructed on money demand relationship, a short- and/or long-term interest rate would be the most appropriate choice to be considered due to the monetary returns it serves people against holding money. Thus we expect that an increase in the opportunity cost would likely to lead to a decrease in holding of monetary balances.[3]

Through the methodological issues of the identification of the money demand phenomenon, in the long-term money market equilibrium conditions where demand for monetary balances (M_d) determined by the behavioral motives of the economic agents is assumed to have been equalized to the actual money supply (M_s), we can specify this relationship in a linear functional form for a given period t as follows:[4]

$$M_{dt} = M_{st} \ (\text{money market equilibrium condition}) \tag{1}$$

$$M_{dt} = f\ (P_t, Y_t, R_t) \tag{2}$$

or more explicitly in a log-linear form with expected signs of the variables:

$$m_{dt} = c_0 + \alpha_t p_t + \beta_t y_t - \gamma_t r_t + \varepsilon_t \tag{3}$$

where P_t is the price variable considered in the money demand equation, Y_t the scale-real income variable, and R_t the interest rate chosen as the alternative cost variable. Lower-case letters denote the natural logarithms of each aggregate except the interest rate in the functional form, and ε_t is assumed to represent a white-noise error term reflecting deviations from the equilibrium conditions.[5]

Of special interest here, for our empirical purposes, is given to the coefficient estimates of the variables. In Eq. 3, if price elasticity of the nominal money demand has been found with a value equal to or near the unity, this means that economic agents have been subject to no money illusion within the period examined leading to that money demand function can be converted to a real money demand equation. This means that doubling the price level will double nominal balances, as well.[6] As briefly expressed above, the elasticity of real income reveals a significant knowledge dealing with the course of the velocity of money in the long-term. In this sense, a unitary coefficient for real income yields a result in favor of a stationary income velocity of money. Finally, we try to appreciate the property of the interest rate elasticity. We must specify that Cooley and LeRoy (1981) try to draw the researcher's attention to the points that focus upon whether or not the interest rate elasticity of money demand has in fact a significant negative coefficient, and through a brief methological discussion, state that "… the *Monetarist* view of the transmission mechanism implies that *Monetarists* are likely to be much more willing than *Keynesians* to reverse their field on the question of the interest rate elasticity if the evidence appears to require it" (Cooley and LeRoy, 1981, p. 830).

In this paper, we try to empirically highlight these issues of interest by constructing money demand models for both narrow and broad definition of the money using data from the Turkish economy. So doing, we aim to test whether the homogeneity of prices and real income can be provided in the money demand equation as well as to examine the statistical significance of the interest rate variable as for the narrowly- and broadly-defined money demand functions. To this end, the contemporaneous time series techniques have been applied to extract the necessary knowledge of the long-term money demand variable space from the data. For this purpose, the next section deals with preliminary data issues and the section 2 describes estimation methodology. The results of the empirical model are presented in the section 3. The last section summarizes results to conclude the paper.

I) PRELIMINARY DATA ISSUES

For empirical purposes, all the data cover the investigation period 1987Q1-2007Q3 with quarterly observations and have been taken from the electronic data delivery system of the Central Bank of the Republic of Turkey (CBRT). The monetary variables used are the currency in circulation (*CC*) as the narrow definition of money and M2 monetary aggregate (*M2*) as the broad definition of money, which is consisted of currency in circulation plus demand and time deposits denominated in the domestic currency in the Turkish banking system. The price variable (*P*) has been represented by the gross domestic product (GDP) deflator. The scale-real income variable (*Y*) for the maximum amount of money balances to be held in hand is the real GDP at constant 1987 prices. Finally the interest rate variable as an alternative cost to holding money balances has been determined as the maximum rate of interest on the Treasury bills (*R*) whose maturity are at most twelve months or less, gathered from the electronic data delivery system of the CBRT. The data take the form of seasonally unadjusted values except the real income variable for which the US Census Bureau's X12 seasonal adjustment program has been used to adjust real income against seasonality. All the series except the interest rate have been converted to their natural logarithms. No exogeneous dummy has been included into the empirical analysis as an additional deterministic variable.

Spurious regression problem analyzed by Granger and Newbold (1974) indicates that using non-stationary time series steadily diverging from long-term mean causes to unreliable correlations within the regression analysis leading to unbounded variance process. This is particularly likely to be happened when the adjusted determination coefficient under the impact of correlated trends is found highly larger than the regression Durbin-Watson statistic which can also be resulted from non-stationary residuals. However, for the mean, variance and covariance of a time series to be constant over time, conditional probability distributions of the series must be invariant with respect to the time, and if only so the conventional procedures of OLS regressions can be applied using a stationary process for the variables.[7] Dickey and Fuller (1981) provide one of the commonly used test methods known as the augmented Dickey-Fuller (ADF) test of detecting whether the time series data are of stationary form. However, in addition to the conventional ADF test widely used in economics literature, Elliot et al. (1996) propose a more powerful modified version of this test in which the data are detrended so that explanatory variables are taken out of data prior to running the test regression. This test is similar to the ADF test, but as suggested by Elliot et al. (1996), has a better performance in terms of small sample size and substantially improved power when an unknown mean or trend is present. The DF^{GLS} tries to substitute the generalized least squares (GLS) detrended data of the series under investigation for the original time series data. Briefly to explain in a formal way, Elliot et al. (1996) define a quasi-difference of a variable X_t that depends on the value α representing the specific point alternative against which we wish to test the null hypothesis:

$$d(X_t \mid \alpha) = \begin{cases} X_t & \text{if } t = 1 \\ X_t - \alpha X_{t-1} & \text{if } t > 1 \end{cases} \tag{4}$$

An OLS regression of the quasi-differenced data $d(X_t \mid \alpha)$ on the quasi-differenced $d(Z_t \mid \alpha)$ yields:

$$d(X_t \mid \alpha) = d(Z_t \mid \alpha)'\delta(\alpha) + \eta_t \tag{5}$$

where Z_t consists of deterministic constant or constant and trend terms and let $\delta(\alpha)$ be the estimated value from an OLS regression. For the value of α, Elliot et al. (1996) consider:

$$\alpha = \begin{cases} 1 - 7/T & \text{if } \mathrm{Z}_t = \{1\} \\ 1 - 13.5/T & \text{if } \mathrm{Z}_t = \{1,\mathrm{t}\} \end{cases} \tag{6}$$

Following these specification issues, generalized least squares (GLS) detrended data X_t^d are:

$$X_t^d \equiv X_t - Z_t'\delta(\alpha) \tag{7}$$

The DF^{GLS} substitutes the GLS detrended X_t^d data for the original X_t data in the ADF equation. While the DF^{GLS} t-ratio follows a Dickey-Fuller distribution in the constant only case, the asymptotic distribution differs when included both a constant and trend. Elliot et al. (1996) simulate the critical values of the test statistic in this latter setting for $T = \{50, 100, 200, \infty\}$. We report below in Tab. 1 results from the DF^{GLS} univariate unit root tests. We find that the null hypothesis of a unit root cannot be rejected for levels of all the variables, but inversely, for the first differences the data turn out to be stationary. The exception here is the first difference of the price deflator data, that is, the quarterly inflation, which yields a result in favor of that the variable *P* seems to have an I(2) process.

However, we know that conventional unit root tests such as the revised and extended versions of the Dickey-Fuller tests tend to be strongly criticized when they have been subject to structural breaks which yield biased estimations. Perron (1989) in his seminal paper on this issue argues that conventional unit root tests used by researchers do not consider that a possible known structural break in the trend function may tend too often not to reject the null hypothesis of a unit root in the time series when in fact the series is stationary around a one time structural break.[8]

Tab. 1 Unit Root Test Results

	in levels		in 1st differences		in 2nd differences	
	τ_c	τ_t	τ_c	τ_t		
CC	-0.90 (5)	-0.75 (4)	-1.97 (3)*	-3.10 (3)*		
M2	-0.12 (3)	-0.85 (3)	-2.14 (2)*	-3.11 (2)*		
P	-1.37 (8)	-2.01 (8)	-1.65 (3)	-2.17 (3)	-9.71 (2)*	-10.77 (2)*
Y	1.67 (0)	-2.74 (0)	-3.84 (1)*	-7.97 (0)*		
R	-1.85 (0)	-2.50 (0)	-2.25 (1)*	-3.14 (1)*		
5% cv	τ_c	-1.95	τ_t	-3.09		

Note that the numbers in parentheses are the lags used for the stationarity tests, which are augmented up to a maximum of 10 lags. The choice of optimum lag for the DF^{GLS} tests was decided on the basis of minimizing the Schwarz information criterion. τ_c and τ_t are the DF^{GLS} test statistics with allowance for only constant and constant&trend tems in the unit root tests, respectively. Asterisks denote that variables are of stationary form.

Following these criticisms, we also apply to the widely-used Zivot and Andrews (1992) (henceforth ZA) test for testing the null hypothesis of a unit root for the first difference of the variable *P*. Briefly to explain, for any given time series y_t, ZA test the equation of the form:

$$y = \mu + y_{t-1} + \varepsilon_t \quad (8)$$

Here the null hypothesis is that the series y_t is integrated without an exogenous structural break against the alternative that the series y_t can be represented by a trend-stationary I(0) process with a breakpoint occurring at some unknown time. The ZA test chooses the breakpoint as the minimum *t*-value on the autoregressive y_t variable, which occurs at time 1 < TB < T leading to $\lambda = TB / T, \ \lambda \in \left| 0.15, 0.85 \right|$, by following the augmented regressions:

Model A:

$$y_t = \mu + \beta t + \theta DU_t(\lambda) + \alpha y_{t-1} + \sum_{j=1}^{k} c_i \Delta y_{t-j} + \varepsilon_t \quad (9)$$

Model B:

$$y_t = \mu + \beta t + \gamma DT^*_t(\lambda) + \alpha y_{t-1} + \sum_{j=1}^{k} c_i \Delta y_{t-j} + \varepsilon_t \quad (10)$$

Model C:

$$y_t = \mu + \beta t + \theta DU_t(\lambda) + \gamma DT^*_t(\lambda) + \alpha y_{t-1} + \sum_{j=1}^{k} c_i \Delta y_{t-j} + \varepsilon_t \quad (11)$$

where DU_t and DT_t are sustained dummy variables capturing a mean shift and a trend shift occuring at the break date respectively, i.e., $DU_t(\lambda) = 1$ if $t > T\lambda$, and 0 otherwise; $DT_t^*(\lambda) = t - T\lambda$ if $t > T\lambda$, and 0 otherwise. Δ is the difference operator, k is the number of lags determined for each possible breakpoint by one of the information criteria and ε_t is assumed to be i.i.d. error term. The ZA method runs a regression for every possible break date sequentially and the time of structural changes is detected based on the most significant t-ratio for α. To test the unit root hypothesis, the smallest t-values are compared with a set of asymptotic critical values estimated by ZA. The results of the ZA test for the first differenced form of the variable P are given below:

Tab. 2 Zivot-Andrews Unit Root Tests for the First-Differenced Price Level

Intercept			Trend			Both		
k	min t	TB	k	min t	TB	k	min t	TB
3	-4.83	94Q2	3	-4.62	97Q1	3	-5.80	94Q2

Note that estimations are carried out with 0.15 trimmed. Lag length is determined by Schwarz Bayesian information criterion. Min t is the minimum t-statistic. 5% critical values –intercept: -4.80 ; trend: -4.42; both: -5.08.

The results indicate that for all three cases of the deterministic components in the ZA equation, the price deflator in its first differenced form turns out to be stationary under an endogenous structural break allowed by the data within the period under investigation. When we look at the dates of the possible structural breaks, we can easily notice that they have been likely to occur in just the time of 1994 economic / financial crisis or in the first quarter of the 1997 which can also be attributed to some changes applied by policy makers in the discretionary monetary policies. Thus, for our empirical purposes in this paper, we assume that all the variables used have been subject to an I(1) process, which allows us to applying multivariate co-integration techniques.

II) ESTIMATION METHODOLOGY

We examine the possible long-term stationary relationships derived from the variable space by applying to the multivariate co-integration methodology proposed by Johansen (1988) and Johansen and Juselius (1990), which constructs an error correction mechanism among the same order integrated variables so that stationary combinations of these variables do not drift apart without bound. Moreover, this technique is superior to the regression-based techniques, e.g. Engle and Granger (1987) two-step methodology, for it enables researchers to capture all the possible stationary relationships lying within the long-run variable space. Let us assume a z_t vector of non-stationary n endogenous variables and model this vector as an unrestricted vector autoregression (VAR) involving up to k-lags of z_t:

$$z_t = \Pi_1 z_{t-1} + \Pi_2 z_{t-2} + \ldots + \Pi_k z_{t-k} + \varepsilon_t \qquad (12)$$

where ε_t is assumed to follow an identically and independently distributed (i.i.d.) process with a zero mean and normally distributed N(0, σ^2) error structure and z is (nx1) and the Π_i is (nxn) matrix of parameters. Gonzalo (1994) indicates that Johansen multivarite co-integration methodology performs better an other estimation methods even when the errors are non-normally distributed. Eq. 12 can be rewritten leading to a vector error correction (VEC) model of the form:

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + \Gamma_2 \Delta z_{t-2} + \ldots + \Gamma_k \Delta z_{t-k+1} + \Pi z_{t-k} + \varepsilon_t \qquad (13)$$

$$\Gamma_i = -I + \Pi_1 + \ldots + \Pi_i \; (i = 1, 2, \ldots, k\text{-}1) \qquad (14)$$

$$\Pi = I - \Pi_1 - \ldots - \Pi_k \qquad (15)$$

Eq. 13 can be arrived by subtracting z_{t-1} from both sides of Eq. 12 and collecting terms on z_{t-1} and then adding $-(\Pi_1 - 1)X_{t-1} + (\Pi_1 - 1)X_{t-1}$. Repeating this process and collecting of terms would yield Eq. 13 (Hafer and Kutan, 1994). This specification of the system of variables carries on the knowledge of both the short- and the long-term adjustment to changes in z_t, via the estimates of Γ_i and Π.[9] Following Harris and Sollis (2003), $\Pi = \alpha\beta'$ where α measures the speed of adjustment coefficient of particular variables to a disturbance in the long-run equilibrium relationship and can be interpreted as a matrix of error correction terms, and β is a matrix of long-run coefficients such that $\beta' z_{t-k}$ embedded in Eq. 13 represents up to (n-1) co-integrating relations in the multivariate model which ensure that z_t converge to their long-run steady-state solutions. To select the appropriate lag length of the unrestricted VAR models, we consider various model selection information criterions, namely sequential modified LR statistics employing small sample modification, the final prediction error (FPE) criterion, the Akaike information criterion (AIC), Schwarz criterion (SC) and Hannan-Quinn (HQ) information criterion. For both narrowly- and broadly-defined money demand equations, the LR, FPE, and AIC statistics suggest to use 4 lag orders while SC and HQ statistics suggest 1 lag to be considered. Therefore, for both models, we tend to choose here 4 lag orders by which autoregressive models are constructed.

III) RESULTS

We now report the results of Johansen co-integration test using max-eigen and trace tests based on critical values taken from Osterwald-Lenum (1992). Following Johansen (1992) and Harris and Sollis (2003), for the co-integration test we restrict into the long-term variable space only an intercept for the broad money demand (MD) model and an intercept & trend for the narrow money demand model in line with the Pantula principle.[10] The rank tests are presented in Tab. 3 and Tab. 4.

From the results in Tab. 3 using narrowly-defined money demand model, trace test indicates no co-integrating vector, whereas max-eigen test supports the existence of one co-integrating vector. For the broadly-defined money demand model in Tab. 4, trace statistics indicate one and max-eigen statistics two potential co-integrating vectors considering 0.05 critical values. Following these findings, therefore, we accept that for both model one potential co-integrating vector is likely to be lying in the long-term variable space.

When we examine the unrestricted co-integrating coefficients for both models, we see that the first vectors with the largest eigenvalue seem to be a theoretically plausible money demand vector.[11] Since the Johansen methodology only gives us the unrestricted coefficients that tend to converge to an econometrically identified stationary relationship in a co-integrating vector, some normalizations are needed to be carried out to give the variables economical meaning. Thus, rewriting the normalized equations for both narrowly- and broadly-defined money demand equations under the assumption of r = 1 yields below (standard errors are given in parentheses):

$$\begin{array}{l} \beta_{CC} z_t = CC - \underset{(0.0360)}{0.6937P} - \underset{(0.2472)}{0.6884Y} + \underset{(0.0615)}{0.0089R} - \underset{(0.0060)}{0.0357TREND} - 0.8829 \end{array} \quad (16)$$

$$\beta_{M2} z_t = M2 - \underset{(0.0287)}{0.9828P} - \underset{(0.3471)}{1.8754Y} + \underset{(0.1394)}{0.3624} + 8.9122 \quad (17)$$

Our estimation results indeed give highly plausible money demand vectors. For both money demand models, price and real income have a statistically significant positive elasticity. Furthermore, the interest rate variable has a negative coefficient, but it is only significant for the broad money demand model. To test price homogeneity, we apply to the unit price elasticity restriction and find $\chi^2(1)$ = 8.0499 (prob. 0.0046) for the *CC* model and $\chi^2(1)$ = 0.2359 (prob. 0.6272) for the *M2* model. When we test real income homogeneity, we obtain $\chi^2(1)$ = 0.9501 (prob. 0.3297) for the CC model and $\chi^2(1)$ = 6.4832 (prob. 0.0108) for the M2

Tab. 3 Co-integration Rank Test for the Narrowly Defined MD Model

Null hypot.	*r*=0	*r*≤1	*r*≤2	r≤3
Eigenvalue	0.3421	0.2014	0.1136	0.0484
λ-trace	62.652	30.414	13.101	3.8182
5% cv	63.876	42.815	25.872	12.518
λ-max	32.238*	17.313	9.2832	3.8182
5% cv	32.118	25.823	19.387	12.518

Unrestricted Co-integrating Coefficients

CC	*P*	*Y*	*R*	*TREND*
-24.829	14.223	17.093	-0.2211	0.8868
14.871	-9.7295	-4.7959	8.0973	-0.7289
8.3259	-2.5667	26.991	-2.0840	-0.9799
6.6828	-4.9536	8.6977	0.4786	-0.3878

Unrestricted Adjustment Coefficients ('D' indicates the difference operator)

D(*CC*)	0.0272	-0.0014	-0.0062	-0.0026
D(*P*)	0.0195	0.0140	0.0005	0.0074
D(*Y*)	-0.0065	0.0040	-0.0057	-0.0010
D(*R*)	0.0176	-0.0529	0.0075	0.0321

Notes: Max-eigen test indicate 1 co-integrating eqn(s) at the 0.05 level. Trace test indicates no co-integration at the 0.05 level. An asterisk denotes rejection of the hypothesis at the 0.05 level.

Tab. 4 Co-integration Rank Test for the Broadly Defined MD Model

Null hypot.	*r=0*	*r≤1*	*r≤2*	*r≤3*
Eigenvalue	0.3104	0.2423	0.0742	0.0145
λ-*trace*	57.782*	28.792	7.1540	1.1402
5% cv	47.856	29.797	15.495	3.8415
λ-max	28.990*	21.638*	6.0138	1.1402
5% cv	27.584	21.132	14.265	3.8415

Unrestricted Co-integrating Coefficients

M2	*P*	*Y*	*R*
11.697	-11.496	-21.937	4.239
-3.2309	3.9521	4.1703	-9.1506
4.9976	-5.9468	7.1578	6.4691
-4.4989	6.0038	-9.3489	-2.2074

Unrestricted Adjustment Coefficients ('D' indicates the difference operator)

D(*M2*)	-0.0168	-0.0088	0.0040	-0.0001
D(*P*)	0.0072	-0.0232	-0.0059	-0.0026
D(*Y*)	0.0016	0.0011	-0.0029	0.0026
D(*R*)	-0.0551	0.0030	-0.0195	-0.0177

Notes: Trace test indicates 1 co-integrating eqn(s) while Max-eigen test indicates 2 co-integration eqn(s) at the 5% level. An asterisk denotes rejection of the hypothesis at the 5% level.

model. If we impose both unit price and real income elasticity restrictions on the co-integrating vectors we find $\chi^2(2)$ = 10.4136 (prob. 0.0055) for the *CC* model and $\chi^2(2)$ = 6.5167 (prob. 0.0385) for the *M2* model. These results indicate that for the narrowly-defined monetary aggregates, represented by currency in circulation, the unit real income assumption cannot be rejected, but no such a finding can be obtained for the unit price elasticity assumption. For the broad monetary aggregates the reverse is true, that is, the unit price elasticity assumption cannot be rejected, but in this case we are unable to support unit real income elasticity. Considering the results in Eq. 17, we can infer that for the broadly-defined balances, money seems to be a luxury good in the eyes of economic agents and this would likely to lead to a decrease in the income velocity of money under the long-term money market equilibrium conditions. Finally, we find that interest rate is only significant for the broad money demand equation. We have also tested such a finding by applying to further zero restrictions using an LR statistic and estimate $\chi^2(1)$ = 0.1115 (prob. 0.9148) for the *CC* model and $\chi^2(1)$ = 10.8546 (prob. 0.0010) for the *M2* model. This means that an alternative costs variable is more appropriate for holding of broad money balances when compared with the demand for narrow money, and the broader the definition of money considered the more likely for economic agents to tend to follow an assets approach considering excess return possibilities. Having estimated the co-integrating models, we give in Tab. 5 and Tab. 6 the relevant adjustment coefficients obtained from normalized vectors on money balances Standard errors are given in parentheses and 'D' indicates the difference operator. We see that 68% deviation from the long-term path of narrow money balances is corrected within one period, while it takes about 5 periods for the broad money balances to fully adjust to their long-term course. Thus we can easily infer that adjustment takes place much more rapidly for narrowly defined balances. When we look at the significance of the adjustment coefficients, for the *CC* model all the variables seem to have an endogenous characteristic. However, the

Tab. 5 Adjustment Coefficients for the *CC* Model

D(*CC*)	D(*P*)	D(*Y*)	D(*R*)
-0.6760	-0.4848	0.1620	-0.4364
(0.1465)	(0.1781)	(0.0722)	(0.6371)

Tab. 6 Adjustment Coefficients for the *M2* Model

D(*M2*)	D(*P*)	D(*Y*)	D(*R*)
-0.1961	0.0837	0.0183	-0.6446
(0.0514)	(0.0865)	(0.0361)	(0.2761)

weak exogeneity of price level and real income cannot be rejected for the *M2* model. Finally, we report in Tab. 7 and Tab. 8 the multivariate statistics for testing stationarity and system VEC serial correlation test results:

Tab. 7 Multivariate Statistics for Testing Stationarity (*CC* Model)

	CC	*P*	*Y*	*R*
$\chi^2(3)$	26.97	26.22	25.97	12.03

Tab. 8 Multivariate Statistics for Testing Stationarity (*M2* Model)

	M2	*P*	*Y*	*R*
$\chi^2(3)$	22.29	22.95	20.23	21.69

As for the non-stationarity of the variables, multivariate statistics for testing stationarity are in line with the univariate unit root test results obtained above in the sense that no variable alone can represent a stationary relationship in the co-integrating vector. Furthermore, we find that the models have good diagnostics and indicate no serial correlation at any order:

Tab. 9 VEC Residual Serial LM Test (*CC* Model)

Null hypothesis: no serial correlation at lag order h

Lags	LM-Stat.	Prob.
1	11.15843	0.7996
2	14.80176	0.5392
3	13.51129	0.6351
4	18.02465	0.3225

Tab. 10 VEC Residual Serial LM Test (*M2* Model)

Null hypothesis: no serial correlation at lag order h

Lags	LM-Stat.	Prob.
1	9.770577	0.8783
2	17.26332	0.3688
3	7.928381	0.9510
4	15.19882	0.5101

CONCLUDING REMARKS

Modeling demand for monetary balances is a useful guide to determine the long-term course of monetary variables and thus examines the main motives leading economic agents to holding money in their hands. In this paper, we try to highlight this issue of interest by estimating some money demand models using both narrowly and broadly defined monetary aggregates for the Turkish economy. So doing, we primarily aim to test homogeneity of prices and real income and the significance of the interest rate as an alternative cost to holding money in the specification of the money demand models. Using contemporaneous multivariate co-integration methodology, our findings indicate that for the narrowly-defined monetary aggregates represented by currency in circulation, the unit real income assumption cannot be rejected, but no such a finding can be obtained for the unit price elasticity assumption. For the broad monetary aggregates the reverse is true, that is, the unit price elasticity assumption cannot be rejected, but unit real income elasticity is rejected by the data. Finally, we find that interest rate is only significant for the broad money demand equation. Of course, these findings require further investigation as a complementary study revealing also dynamic properties of the data upon which money demand models have been constructed. Testing the stability of such functional forms must also be elaborately examined so that more accurate policy inferences derived from a money demand model can be obtained by researchers and policy makers.

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ENDNOTES

[1] Using data from the Turkish economy, Metin (1994) tries to estimate a narrow money demand function with a long-span data, while Civcir (2003) presents a comprehensive empirical paper for the estimation of a real broad money demand equation conditioned upon a large set of assets. Altınkemer (2004) also investigates the properties of both a base and a broad money demand function under an assumption of rational expectations. See Sriram (2001) for a survey of recent international evidence of empirical money demand studies.

[2] On this issue, see e.g. Ozmen (1998) estimating a currency seigniorage model for the Turkish economy.

[3] An additional alternative cost variable can also be considered as the level of or the change in the price of the exchange rate for a high inflation country. Choudhry (1995) emphasizes that a significant presence of the rate of change of exchange rate in the money demand function may provide evidence of currency substitution in high inflation countries. In a recent empirical paper upon the Turkish economy, Bahmani-Oskooee and Karacal (2006) emphasize that stability of demand for money can be affected by the (non-)inclusion of exchange rate variable representing currency substitution phenomenon into the functional relationship. However, in this paper, since our main purpose is to examine the homogeneity property of the real income and prices in a money demand equation, we tend to abide by the conventional money demand consideration, and omit the existence of any price variable for exchange rate

changes in the functional form. Of course, this would be an issue of further interest to be examined in the future papers.

[4] Such a relationship can also be attributed to the determination of a money supply process. Thus, we omit here the methodological differences between the identification of money demand and money supply, and assume that in the long-term money market equilibrium conditions both of them tend to be equal to each other.

[5] For such variables as nominal interest rate and inflation that have already been used as a ratio in a long-term money demand analysis, whether or not any logarithmic transformation is needed has been of a controversial issue of interest in the economics literature. For different approaches briefly stated, see, e.g., Friedman and Kuttner (1992) and Hoffman and Rasche (1996).

[6] In most empirical papers, it is implicitly assumed that people have been subject to no money illusion in the long-term. This assumption in turn leads to the estimation of a real money demand equation. But in this paper, we make no such an assumption and tend to explicitly test whether or not this implicit inference as for the money demand equation can be supported by the actual data. For some papers applying a similar methodology, at least to some extent, see Miller (1991), Thornton (1998) and Funke and Thornton (1999) for the use of nominal balances in the demand equation. Dekle and Pradhan (1997) also estimate both a real and a nominal demand equation to appreciate the results of the money demand function considered.

[7] The discussion of such an issue goes back to the earlier analysis of Jevons (1884) and Yule (1926).

[8] For a more detailed investigation of the Perron methodology, see Göktaş (2005).

[9] In Eq. 13, a vector of constants can be included to allow for the possibility of deterministic drift in the data.

[10] For the broad money demand model, restricting a deterministic trend into the long-term variable space yields no co-integrating vectors. Note also that in our paper, we do not allow for quadratic deterministic trends lying in both the co-integrating space and the dynamic vector error correction models.

[11] Notice also that for the broadly-defined vector for which the max-eigen tests yield a rank test result in favor 2 co-integrating vectors, the second potential vector also seems to be a money demand vector with appropriate theoretical signs of the variables. Following Thornton (1998), we must state that normalized equations of these vectors may represent money demand, money supply or some more complicated implications, and the only way to which we apply here is to appreciate the signs and magnitudes of the variables through our *a priori* economic model construction.

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