MPRA
Munich Personal RePEc Archive

Turkish money demand, revisited: some implications for inflation and currency substitution under structural breaks

Levent, Korap

Istanbul University Institute of Social Sciences, Besim Ömer Paşa Cd. Kaptan-ı Derya Sk. 34452 Beyazıt /ISTANBUL

2007

Online at https://mpra.ub.uni-muenchen.de/19479/
MPRA Paper No. 19479, posted 21 Dec 2009 06:06 UTC

TURKISH MONEY DEMAND, REVISITED: SOME IMPLICATIONS FOR INFLATION AND CURRENCY SUBSTITUTION UNDER STRUCTURAL BREAKS

CEM SAATÇİOĞLU* **LEVENT KORAP****

ABSTRACT

In this paper, a money demand model constructed on currency in circulation is used to determine the appropriate alternative cost to hold monetary balances in the Turkish economy. Our estimation results, using contemporaneous multivariate co-integration methodology, indicate that the most significant alternative cost to demand for money is the depreciation rate of the nominal exchange rate. This brings out the importance of having a currency substitution phenomenon settled in the economy when economic agents make their decisions as to their monetary transactions. Moreover, we find that domestic inflationary framework has been subject to a weakly exogenous characteristic and conclude that the main factors leading to domestic inflation are determined out of the money demand variable space.

***Key words*: money demand, inflation, currency substitution, Turkish economy.**
JEL Classification: C32, E41, E52.

TÜRKİYE'DE PARA TALEBİNİN TEKRAR İNCELENMESİ: YAPISAL KIRILMALAR ALTINDA ENFLASYON VE PARA İKAMESİ İÇİN BAZI ÇIKARSAMALAR

ÖZET

Bu çalışmada, dolaşımdaki para miktarı üzerine kurulmuş olan bir para talebi modeli elde para tutumu için uygun almaşık maliyet unsurunun Türkiye ekonomisi koşullarında belirlenebilmesi amacıyla kullanılmaktadır. Çağdaş çok değişkenli eş-bütünleşim çözümlemesi kullanılarak elde ettiğimiz bulgular para tutumu için en anlamlı almaşık maliyet unsurunun parasal döviz kuru değer kayıpları olduğunu göstermekte ve bu durum ekonomide yerleşik para ikamesi olgusunun ekonomik birimler parasal işlemleri ile ilgili kararlarını alırken taşıdığı önemi ortaya koymaktadır. Ayrıca, bulgularımız yurtiçi enflasyonist yapının zayıf dışsal bir özelliğe sahip olduğunu göstermekte ve yurtiçi enflasyonu belirleyen temel etkenlerin para talebi değişken uzayı dışında belirlendiği sonucuna ulaşılmaktadır.

Anahtar kelimeler: para talebi, enflasyon, para ikamesi,Türkiye ekonomisi.
JEL Sınıflaması: C32, E41, E52.

---

* Cem Saatçioğlu is an Assistant Professor in the Department of Economics at İstanbul University, 34452, Beyazıt , İstanbul, Turkey. E-mail: saatcic@istanbul.edu.tr
** Levent Korap is an Economist. E-mail: korap@e-kolay.net

Research on demand for monetary balances gives economic agents knowledge of how policy makers must direct monetary policy. Inferences derived from money demand equations can be used to reveal prerequisites in applying stabilization programs, and appropriate tools used for this purpose can be chosen to achieve program targets. These will enable policy makers to form policy rules for major economic problems in the economy. For example, implications for the income velocity of money and exogenous/endogenous characteristics of the factors which affect the money demand variable space will yield results for the stability of functional relationships in the monetary markets. If stability of money demand can be indicated, this will also indirectly mean that, in line with a well-known quantity theoretical relationship, variations in the velocity of money can be foreseen and explained by economic agents considering a stationary economic relationship. On the other side, if domestic inflationary framework cannot be indicated as a function of monetary aggregates on which money demand variable space is constructed, for example, due to the weakly exogenous characteristic of inflation, such a conclusion will explicitly contradict the quantity theoretical approaches (Özmen, 2003). In this case, money cannot be considered a forcing variable for inflation, and this means that the main factors leading to inflation are determined out of a money demand variable space and that money demand equations should not be appreciated as price equations (MacKinnon and Milbourne, 1988). Therefore, a policy design process that takes into account all of these policy matters will help researchers extract what motives drive the expectations of economic agents.

Dotsey and Hornstein (2003), in their calibrating model on the US economy, warn that even though money communicates information on some policy aggregates such as aggregate output, it is of limited use for policy makers in the sense that it would be a useful signal in an environment driven by productivity shocks, but using it as a signal would have adverse consequences in the presence of money demand disturbances. They suggest that time variation in the behavior of money demand disturbances would imply time variation in a policymaker's responsiveness to money. Likewise, Estrella and Mishkin (1997) focus on the role of monetary aggregates as information variables considering a monetary policy rule perspective. They show that, for the post-1979 period in the US economy, monetary aggregates represented by either a monetary base or a M2 monetary aggregate fall considerably short of this requirement, and results with German M3 broad money supply measures are hardly more favorable. They conclude that monetary aggregates cannot be used in a straightforward way to signal the stance of monetary policy since they do not seem to provide adequate or consistent information. Thus, the existence of a well-specified and stable relationship between money, income and alternative costs to hold money can be seen as a prerequisite for the use of monetary aggregates in the conduct of monetary policy (Goldfeld and Sichel, 1990). Otherwise, disorderly velocity shocks will lead policy makers to fail in the conduct of monetary policy due to the persistent deviations of the growth of monetary aggregates from estimated values. Beginning with the time of well-known missing money arguments and the stability controversies of the demand for money function (Goldfeld, 1973; Goldfeld, 1976), great importance has been attributed to this subject in the economics literature.

For empirical purposes, two approaches can be related to the behavioral assumptions leading economic agents to demand for money, i.e., the transactions and the asset or portfolio balance approaches. The transactions motive emphasizes mainly money's role as a medium of exchange. For this approach, money is viewed essentially as an inventory held for transaction purposes. The transaction costs of going between money and other liquid financial assets justify holding such inventories even though other assets offer higher yields (Judd and Scadding, 1982). Especially the seminal papers by Baumol (1952) and Tobin (1956) develop the underpinnings of this approach, according to which demand for money balances increases proportionally with the volume of transactions in the economy. On the other side, the portfolio balance approaches consider that people hold money as a store of value and that money is only one of the assets among which people distribute their wealth. People give more importance to the expected rate of return for the assets held relative to the transactions necessities and take into account the risk factor for these assets because of the changing ratio of returns against each other. Friedman (1956) and also Friedman (1959), in an influential empirical study which highlights the new quantity theory, and Tobin (1958) can be considered the main pioneering studies in economics literature emphasizing the importance of risk factor and portfolio decision to demand for money.

Given the importance of a stable money demand relationship, many studies in recent years have been conducted on various country cases by researchers such as Sriram (1999), Nachega (2001), Kontolemis (2002), Ramachandran (2004) and Dreger et al. (2006). On the other side, Metin (1994), Civcir (2000), Civcir (2003), Bahmani-Oskooee and Karacal (2006) and some papers by the Central Bank of the Republic of Turkey (CBRT) researchers such as Mutluer and Barlas (2002), Akıncı (2003) and Altınkemer (2004) try to test the demand for money relationship for the Turkish economy. In our paper, we examine these issues of interest by considering the demand for currency in circulation as a function of real income, domestic inflation and exchange rate depreciation. Such a model specification upon narrowly defined monetary aggregates can help us attain implications for the stability of monetary velocity and reveal the structural breaks incurred by velocity shocks. For this purpose, the next section is devoted to a detailed modeling of the Turkish money demand, and the last section concludes.

MODEL

Preliminary Data Issues

We now construct a money demand model for the investigation period 1987Q1-2007Q2 using quarterly observations. The monetary variable we consider (*m*) is the currency in circulation in natural logarithms. The real gross domestic product (GDP) data at constant 1987 prices are used for the scale-real income variable (*y*). The variables representing alternative cost to hold money are 12-months weighted time deosit rate (*r*), annualized quarterly inflation based on GDP-deflator (*p*) and annualized quarterly change in the TL/US$ exchange rate (*e*). Choudry (1995) emphasizes 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. He indicates that for three high inflation countries, i.e., Argentina, Israel and Mexico, the stationary long-run money demand relationship only holds with the inclusion of currency depreciation in the money demand function. Furthermore, Bahmani-Oskooee and Karacal (2006) reveal that the stability of demand for money would be affected by the non-inclusion of exchange rate variable representing currency substitution in the functional form.

Calvo and Leiderman (1992) and Easterly et al. (1995) state that if sequential values of an economic time series (*x*) for the alternative cost to hold money are not very close to each other, the cost of holding money can be considered such as [$x/(1+x)$], which fits well with the Turkish case. Following such a variable specification, we transform variables (*r*), (*p*) and (*e*) into $r2$ = [$r/(1+r)$], $p2$ = [$p/(1+p)$] and $e2$ = [$e/(1+e)$], respectively. All the data used have been taken from the electronic data delivery system of the CBRT and indicate seasonally unadjusted values. We additionally assume that own rate of return for narrowly defined money balances is zero for simplicity and no impulse-dummy variable representing the exogenous shocks witnessed by the economy is considered in the money demand variable space. We use the GDP deflator to deflate the money supply.

The spurious regression problem analyzed by Granger and Newbold (1974) indicates that using non-stationary time series steadily diverging from long-run mean will produce biased standard errors, which causes unreliable correlations and an unbounded variance process within regression analysis. In this way, the standard OLS regression will produce a good fit and predict statistically significant relationships between variables where none really exists (Mahadeva and Robinson, 2004). This means that variables must be differenced (d) times to obtain a covariance-stationary process. Therefore, the individual time series properties of variables should be elaborately considered. Dickey and Fuller (1979, 1981) provide one of the commonly used test methods, known as augmented Dickey-Fuller (ADF) test, for detecting whether time series is stationary. This can be formulated such that:

$$\Delta y_t = \mu + \beta t + (\rho\text{-}1) y_{t\text{-}1} + \sum_{i=1}^{k} \gamma_i \Delta y_{t\text{-}i} + \varepsilon_t \tag{1}$$

where y_t is the variable of interest and t is a time trend. The k lagged differences are to ensure a white noise error series and the number of lags is determined by a test of significance on the coefficient γ_i. The null hypothesis of the ADF test is the presence of a unit root (ρ=1) against alternative stationary hypothesis. For y_t to be stationary, (ρ-1) should be negative and significantly different from zero. We compare the estimated ADF statistics with the simulated MacKinnon (1991, 1996) critical values. For the case of stationarity, we accept that these statistics must be larger than critical values in absolute value and have a minus sign.

Besides the conventional ADF test in Eq. (1), we also consider DF^{GLS} test of Elliot et al. (1996), which proposes a more powerful modified version of the ADF test. In DF^{GLS} test, 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} substitutes the generalized least squares (GLS) detrended y_t^d data for the original y_t data in Eq. 1 above. The DF^{GLS} t-ratio follows a Dickey-

Fuller distribution in the constant only case, while asymptotic distribution differs when both a constant and trend are included in the test equation.

However, due to the evidence yielded by DeJong et al. (1989), the Dickey-Fuller type tests may have low power against plausible stationary alternatives and the null hypothesis of a unit root tends to be accepted unless there is strong evidence against it. Considering these facts, the ADF tests are supplemented by the tests proposed by Kwiatkowski et al. (1992), known as KPSS tests. KPSS tests are designed to test the null hypothesis of stationarity against the unit root alternative. Yavuz (2004) highlights the properties of the ADF-type and KPSS tests and tries to compare them by using Turkish stock exchange data. We report below in Table 1 results from univariate unit root tests. The numbers in parantheses are the lags used for the ADF and DF^{GLS} stationarity tests, which are augmented up to a maximum of 10 lags, and automatic bandwidth selections for the KPSS test. The choice of optimum lag for the ADF and DF^{GLS} tests was decided on the basis of minimizing the Schwarz information criterion (asterisks denote that variables are of stationary form).

Table 1
Unit Root Tests

Var.	τ_C	τ_T	τ_C^{GLS}	τ_T^{GLS}	$Z(\tau_C)$	$Z(\tau_T)$
m	0.86 (4)	-0.18 (4)	0.80 (4)	-0.86 (4)	0.79 (6)	0.28 (6)
Δm	-3.63 (3)*	-4.12 (3)*	-3.47 (3)*	-3.30 (3)*	0.30*(26)	0.14*(19)
y	-0.08 (8)	-2.21 (8)	0.76 (8)	-1.52 (8)	1.22 (6)	0.22 (6)
Δy	-3.02 (7)*	-4.03 (7)*	-2.68 (7)*	-3.23 (7)*	0.10(12)*	0.05(12)*
$p2$	0.40 (4)	-1.15 (4)	0.48 (4)	-0.86 (4)	0.77 (6)	0.26 (6)
$\Delta p2$	-7.77 (3)*	-8.05 (3)*	-1.57 (6)	-7.60 (3)*	0.09 (1)*	0.03 (1)*
$e2$	-1.30 (5)	-1.95 (5)	-1.31 (5)	-1.82 (5)	0.57 (6)	0.22 (6)
$\Delta e2$	-4.04 (4)*	-4.05 (4)*	-3.19 (4)*	-3.89 (4)*	0.06 (2)*	0.03 (2)*
$r2$	-1.57 (0)	-2.24 (0)	-1.46 (0)	-1.68 (0)	0.48 (6)	0.30 (6)
$\Delta r2$	-7.75 (1)*	-7.97 (1)*	-7.80 (1)*	-8.00 (1)*	0.38 (9)*	0.09 (13)*
5% cv.	-2.90	-3.52	-1.95	-3.11	0.46	0.15

Above, τ_C and τ_T are the ADF test statistics with allowance for only constant and constant&trend tems in the unit root tests respectively, and τ_C^{GLS}, τ_T^{GLS}, $Z(\tau_C)$ and $Z(\tau_T)$ are the relevant DF^{GLS} and KPSS statistics. 'Δ' denotes the first difference operator. Results from the unit root tests reveal that null hypothesis of a unit root cannot be rejected for all the variables in the level form, but inversely, for the first differences the stationary alternative hypothesis cannot be rejected.

Testing Endogenous Breaks in the Unit Root Procedure

The unit root tests applied above indicate that variables can be characterized as a random walk process, which requires differencing to achieve a stationary time series. However, these tests are criticized strongly in the contemporaneous economics literature 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. Contrary to the general evidence of many earlier papers which conclude that the US post-war GNP series can be represented by a unit root process, Perron (1989) finds that if the first oil shock in 1973 is treated as a structural breakpoint in the trend function, then the unit root hypothesis of the US post-war GNP series can be rejected in favor of a trend stationary hypothesis.

Selecting the date of structural break, that is, assuming that time of break is known *a priori*, however, may not be the most efficient methodology. The actual dates of structural breaks may not coincide with dates chosen exogenously. To address this issue, several methodologies including Perron (1990), Zivot and Andrews (1992) and Banerjee, Lumsdaine and Stock (1992) have been suggested to allow for the determination of the date of structural breaks endogenously. Considering these issues, in our paper, we follow first the Zivot and Andrews (henceforth ZA) methodology, allowing the data to indicate breakpoints endogenously rather than imposing a breakpoint from outside the system. We then allow for some extensions of this test by following Clemente et al. (1998).

The ZA methodology as a further development on Perron (1989) methodology can be explained by considering three possible types of structural breaks in a series, i.e., Model A assuming shift in intercept, Model B assuming change in slope and Model C assuming change in both intercept and slope. For any given time series y_t, ZA (1992) test the equation of the form:

$$y = \mu + y_{t-1} + e_t \tag{2}$$

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:

$$y_t = \mu + \beta t + \theta DU_t(\lambda) + \alpha y_{t-1} + \sum_{j=1}^{k} c_j \Delta y_{t-j} + \varepsilon_t \tag{3}$$

Model B:

$$y_t = \mu + \beta t + \gamma DT_t^*(\lambda) + \alpha y_{t-1} + \sum_{j=1}^{k} c_j \Delta y_{t-j} + \varepsilon_t \tag{4}$$

Model C:

$$y_t = \mu + \beta \mathrm{t} + \theta DU_t(\lambda) + \gamma DT_t^*(\lambda) + \alpha y_{t-1} + \sum_{j=1}^{k} c_j \Delta y_{t-j} + \varepsilon_t \tag{5}$$

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. We must note that critical values in the ZA methodology are larger in absolute sense than the conventional ADF critical values since the ZA methodology is not conditional on the prior selection of the breakpoint. Thus, it is more difficult to reject the null hypothesis of a unit root in the ZA test. For the appropriate lag length, we consider the Schwarz's Bayesian information criterion (SBIC)-minimizing value.

In addition, considering the case of multiple breakpoints for an economic time series, Clemente et al. (1998) suggest a unit root test that allows for two changes in the mean of a series under the assumption of either innovational (IO) or additive outliers (AO). For the case where the two breaks belong to the IO, we estimate the following regression:

$$\Delta y_t = \mu + d_1 DTB_{1t} + d_2 DTB_{2t} + \theta_1 DU_{1t} + \theta_2 DU1_{2t} + \alpha y_{t-1} + \sum_{j=1}^{k} c_i \Delta y_{t-i} + \varepsilon_t \qquad (6)$$

where DTB_i (i =1,2) are pulse variables that take the value 1 if $t = TB_i + 1$ and zero otherwise, DU_i are defined as above, and TB_1 and TB_2 are the dates when the shifts in mean occur. Eq. (6) again is sequentially estimated and the unit root hyothesis is tested by obtaining minimal value of the t-statistic for the hypothesis α=0 for all break time combinations. An application of the methodology of Clemente et al. (1998) can be found in a recent paper by Abu-Qarn and Abu-Bader (2007).

For estimation purposes, we used EViews 5.1 for the ADF, DF^{GLS} and KPSS tests, and Stata 9.0 for the ZA and Clemente et al. (1998) unit root tests. When we consider the ZA unit root test in Table 2 above allowing endogenous breaks in the time series used, no change occurs in the non-stationary characteristics of the variables. Table 3 presents the results of Clemente et al. (1998) unit root test considering two shifts in the mean of the series for both the AO and IO cases. Allowing two structural breaks in the mean of the series verify the estimation results found above.

Table 2

Zivot-Andrews Unit Root Test

Variable	Intercept			Trend			Both		
	k	min t	TB	k	min t	TB	k	min t	TB
m	2	-2.70	(2003Q3)	2	-3.67	(2002Q1)	2	-3.65	(2001Q2)
y	0	-2.97	(1998Q4)	0	-3.55	(1998Q2)	0	-3.86	(1997Q1)
p2	0	-4.94	(2001Q2)	0	-4.25	(1996Q4)	0	-4.95	(1993Q2)
e2	0	-4.33	(1993Q3)	0	-4.76	(1993Q3)	0	-4.99	(1993Q1)
r2	0	-3.98	(2001Q1)	0	-4.12	(2001Q2)	0	-4.80	(2001Q1)

[1] Estimation with 0.15 trimmed. Lag length is determined by Schwarz's Bayesian information criterion. min t is the minimum t-statistic calculated.
[2] Critical values – intercept: -5.43 (1%), -4.80(5%); trend: -4.93 (1%), -4.42 (5%); both: -5.57 (1%), -5.08 (5%)

Table 3

Clemente-Montañés-Reyes Unit Root Test with Double Mean Shift

Variable	Additive Outliers		Innovative Outliers	
	min t	Optimal Breakpoints	min t	Optimal Breakpoints
m	-2.56	1999Q2, 2003Q2	-2.27	1999Q3, 2003Q2
y	-2.90	1995Q3, 2001Q2	-3.05	1994Q1, 1999Q2
p2	-2.06	1999Q3, 2002Q4	-2.55	1998Q1, 2001Q4
e2	-4.07	1991Q2, 2004Q4	-3.08	1991Q1, 2003Q1
r2	-4.15	1994Q2, 2001Q1	-4.55	1994Q1, 2001Q1

[1] Estimation with 0.15 trimmed. min t is the minimum t-statistic calculated.
[2] 5% critical values – two breaks: -5.49

Econometric Methodology

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:

$$z_t = \Pi_1 z_{t-1} + \Pi_2 z_{t-2} + \ldots + \Pi_k z_{t-k} + \varepsilon_t \tag{7}$$

where ε_t follows an i.i.d. process and z is ($nx1$) and the Π_i is an (nxn) matrix of parameters. Eq. 7 can be rewritten leading to a vector error correction (VEC) model of the form:

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + \Gamma_2 \Delta z_{t-2} + \ldots + \Gamma_{k-1} \Delta z_{t-k+1} + \Pi z_{t-k} + \varepsilon_t \qquad (8)$$

where:

$$\Gamma_i = -I + \Pi_1 + \ldots + \Pi_\mathrm{i} \;\; (i = 1, 2, \ldots, k\text{-}1) \qquad \text{and} \qquad \Pi = I - \Pi_1 - \Pi_2 - \ldots - \Pi_k \qquad (9)$$

Eq. 8 can be arrived at by subtracting z_{t-1} from both sides of Eq. 7 and collecting terms on z_{t-1} and then adding $-(\Pi_1 - 1)X_{t-1} + (\Pi_1 - 1)X_{t-1}$. Repetition of this process and the collection of terms would yield Eq. 8 (Hafer and Kutan, 1994). This specification of the system of variables carries on the knowledge of both short- and long-run adjustment to changes in z_t, via the estimates of Γ_i and Π. Following Harris (1995), $\Pi = \alpha\beta'$ 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, while β is a matrix of long-run coefficients such that $\beta' z_t$ embedded in Eq. 8 represents up to (n-1) cointegrating relations in the multivariate model which ensure that z_t converge to their long-run steady-state solutions. Note that all terms in Eq. 8 which involve Δz_{t-i} are $I(0)$ while Πz_{t-k} must also be stationary for $\varepsilon_t \sim I(0)$ to be white noise of an $N(0, \sigma_\varepsilon{}^2)$ process.

Dealing with the rank conditions, there alternative cases can be considered. If the rank of Π matrix equals zero, there would be no co-integrating relation between the endogenous variables, which means that there would be no linear combinations of the z_t that are $I(0)$. This requires that Π would be (nxn) matrix of zeros. In this case, a VAR model consisting of a set of variables in first differences could be suggested to examine the variable system. If the Π matrix is of full rank when $r = n$, then all elements in z_t would be stationary in their levels. Another case is the possibility that there exist r co-integrating vectors in $\beta' z_t \sim I(0)$ and (n-r) common stochastic trends when Π has reduced rank, i.e., $0 < r \leq (n\text{-}1)$. That is, first r

columns of β are the linearly independent combinations of the endogenous variables settled in vector z_t which represents stationary relationships. Whereas, the latter (n-r) columns constitute the non-stationary vectors of $I(1)$ common trends, which require also that the last (n-r) columns of α take insignificantly values highly close to zero, impeding the feedback effects of deviations from the long-run stationary equilibrium process. Thus this method is equivalent to testing which columns of α are zero (Harris, 1995). Gonzalo (1994) indicates that this method performs better than other estimation methods even when the errors are non-normal distributed or when the model is over-parameterized by including additional lags in the error correction model. Further, this method does not suffer from problems associated with normalization (Johansen, 1995).

Model Specification

We now construct an unrestricted VAR model consisting of an endogenous variable vector $(m, y, p2, e2, r2)'$ for the potential long-run money demand space and test whether the expression embedded in (10) below using the multivariate co-integration methodology of the same order integrated variables holds:

$$\beta' z : (m, y, p2, e2, r2) \sim I(0) \tag{10}$$

For the lag length of unrestricted VAR, we consider various information criterions to select appropriate model between different lag specifications, i.e., sequential modified LR statistics employing small sample modification, the minimized Akaike information criterion (AIC), the final prediction error criterion (FPE), the Schwarz information criterion (SC) and the Hannan-Quinn information criterion (HQ). Considering the maximum lag 5 for the unrestricted VAR model of quarterly frequency data, LR, AIC and FPE criterions suggest to use 3 lags, while SC and HQ information criterion suggests 1 lag order. Thus we choose the lag length 3 to construct unrestricted VAR model. We add a set of centered seasonal dummies which sum to zero over a year as exogenous variable so that 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). For instance, the second quarter takes the value of 0.75 while the sum of the remaining three quarters' dummies is -0.75.

As a next step, we estimate the long run co-integrating relationships between the variables by using two likelihood test statistics. Briefly, the null hypothesis that there are at most r cointegrating vectors and k-r unit roots amounts to:

$$H_0: \lambda_i = 0, \qquad i = r+1, \dots, n \tag{11}$$

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 maximized likelihood function for the restricted model is compared to the log of the maximized likelihood function of the unrestricted model and a standard LR test computed. Using the trace statistic we can test the null hypothesis:

$$\lambda_{trace} = -2 \log (Q) = -T \sum_{i=r+1}^{n} \log (1-\lambda_i) \quad \text{and} \quad r = 0, 1, 2, \dots, n\text{-}2, n\text{-}1 \tag{12}$$

where Q = (restricted maximized likelihood / unrestricted maximised likelihood), T is the sample size. Another test of the significance of the largest λ_i is the maximal-eigenvalue statistic:

$$\lambda_{max} = -T \log (1-\lambda_{r+1}) \quad \text{and} \quad r = 0, 1, 2, \dots, n\text{-}2, n\text{-}1 \tag{13}$$

which tests that there are r co-integration vectors against the alternative that $r+1$ exist as expressed above. Table 4 below reports the results of the Johansen co-integration test using max-eigen and trace tests based on critical values taken from Osterwald-Lenum (1992) and on newer p-values for the rank test statistics from MacKinnon et al. (1999). Following Johansen (1992), for the co-integration test we restrict intercept and trend factor into our long run variable space following the so-called Pantula principle.

Table 4

Co-integration Test

Null hypothesis	r=0	r≤1	r≤2	r≤3	r≤4
Eigenvalue	0.49	0.29	0.17	0.14	0.11
λ trace	108.38	58.40	32.93	19.26	8.40
5% critical value	88.80	63.88	42.92	25.87	12.52
Prob.	0.00	0.13	0.34	0.27	0.22
λ max	45.97	25.47	13.67	10.86	8.40
5% critical value	38.33	32.12	25.82	19.38	12.52
Prob.	0.00	0.26	0.75	0.53	0.22

Unrestricted Co-integratig Coefficients

m	*y*	*r2*	*e2*	*p2*	*trend*
-6.702973	2.010034	-18.52946	-9.561195	7.081982	0.034179
-0.148580	3.069665	40.61782	-13.33160	-14.81748	-0.059962
-3.729614	-15.92201	-0.250599	10.79655	-22.07708	0.155549
2.163197	-11.19938	-3.721771	-13.35036	36.64724	0.188690
10.92687	-21.27735	12.59246	-5.197506	14.38208	0.150727

Unrestricted Adjustment Coefficients (alpha)

D(*m*)	0.034925	0.004218	-0.001360	-0.000116	-0.005964
D(*y*)	0.010174	-0.002081	-0.006164	-0.000180	0.008177
D(*r2*)	-0.005663	-0.012033	-0.006858	-0.002460	-0.005250
D(*e2*)	-0.016273	0.015933	-0.012363	-0.013421	-0.003651
D(*p2*)	-0.005627	0.004047	0.004372	-0.011299	0.001505

1 Co-integrating Equation (t-stat. in paranthesis): Log likelihood 699.3562

m	*y*	*r2*	*e2*	*p2*	*trend*
1.000000	-0.299872	2.764364	1.426411	-1.056543	-0.005099
	(-2.07409)	(2.21786)	(2.79696)	(-1.14854)	(-0.83007)

Adjustment coefficients (‘D’ indicates the first difference operator)

D(*m*)	D(*y*)	D(*r2*)	D(*e2*)	D(*p2*)
-0.234101	-0.068198	0.037957	0.109075	0.037716
(-6.28356)	(-2.39938)	(1.24870)	(2.02238)	(0.89463)

Multivariate Statistics for Testing Stationarity

	m	*y*	*r2*	*e2*	*p2*
$\chi^2(4)$	35.32861	34.76605	20.31421	16.6645	25.42053

Unit Income Homogeneity Restriction

b(1,2)=-1, $\chi^2(1)$ = 1.430427 Probability 0.231695

Results

From Table 4, both trace and max-eigen tests indicate 1 potential co-integrating vector lying in the long-run variable space:

Rewriting the normalized money demand equation under the assumption of r = 1 and applying to the unit income homogeneity restriction yield below (t-stats. in parantheses):

$$\beta' z = m - y + 2.764364 r2 + 1.426411 e2 - 1.056543 p2 + 0.005099 trend \sim I(0) \quad (14)$$
$$(2.21786) \quad (2.79696) \quad (-1.14854) \quad (0.83007)$$

The homogeneity restriction applied to the coefficient of real income is well-accepted by $\chi^2(1)$ = 1.43 under the null hypthesis. Moreover, co-integrating vector has good diagnostics and fit well with the data generating process in the VEC model using LM(1) = 33.19961 (prob. 0.1262), LM(4) = 28.06211 (prob. 0.3050), Skew(5) = 7.906535 (prob. 0.1615), Kur(5) = 18.38728 (prob. 0.0025), JB(10) = 26.29381 (prob. 0.0034), and Het (525) = 490.7403 (prob. 0.8555), where LM(1) and LM(4) are the 1st and 4th order VEC system residual serial correlation lagrange multiplier statistics under the null of no serial correlation, Skew the skewness, Kur the kurtosis, and JB the Jarque-Bera VEC residual normality statistics assuming Cholesky orthogonalization of Lütkepohl (1991) under the null hypothesis that system residuals are multivariate normal, which indicates no significant outliers in the model. Under the null of no heteroskedasticity or no misspecification, the VEC residual heteroskedasticity test accepts the null hypothesis. For the VEC system residual serial correlation test, probs. come from $\chi^2(16)$, and the values in parantheses for the system normality and heteroskedasticity tests are the degrees of freedom (d.o.f.) values considered. 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.

In Eq. 14 above, we find that the null hypothesis of homogeneity cannot be rejected for the real income elasticity of money demand. Between the alternative cost variables, the depreciation rate of nominal exchange rate and time-deposit rate have found in line with *a priori* expectations in a significant way. However, the coefficient of inflation rate has a wrong sign and is statistically insignificant as well. Considering *t*-statistics, the most significant alternative cost for economic agents to hold narrowly defined monetary balances is the depreciation rate of nominal exchange rate inside the period examined. This brings out the importance of currency substitution phenomenon settled in the economy when economic agents make their decisions about their monetary transactions.

In addition, an important policy conclusion which can be extracted from Table 4 is that domestic inflation is found weakly exogenous in the money demand variable space since the unrestricted adjustment coefficient for domestic inflation is highly close to zero. This requires that no feedback effect of disturbances from the steady-state money demand functional form can be constructed as a dynamic VEC model upon domestic inflation, and such a case reveals explicitly that the main factors leading to the domestic inflation are determined out of the money demand variable space considered in this paper. Whereas, in line with a quantity theoretical perspective, excess money derived from a money demand equation should have a positive significant effect on the inflation (Civcir, 2000). We give the graph of the co-integrating relationship:

Figure 1

Co-integrating Relation

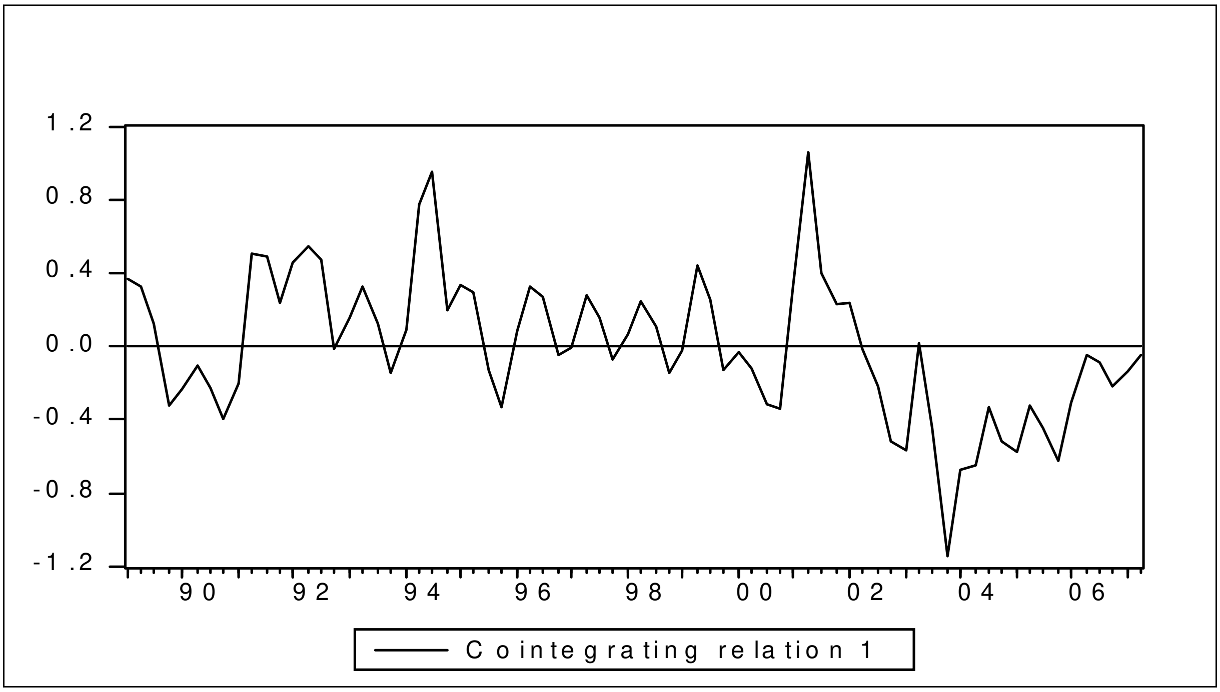


Table 5

Parsimonious VEC Model on Money Demand

Redundant Variables: Dm_{-1} Dm_{-3} Dy_{-1} Dy_{-3} $Dr2_{-1}$ $Dr2_{-2}$ $De2_{-1}$ $De2_{-3}$ $Dp2_{-2}$ $Dp2_{-3}$

F-statistic	0.750893	Prob.	0.673390
Log Likelihood Ratio	10.19797	Prob.	0.423300

Test Equation:

Dependent Variable: Dm

Method: Least Squares

Sample(adjusted): 1989Q2 2006Q4

White HCSE & Covariance

Variable[1,2]	Coefficient	Std. Error	t-Statistic	Prob.
C	0.022714	0.008226	2.761354	0.0079
EC_{-1}	-0.085659	0.030973	-2.765605	0.0078
D_Q2	-0.412219	0.117301	-3.514208	0.0009
D_Q3	-0.264375	0.106879	-2.473583	0.0167
D_Q4	-0.248697	0.042127	-5.903502	0.0000
Dm_{-2}	-0.302120	0.130452	-2.315940	0.0245
Dy_{-2}	-0.592103	0.189142	-3.130473	0.0029
$Dr2_{-3}$	-0.584205	0.158044	-3.696465	0.0005
$De2_{-2}$	-0.087062	0.029715	-2.929916	0.0050
$Dp2_{-1}$	0.373352	0.144009	2.592553	0.0123
Adj. R^2	0.617524	Akaike info criterion		-2.638895
S.E. of regression	0.060101	Schwarz criterion		-2.295809
Durbin-Watson Stat.	2.193092	F-statistic		11.94302

[1] 'D' indicates the first difference operator

[2] D_Q2, D_Q3 and D_Q4 are the centered seasonal dummies

Having established the long-run money demand co-integrating equilibrium model, we now estimate the dynamic VEC model using both a reduced form model with the econometrically meaningful variables shown and the estimated error correction term produced in the co-integrating relationship. Since all the variables in the model are now of stationary form, statistical inferences using standard t and F tests are valid. We have calculated the t-statistics of each variable by dividing the relevant coefficient by its standard error. We also apply an F-test for the reduction of insignificant variables in our model. EC is the estimated error correction coefficient upon money demand equation and the latter 'D' indicates the first difference operator. As Sriram (1999) emphasizes, 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 a 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. The parsimonious model has good diagnostics except non-normality problem due to excess kurtosis, using $LM(4)$=1.75 (0.09), Nor=0.02 (0.99), where LM is the Breusch-Godfrey Serial Correlation LM Test, and Nor is the Jarque-Bera Normality Test. Probs. are in parantheses. In Table 5, we find that nearly 8.5% of the adjustment in money demand disequilibrium conditions to long run static equilibrium is realized within one period.

Stability Tests

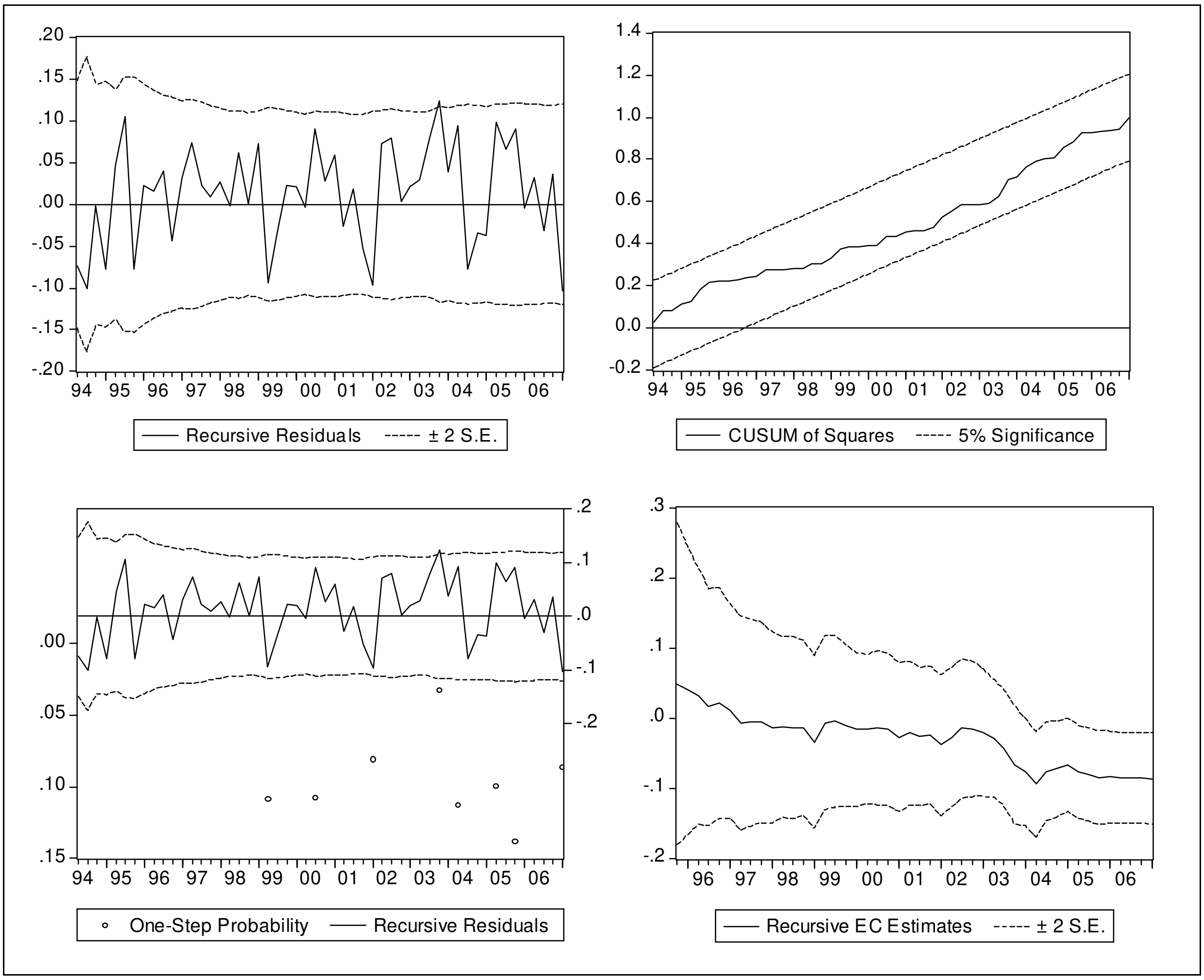
Establishing co-integration in the money demand variable space with appropriate signs as a long-run steady-state economic relationship may be interpreted as a sign of stable money demand functional relationship. However, Bahmani-Oskooee and Bohl (2000) and Bahmani-Oskooee and Karacal (2006) emphasize that evidence of co-integration does not imply constancy of estimated coefficients in co-integrating space. Following Laidler (1993), hence, possible break points inside the period as to our model specification of long-run money demand functional form should be sought.

In Figure 2 above, we first present the plot of recursive residuals about a zero line for the parsimonious error correction model. Considering ±2 standard error bands, residuals outside the standard error bands suggest instability in the parameters of the equation. A complementary test to the recursive residuals is the one-step forecast test that produces a plot of the recursive residuals and standard errors using sample points whose probability value is at or below 15%. The upper portion of the plot repeats the recursive residuals and standard errors displayed by the recursive residuals and the lower portion of the plot shows the probability values for those sample points where the hypothesis of parameter constancy would be rejected at the 5, 10, or 15% levels. The points with p-values less the 0.05 correspond to those points where the recursive residuals go outside the two standard error bounds.

Considering these tests, model stability has been in general satisfied. The possible parameter instability occurs for the post-2003 period.

Figure 2

Recursive Estimates



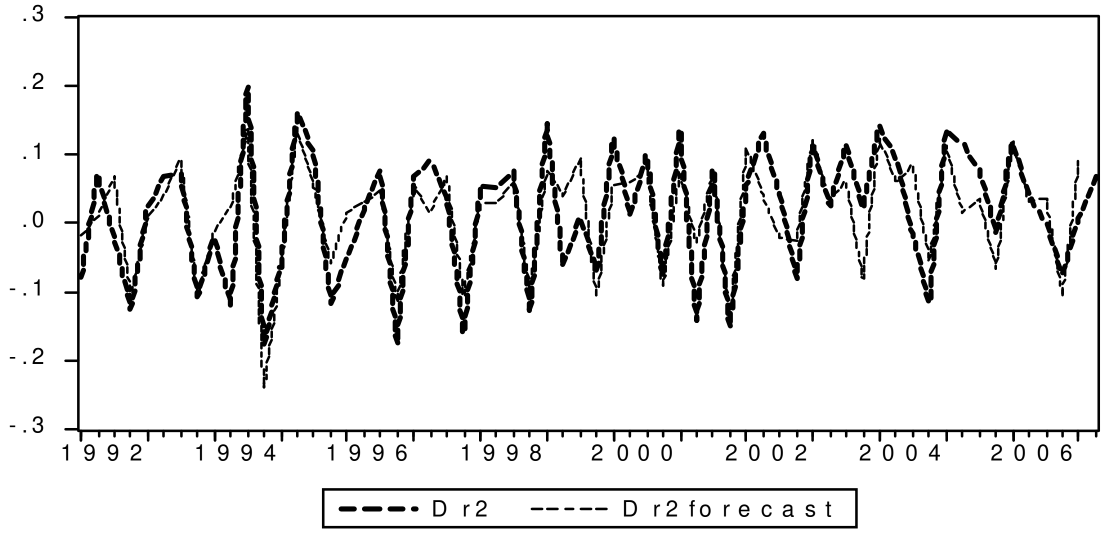
As with the CUSUM of Squares test, movement outside the critical lines is suggestive of parameter or variance instability. The cumulative sum of squares is within the 5% significance lines suggesting that the residual variance is stable, but they tend to approach to the margin of 5% significance level for the post-2003 period. Finally, recursive error correction (EC) estimates plot the evolution of estimates of the error correction coefficient which comes from the long-run co-integrating model as more and more of the sample data are used in the estimation. If the coefficient displays significant variation as more data is added to the estimating equation, this would be an indicator of instability. Recursive EC estimates yield results in line with recursive residual and one-step forecast tests for the narrowly defined monetary balances in the sense that major instabilities occur for the post-2003 crisis period.

Fitness to Turning Points

We now calculate whether the multi-step out-of-sample forecasts of the model can capture the turning points from the actual data of real money balances. For this purpose, we re-estimate the parsimonious money demand model by computing fully dynamic predictions from 1992 Q1 to 2007Q2 so that previously forecasted values of the lagged real money balances are used in forming the forecasts of the current values of real money balances. Such a forecasting methodology will differ from static forecasts using the actual values in estimation process. In Figure 3, we present a comparison of actual series and dynamic forecasts of real money balances:

Figure 3

Comparison of Actual and Forecated Real Money Balances



In Figure 3, we estimate that VEC modeling is highly successful in tracking down the path of actual data and that the model captures the turning points of actual real money balances well.

CONCLUDING REMARKS

Modeling demand for monetary balances is a useful guide to determine the long-run course of monetary policy and can give policy makers and researchers the knowledge of appropriate tools for stabilization policies. In our paper, we examine main determinants of currency in circulation by constructing a money demand model upon the Turkish economy. Our estimation results reveal that the most significant alternative cost to demand for money is the depreciation rate of nominal exchange rate and such a finding brings up the importance of the currency substitution phenomenon settled in the economy when economic agents make their decisions as to their monetary transactions. In addition, we find that domestic inflation is weakly exogenous in the money demand variable space so that no feedback effect of disturbances from the steady-state money demand functional form can be warranted to construct a dynamic vector error correction model upon domestic inflation. In this line, we concluded that main factors leading to the domestic inflation have been determined out of the money demand variable space, which contradicts especially what proponents of the Monetarist school of economic thought put forward as to the determination of the inflationary process. These all, of course, require additional researches and future papers considering money demand equations constructed upon various other monetary aggregates, both narrowly and broadly defined, in order to examine the robustness of the estimation results obtained in this paper.

Acknowledgement

This paper has been presented to Dündar M. Demiröz of the İstanbul University Department of Economics as a working paper by Levent KORAP during the author's PH. D. courses. The author would like to thank Mr. Demiröz for his criticisms, if can be any, upon the arguments advocated in this paper. The author is also grateful to Özlem Göktaş of the İstanbul University Department of Econometrics for the courses given to the author.

REFERENCES

Abu-Qarn, A.S. and Abu-Bader, S. (2007). "Structural Breaks in Military Expenditures: Evidence for Egypt, Israel, Jordan and Syria," *Monaster Center for Economic ResearchDiscussion Paper*, June, No. 07-04.

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

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

Bahmani-Oskooee, M. and Bohl, M. (2000). "German Monetary Unification and the Stability of Long-Run German Money Demand Function," *Economics Letters*, 66: 203-208.

Bahmani-Oskooee, M. and Karacal, M. (2006). "The Demand for Money in Turkey and Currency Substitution," *Applied Economics Letters*, 13: 635-642.

Banerjee, A., Lumsdaine, R.L., and Stock, J.H. (1992). "Recursive and Sequential Tests of the Unit Root and Trend Break Hypothesis," *Journal of Business and Economic Statistics*, 10: 271-287.

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

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

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

Civcir, I. (2000). "Broad Money Demand, Financial Liberalization and Currency Substitution in Turkey," *Paper Presented to the 7th Annual Conferenceof ERF*, 26-29 September, Amman.

Civcir, I. (2003). "Broad Money Demand and Currency Substitution in Turkey," *The Journal of Developing Areas*, 36: 127-144.

Clemente, J., Montanes, A., and Reyes, M. (1998). "Testing for a Unit Root in Variables with a Double Change in the Mean," *Economics Letters*, 59(2): 175-182.

DeJong, D.N., Nankervis, J.C., Savin, N.E., and Whiteman, C.H. (1989). "Integration versus Trend-Stationarity in Macroeconomic Time-Series," *University of Iowa Working Paper Series*, December, No. 89/31.

Dickey, D.A. and Fuller, W.A. (1979). "Distribution of the Estimators for Autoregressive Time Series with a Unit Root," *Journal of the American Statistical Association*, 74: 427-431.

Dickey, D.A. and Fuller, W.A. (1981). "Likelihood Ratio Statistics for Autoregressive Time Series with Unit Roots," *Econometrica*, 49: 1057-1072.

Dotsey, M. and Hornstein, A. (2003). "Should a Monetary Policymaker Look at Money?" *Journal of Monetary Economics*, 50: 547-579.

Dreger, C., Reimers, H.E., and Roffia, B. (2006). "Long-Run Money Demand in the New EU Member States with Exchange Rate Effects," *European Central Bank Working Paper Series*, May: No. 628.

Easterly, W.R., Mauro, P., and Schmidt-Hebbel, K. (1995). "Money Demand and Seigniorage-Maximizing Inflation," *Journal of Money, Credit and Banking*, 27(2): 583-603.

Elliott, G., Rothenberg, T.J., and Stock, J.H. (1996). "Efficient Tests for an Autoregressive Unit Root," *Econometrica*, 64: 813-836.

Estrella, A. and Mishkin, F.S. (1997). "Is There a Role for Monetary Aggregates in the Conduct of Monetary Policy?" *Journal of Monetary Economics*, 40: 279-304.

Friedman, M. (1956). "The Quantity Theory of Money - a Restatement," in M. Friedman (ed.), *Studies in the Quantity Theory of Money*: 3-21. The University of Chicago Press.

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

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

Goldfeld, S.M. (1976). “The Case of Missing Money,” *Brookings Papers on Economic Activity*, 3: 683-730.

Goldfeld, S.M. and Sichel, D.E. (1990). “The Demand for Money,” in B. M. Friedman and F.H. Hahn (eds.), *Handbook of Monetary Economics*, 1: 300-356.

Gonzalo, J. (1994). “Five Alternative Methods of Estimating Long-Run Equilibrium Relationships,” *Journal of Econometrics*, 60: 203-233.

Granger, C.W.J. and Newbold, P. (1974). “Spurious Regressions in Economics,” *Journal of Econometrics*, 2(2): 111-120.

Hafer, R.W. and Kutan, A.M. (1994). “Economic Reforms and Long-Run Money Demand in China: Implications for Monetary Polic,” *Southern Economic Journal*, 60(4): 936-945.

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

Johansen, S. (1992). “Determination of Cointegration Rank in the Presence of a Linear Trend,” *Oxford Bulletin of Economics and Statistics*, 54(3): 383-397.

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

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

Kontolemis, Z.G. (2002). “Money Demand in the Euro Area: Where Do We Stand (Today)?” *IMF Working Paper*, No. 02/185.

Kwiatkowski, D., Phillips, P.C.B., Schmidt, P., and Shin, Y. (1992). “Testing the Null Hypothesis of Stationary against the Alternative of a Unit Root,” *Journal of Econometrics*, 54: 159-178.

Laidler, E.W.D. (1993). *The Demand for Money: Theories, Evidence and Problems*, 4th ed. New York: Harper Collins College Publishers.

Lütkepohl, H. (1991). *Introduction to Multiple Time Series Analysis*. New York: Springer-Verlag.

MacKinnon, J.G. (1991). "Critical Values for Cointegration Tests," in R.F. Engle and C.W.J. Granger (eds.), *Long-run Economic Relationships: Readings in Cointegration*: Ch. 13. Oxford: Oxford University Press.

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

MacKinnon, J.G., Haug, A.A., and Michelis, L. (1999). "Numerical Distribution Functions of Likelihood Ratio Tests for Cointegration," *Journal of Applied Econometrics*, 14: 563-577.

MacKinnon, J.G. and Milbourne, R.D. (1988). "Are Money Demand Equations Really Price Equations on Their Heads", *Journal of Applied Econometrics*, 3: 295-305.

Mahadeva, L. and Robinson, P. (2004). "Unit Root Testing to Help Model Building," in A. Blake and G. Hammond (eds.), *Handbooks in Central Banking*, Centre for Central Banking Studies, Bank of England, July, No. 22

Metin, K. (1994). "Modelling the Demand for Narrow Money in Turkey," *METU Studies in Development*, 21: 231-256.

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

Nachega, J.C. (2001). "A Cointegration Analysis of Broad Money Demand in Cameroon," *IMF Working Paper*, WP/01/26.

Osterwald-Lenum, M. (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.

Özmen, E. (2003). “Testing the Quantity Theory of Money in Greece,” *Applied Economics Letters*, 10: 971-974.

Perron, P. (1989). “The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis,” *Econometrica*, 57: 1361-1401.

Perron, P. (1990). “Testing for a Unit Root in a Time Series with Changing Mean,” *Journal of Business and Economic Statistics*, 8: 153-162.

Ramachandran, M. (2004). “Do Broad Money, Output and Prices Stand for a Stable Relationship in India?” *Journal of Policy Modeling*, 26: 983,1001.

Sriram, S.S. (1999). “Demand for M2 in an Emerging-Market Economy: An Error-Correction Model for Malaysia,” *IMF Working Paper*, WP/99/173.

Tobin, J. (1956). “The Interest-Elasticiy of Transactions Demand for Cash,” *The Review of Economics and Statistics*, 38(3): 241-247.

Tobin, J. (1958). “Liquidity Preference as Behavior towards Risk,” *The Review of Economic Studies*, 25(2): 65-86.

Yavuz, N.C. (2004). “Durağanlığın Belirlenmesinde KPSS ve ADF Testleri: İMKB Ulusal-100 Endeksi ile bir Uygulama,” *İ.Ü. İktisat Fakültesi Mecmuası*, 54(1): 239-247.

Zivot, E. and Andrews, D.W.K. (1992). “Further Evidence of Great Crash, the Oil Price Shock and the Unit Root Hypothesis,” *Journal of Business and Economic Statistics*, 10: 251-270