Long-run relations between money, prices and output: the case of Turkey

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LONG-RUN RELATIONS BETWEEN MONEY, PRICES AND OUTPUT: THE CASE OF TURKEY

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ABSTRACT

In this paper, the long-run relationships between monetary aggregates, prices and real output level have been examined in a quantity theory of money perspective for the Turkish economy. Using some contemporaneous econometric techniques, our findings exhibit that stationary characteristics of the velocities of narrowly and broadly defined monetary aggregates cannot be rejected. However, monetary aggregates seem to have an endogeneity for the long-run evolution of prices and real income. Furthermore, some parameter instabilities and structural breaks have been attributed to the estimated model especially for the 1994 and 2001 economic crisis periods in the Turkish economy. We have concluded that given the endogenous characteristics of the monetary variables, monetary authority follows an accommodative monetary policy inside the period.

Keywords: *Quantity Theory of Money, Co-integration, Turkish Economy*

PARA, FİYATLAR VE ÇIKTI ARASINDAKİ UZUN-DÖNEMLİ İLİŞKİLER: TÜRKİYE ÖRNEĞİ

ÖZET

Bu çalışmada, parasal büyüklükler, fiyatlar ve reel çıktı seviyesi arasındaki uzun dönemli ilişkiler paranın miktar kuramı çerçevesinde Türkiye ekonomisi için incelenmektedir. Çağdaş bazı ekonometrik yöntemler kullanılmak suretiyle elde ettiğimiz bulgular dar ve geniş kapsamlı tanımlanan parasal büyüklüklere ait dolanım hızlarının durağan yapısının reddedilemeyeceğini göstermektedir. Bununla birlikte, parasal büyüklükler fiyatların ve reel gelirin uzun-dönemli gelişimi açısından içsel bir yapıda görülmektedir. Ayrıca, bazı parametre istikrarsızlıkları ve yapısal kırılmalar özellikle Türkiye ekonomisindeki 1994 ve 2001 ekonomik kriz dönemleri için tahmin edilen modelle ilişkilendirilmektedir. Sonuç olarak parasal değişkenlerin içsel yapılarının veri olduğu bir ortamda parasal yetkililerin uyumlaştırıcı bir para politikası izlediği sonucuna ulaşılmaktadır.

Anahtar Kelimeler: *Paranın Miktar Kuramı, Eş-Bütünleşim, Türkiye Ekonomisi*

1. INTRODUCTION

The functional relationships between persistent changes in price level, quantity of money and output level have been one of the main controversial theoretical issues of interest for economists, going back to the earlier times of capitalist development as discussed by David Hume (1970). The basic policy implications extracted from the hypotheses on which the quantity theory of money (henceforth, the QTM) is constructed have been of a special importance for researchers testing the role of money in assessing business cycles characteristics of an economy. Thus, revealing long-run stationary as well as short-run dynamic links leading to the quantity theoretical economic approaches would help researchers to determine how successful the *ex-ante* designed policies would be and which policy tools should be used to attain the desired policy conclusions for stabilization purposes. Resurrecting the interest upon the QTM, Friedman (1956) relates the QTM to the existence of stable functional relations that affect the quantity of money demanded and such a consideration in turn leads to the additional implication of the QTM that causes of variations in the velocity of money can be foreseen and explained by economic agents considering a stationary relationship as for the various phases of business cycles.

However, the role of money in providing adequate information for economic agents and policy makers have been criticized in various respects. Dotsey and Hornstein (2003) emphasize in their calibrating model upon the US economy that even though money provides sufficient information for aggregate output, it is of limited use for a policy maker 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 policy makers' responsiveness to money. Likewise, Estrella and Mishkin (1997) focus on the role of monetary aggregates as information variables and indicate that for the post-1979 period in the US economy, the monetary aggregates represented by either monetary base or M2 monetary aggregate fall considerably short of this requirement and results with German M3 broad money supply measures are hardly more favorable, which lead them to infer that since the monetary aggregates do not seem to provide adequate and consistent information, they cannot be used in a straightforward way to signal the stance of monetary policy. Therefore, as Meltzer (1998) stated, most researchers and policy makers, in recent times, tend to rely on the analyses based on the Phillips' curve or *atheoretical* relations in the construction of economic policies rather than on the money growth rates to predict the basic characteristics of the inflationary framework.

Such issues can also be related to the criticisms of Lucas (1981) that examines both the empirical model evaluation process of researchers and the changes in the motives that determine the decision making of economic agents and policy makers. Considering the well-known Lucas' critique, since the optimal decision rules on which the structure of econometric models are based have been varied with changes in the structure of series that represent the behavior of economic agents, the structure of econometric models used for estimation purposes will have been also altered by the systematic changes in the policy choices. Following Lucas, such a proposal is of great interest for policy makers and assuming that the critique holds gives rise to that comparisons of the effects of alternative policy rules using current macroeconomic models will be invalid regardless of the performance of these models over the sample period or in short-run forecasting. Assuming also that expectations are constructed rationally by economic agents leads us easily to infer that policy evolution processes considered to have an exogeneous characteristic in Keynesian and Monetarist models have been imposed with an endogenous expectation formation process conditional upon forecasts for the results of policy implementations (Ardıç, 1996). Thus, rigid assumptions of the one-way causal relationship between the variables of the QTM long-run equilibrium space without elaborately testing them, e.g. assuming *a priori* long-run exogeneity of money supply changes and endogeneity of the changes in price level which respond to the former when relating them to each other, would lead to the inconsistent long-run economic forecasts following specified model construction and such a case would invalidate the policy conclusions derived from structural econometric models. These all bring out the importance of the stability of functional relationships once again for the construction of the QTM and the critical assumptions used for this purpose must be elaborately examined to search for whether they can be supported in a way providing internal consistency of the theory. Following Lucas (1980), this would help us to provide solutions of explicit theoretical models of idealized economies to explore why one might expect the theory to hold in reality and to explain the conditions under which the theory might be expected to break down. On these issues of interest, Stanley (2000) gives a review of theoretical and empirical papers.

Considering these criticisms in the model construction process, in this paper, our aim is to examine the validity of the QTM relationship for the Turkish economy in an empirical way. For this purpose, the next section is devoted to the theoretical background of the QTM and a contemporaneous literature review is presented in Section 3. Section 4 is devoted to the data processing and econometric model construction issues and tries to conduct an empirical model upon the Turkish economy. Finally, the last section summarizes results and concludes.

2. THEORETICAL FUNDAMENTALS

2.1. The Main Model

The QTM based on the classic book by Fisher (1911) can be described by the well known exchange identity:

$$M V_T = P T \qquad (1)$$

where M is the money supply, V_T the transactions velocity of money, P the general price level and T the economic transactions volume in the economy in a given time period. Following Mishkin (1997), however, because the nominal value of transactions T is difficult to measure, it can be replaced by aggregate output level Y under the simplifying assumption that T would be proportional to Y as follows:

$$T = \upsilon Y \tag{2}$$

where υ is a constant of proportionality. Substituting υY for T would yield:

$$M\,V = \upsilon\,P\,Y \tag{3}$$

where now V, the income velocity of money as a function of institutional structure of the financial system *ex-ante* assumed time invariant, equals VT / υ. Following Pigou (1917) and considering the importance of money demand relationship in explaining the implications related to the QTM, Eq. 3 can also be re-written as follows:

$$M / P = kY \tag{4}$$

where k equals the inverse of income velocity of money and indicates the ratio of money balances demanded by economic agents in proportion to real income. Eq. 4 assumes that economic agents have been subject to no money illusion which requires that if prices increase then people want to hold more money so that money would buy the same amount of goods and services (Dwyer and Hafer, 1999). It reveals that the larger the aggregate income level, the larger would be the aggregate spending in turn leading economic agents to increasing their money holdings with a k proportion of income, which is also called the Cambridge k. Thus re-specifying the QTM in this way would allow researchers to examine the factors that affect the quantity of money demanded, which must be consisted of a set of opportunity costs to hold money other than the scale-real income variable representing maximum limit of money balances economic agents can hold. An important contribution of the Cambridge k to the quantity theory is to indicate that if the demand for money by economic agents has been of an unstable form resulting from the variation in the opportunity costs of money, e.g., due to the changes in expected inflation and interest structure dominated in the economy leading also to the changes in the demand for monetary balances, these latter factors can also give the QTM relationship an unstable functional form that destabilizes the implications based on the stable velocity of money.

A critical assumption extracted from this relationship is that quantity of money demanded and supplied in the aggregate level equal at least over the long horizons so that if the quantity of money supplied increases, either the desired ratio of money holdings to real income or the nominal income must increase (Dwyer and Hafer, 1988).[1] Otherwise, in terms of the new quantity theory of Friedman and following Fitzgerald (1999), the price level would work to equate the quantity of money demanded with the quantity of money supplied.

2.2. Some Extensions

Having specified the construction of the QTM relationship in a two related way, some other policy implications can be derived more explicitly. Assume the QTM in terms of the growth rates:

$$m + v = p + y \qquad (5)$$

where the lower case letters denote the growth rates. The QTM relationship requires that there exist proportional relationships between the growth rates of money supply and price level and that money must be (super)neutral which is resulted from stationary velocity of money and unaffected real output level in the long-run following the permanent changes in the growth rate of money supply.

Note here that testing a variable vector $X = (\Delta Y, \Delta M)'$, where logarithm of the money stock, M, and logarithm of the real output, Y, are assumed to follow an I(1) process, means to examine the neutrality of money, whereas if the process describing M is I(2) rather than I(1) then we test the concept of (super)neutrality by using the variable vector $X = (\Delta Y, \Delta 2M)'$. King and Watson (1997) emphasize that long-run neutrality cannot be tested in a system in

[1] The authors thank Merih PAYA who draws their attention on this issue.

which output is I(1) and money is I(2). This is because neutrality of money refers to the hypothesis that changes in the quantity of money affect the nominal variables in the macroeconomic system and concern the relationship between shocks to the level of money and the level of output. However, if an I(2) process dominates the money supply, shocks in this case would affect the rate of money growth and there would be no shocks to the level of money. Fisher and Seater (1993) and Bullard (1999) argue various cases for long-run neutrality and (super)neutrality of money that depend on the integration of the variables.

Following Ozmen (2003) and Grauwe and Polan (2005), for empirial purposes, the QTM requires that each of *m*, *p* and *y* or their linear combination with a coefficient vector (-1 1 1) must be stationary. That is, a long-run I(0) process must dominate this variable space leading to that velocity of money (*v*) has been subject to a stationary long-run process. Ozmen states that even if this requirement constitutes a necessary condition for the quantity theory, this is not a sufficient condition, since the QTM contains also the exogeneity of money in the velocity variable system which requires that money supply must be weakly exogenous for the long-run evolution of prices and real income. Otherwise, an endogenous money supply framework would be validated within the quantity theory variable system.

3. LITERATURE REVIEW

Geweke (1986) using a century of annual US data as well as postwar monthly data for the US economy, King and Watson (1987) testing various long-run neutrality propositions using postwar US data, Serletis and Krause (1996) and Serlestis and Koustas (1998) using a low frequency data from ten developed countries over one hundred years, and Koustas (1998) testing neutrality using post WWII data for the Canadian economy give in general strong support for the long-run neutrality proposition. Likewise, Bullard (1999) reports a large review of papers upon long-run monetary neutrality and (super)neutrality propositions and emphasizes that there exists a general evidence in favor of the neutrality proposition but no clear-cut inference can be drawn from the international evidence of (super)neutrality.

Karfakis (2002) tests the predictability of income velocity and the proportionality of nominal income (or, prices) and money using Greek data. He finds that proportionality is supported by the data and that velocity does not fluctuate widely and movements in the velocity would be predictable. However, Ozmen (2003) re-examining the Greek data used by Karfakis (2002) reveals that contrary to the findings of Karfakis the Greek data strongly reject the exogeneity of money in a velocity variable system. He concludes that money and nominal income (or, prices) appear to be jointly determined in a consistent way with an endogenous money hypothesis. In reply to the Ozmen (2003), Karfakis (2004) addresses the issues raised by Ozmen and demonstrates that money can be treated as a long-run driving variable for nominal income in Greece and expresses that stationarity of the income velocity of money and validity of proportionality support the QTM by using Greek data.

Ashra et al. (2004) examine the relationship between money, output and price level for the case of a developing country, i.e., India. They emphasize that the Monetarist strategy to monitor money supply to check inflation assumes, *inter alia*, exogeneity of money. However, their findings indicate that there exists a bidirectional causality between money and price level and that money is non-neutral so that it is not exogenous in the long-run. Grauwe and Polan (2005) using a large panel of low- and high-inflation countries find that the QTM prediction that an expansion of the money stock does not increase output in the long-run is confirmed. Finally, Herwartz and Reimers (2006) analyse the dynamic relationships between money, real output and prices for an unbalanced panel of 110 economies and find that particularly for high inflation countries homogeneity between prices and money cannot be rejected. They suggest that central banks, even in high inflation countries, can improve price stability by controlling monetary growth.

4. EMPIRICAL MODEL

4.1. Data

In this section, we consider data for the investigation period 1987Q1 – 2007Q2 using quarterly observations for the model construction purposes. All the data take the form of seasonally unadjusted values in their natural logarithms and are taken from the electronic data delivery system of the Central Bank of the Republic of Turkey (CBRT). Following Lucas (1980), for the appropriate money supply variable two variable specifications are considered to verify the consistency of results for different monetary aggregates, represented by either narrow money supply, i.e., M1 monetary aggregate ($m1$) as a sum of currency in circulation plus sight deposits in the banking system, or broad money supply, i.e., M2 monetary aggregate (*m*2) as a sum of M1 monetary aggregate plus time deposits in the banking system. The gross domestic product (GDP) deflator (*p*) is used to represent the relevant price measure for which the log-differenced form of the deflator would be the quarterly inflation. Real income variable (*y*) is scaled by the real GDP data, as well. Two impulse-dummy variables which take on values of unity from 1994Q1 till 1994Q4 and from 2001Q1 till 2001Q4 concerning the financial crises occured in 1994 and 2001 are considered exogenous variables.

4.2. Testing Unit Roots Allowing for Endogenous Breaks

Spurious regression problem analysed by Granger and Newbold (1974) indicates that using non-stationary time series steadily diverging from long-run mean would produce biased standard errors, which causes to unreliable correlations within the regression analysis leading to unbounded variance process. This means that the variables must be differenced (*d*) times to obtain a covariance-stationary process. Therefore, individual time series properties of the variables should be elaborately considered. However, conventional tests for identifying the unit roots in a time series, e.g., the augmented Dickey-Fuller test (Dickey and Fuller, 1979) and Phillips-Perron test (Phillips and Perron, 1988) 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 not considering 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.

Selecting the date of the structural break, however, may not be the most efficient methodology, because the actual dates of structural breaks may not be coincided with the dates chosen exogenously. To address this issue, several methodologies have been suggested to allow for the determination of the date of the structural break endogenously, including those advanced by Zivot and Andrews (1992), Banerjee, Lumsdaine and Stock (1992) and Perron (1990). For this purpose, we follow first the Zivot and Andrews (henceforth ZA) methodology allowing the data themselves to indicate breakpoints endogenously rather than imposing a breakpoint from outside the system. We then allow some extensions of this test by following Clemente et al. (1998) that employ unit root tests of double changes in the mean.

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 test the equation of the form:

$$y = \mu + y_{t\text{-}1} + e_t \tag{6}$$

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 occuring at some unknown time. The ZA test chooses the breakpoint as the minimum *t*-value on the autoregressive *yt* variable, which occurs at time 1 < TB < T leading to $\lambda = TB / T$, $\lambda \quad |0.15, 0.85|$, by following the augmented regressions:

Model A:

$$y_t = \mu + \beta t + \theta DU_t(\lambda) + \alpha y_{t\text{-}1} + \sum_{j=1}^{k} c_j \, \Delta y_{t\text{-}j} + \varepsilon_t \tag{7}$$

Model B:

$$y_t = \mu + \beta t + \gamma DU_t(\lambda) + \alpha y_{t\text{-}1} + \sum_{j=1}^{k} c_j \, \Delta y_{t\text{-}j} + \varepsilon_t \tag{8}$$

Model C:

$$y_t = \mu + \beta t + \theta DU_t(\lambda) + \gamma DU_t(\lambda) + \alpha y_{t\text{-}1} + \sum_{j=1}^{k} c_j \, \Delta y_{t\text{-}j} + \varepsilon_t \tag{9}$$

trend shift occuring at the break date respectively, i.e. $DUt(\lambda) = 1$ 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 identically and independently distributed (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 the 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.

Besides, 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 additional outliers (AO). Following Clemente et al. (1998), for the case where the two breaks belong to the AO, we can test the unit root null hypothesis through the following two-step procedure. First, we should remove the deterministic part of the variable:

$$y_t = \mu + d_1 DU_{1t} + d_2 DU_{2t} + \tilde{y} \tag{10}$$

and, subsequently, carry out the test by searching for the minimal t-ratio for the $\alpha = 1$ hypothesis in the following model:

$$\tilde{y} = \sum_{j=0}^{k} \omega_{1i}\mathrm{DTB}_{1t-i} + \sum_{j=0}^{k} \omega_{2i}\mathrm{DTB}_{2t-i} + \alpha\tilde{y}_{t-1} + \sum_{j=1}^{k} \theta_i \Delta\tilde{y}_{t-i} + e_t \tag{11}$$

and if we consider that the two breaks belong to the innovational outlier, we can also test the unit root hypothesis by first estimating the following model:

$$y_t = \mu + \alpha y_{t-1} + \delta_1 DTB_{1t} + \delta_2 DTB_{2t} + d_1 DU_{1t} + d_2 DU_{2t} + \sum_{j=1}^{k} c_i \Delta y_{t-i} + e_t \tag{12}$$

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 the mean occur. Eq. (12) is sequentially estimated and the unit root hyothesis is tested by obtaining the minimal value of the t-statistic for the hypothesis $\alpha = 1$ for all break time combinations.

Table 1: Zivot-Andrews Unit Root Test

Var	Intercept			Trend			Both		
	k	min *t*	TB	*k*	min *t*	TB	*k*	min *t*	TB
m1	1	-2.77	(2004Q2)	2	-4.29	(2001Q4)	2	-4.65	(2001Q3)
m2	1	-1.66	(1994Q2)	1	-2.99	(2000Q1)	1	-2.49	(2001Q1)
p	0	-2.75	(1998Q3)	0	-3.80	(1997Q4)	0	-4.15	(1998Q3)
y	0	-3.49	(1998Q4)	0	-3.31	(2003Q2)	0	-4.12	(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 2: Clemente-Montañés-Reyes Unit Root Test with Double Mean Shift

Variable	Additive Outliers		Innovative Outliers	
	min *t*	Optimal Breakpoints	min *t*	Optimal Breakpoints
m1	-2.09	1993Q4, 1999Q2	-2.22	1992Q3, 1996Q1
m2	-2.98	1994Q4, 1999Q4	-4.13	1988Q2, 1999Q4
p	-3.09	1999Q4, 2002Q4	-4.66	2000Q1, 2003Q1
y	-3.44	1999Q4, 2004Q1	-2.02	1999Q2, 2003Q4

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

For estimation purposes, we used Stata 9.0 for ZA (1992) test and Clemente et al. (1998) unit root test of double changes in the mean, for which the latter test procedures can be obtained from the web site of instructor as routines clemao2 and clemio2.[2] Using the ZA procedure, the time of structural breaks is detected based on the most significant *t*-ratio for α. When we consider the ZA unit root tests in Tab. 1 above allowing one-time endogenous structural break in the time series used, we cannot reject the unit root null hypothesis for all the variables. The

[2] The authors thank Ozlem GOKTAS-YILMAZ, Ferda YERDELEN, Burak GURIS and Veli YILANCI for their kind support in implementing estimation procedures using software Stata.

breakpoints for the money supply variables $m1$ and $m2$ coincide in general with either economic crisis periods such as 1994 and 2001 economic crises or periods of structural changes in the economy such as 2000 stabilization program. For the deflator-based price level, the 1997-1998 period repesents a structural break which may be related to the policy changes of the monetary authority in favor of monitoring monetary variables against domestic inflationary framework in the Turkish economy. Likewise, the real income variable has been subject to structural breaks for the economic stagnation or crisis periods of 1998Q4 and 2001Q1. These results are also supported by the Clemente et al. (1998) unit root tests in Tab. 2 for up to two shifts in the mean of the series for both the AO and IO cases. The 1999 economic stagnation period is a common breakpoint for both additive and innovative outliers in all the time series. Despite the structural breaks, therefore, we are unable to reject the null hypothesis of a unit root.

4.3. 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:

$$z_t = \Pi_1 z_{t-1} + \Pi_2 z_{t-2} + \ldots + \Pi_k z_{t-k} + \varepsilon_t \tag{13}$$

where ε_t follows an i.i.d. process and z is ($nx1$) and the Π_i an (nxn) matrix of parameters. Eq. 13 can be re-written leading to a vector error correction (VEC) model of the form:

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + \Gamma_2 \Delta z_{t-2} + \ldots + \Gamma_{k-1} \Delta z_{t-k+1} + \Pi z_{t-k} + \varepsilon_t \tag{14}$$

where:

$$\Gamma_i = -I + \Pi_1 + \ldots + \Pi_i \; (i = 1, 2, \ldots, k\text{-}1) \text{ and } \Pi = I - \Pi_1 - \Pi_2 - \ldots - \Pi_k \tag{15}$$

Eq. 14 can be arrived by subtracting z_{t-1} from both sides of Eq. 13 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. 14 (Hafer and Kutan, 1994). This specification of the system of variables carries on the knowledge of both the short- and the longrun 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 $\beta' z_{t-k}$ embedded in Eq. 14 represents up to (n-1) cointegrating relations in the multivariate model which ensure that zt converge to their long-run steady-state solutions. Note that all terms in Eq. 14 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, σ_ε^2) process.

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, minimized Akaike information criterion (AIC), final prediction error criterion (FPE), Schwarz information criterion (SC) and Hannan-Quinn information criterion (HQ). Considering the maximum lag length 5 for the unrestricted VAR model using quarterly frequency data, LR, AIC, FPE and HQ criterions suggest to use lag lenght 3 for the model using M1 monetary aggregate, while SC information criterion suggests lag length 1. For the model using M2 monetary aggregate, LR, AIC, FPE and HQ suggest to use lag length 4, while SC statistic suggests lag lenght 3. Thus we choose the lag length 3 for the first and the lag length 4 for the second 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 by using two likelihood test statistics known as maximum eigenvalue for the null hypothesis of r versus the alternative of r+1 co-integrating relations and trace for the null hypothesis of r cointegrating relations against the alternative of n co-integrating relations, for r = 0,1, ... ,n-1 where n is the number of endogenous variables.

4.4. Co-integration Results

Tab. 3 and Tab. 4 below report the results of Johansen co-integration test using max-eigen and trace tests based on critical values taken from Osterwald-Lenum (1992) and 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 a long-run deterministic trend into our long run variable space following the *so-called* Pantula principle, but no deterministic trend is assumed for the dynamic VEC model. This requires a test procedure which moves through from the most restrictive model and at each stage compares the trace or max-eigen test statistics to its critical value and only stop the first time the null hypothesis is not rejected.

From Tab. 3 and Tab. 4, both LR tests verify the existence of 1 potential cointegrating vector lying in the long-run variable space. Rewriting the normalized QTM equation upon the money supply variable $m1$ under the assumption of $r = 1$ and applying to the homogeneity and symmetry restrictions in line with the quantity theory yield below:

$$\beta'_{m1}z_t = m1 - p - 2.323627y + 0.024939trend + 13.97589 \sim \mathrm{I}(0) \tag{16}$$

$$\beta'_{m1}z_t = m1 - p - y + 0.012462trend + 1.050425 \sim \mathrm{I}(0) \tag{17}$$

The restrictions are well-accepted by the χ^2 tests. In Tab. 3, we accept the homogeneity restriction for only price level variable with $\chi^2(1) = 0.127419$ and for both price and output variables with $\chi^2(2) = 3.226091$ under the null hypothesis. Likewise, the normalized equation inclusive of $m2$ money supply variable can be given below:

$$\beta'_{m2}z_t = m2 - 1.579841p - 7.871581y + 0.146396trend + 61.18190 \sim \mathrm{I}(0) \tag{18}$$

However, the symmetry and homogeneity restrictions here cannot be accepted under the usual significance levels which yield prob. values under 5% for the null hypothesis. Both co-integrating vectors fit well to the data generating process in the VEC models when we consider the diagnostic estimation results. 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 Tab. 3 and Tab. 4, we find that estimation results are consistent with quantity theory as for the signs of the variables in a significant way and long-run exclusion of the each variable from the stationary variable space can also be rejected. We accept the symmetry and homogeneity restrictions for the model using M1 monetary aggregate, as well. For the model using M2 monetary aggregate,

Table 3: Co-integration Test (using M1 monetary aggregate)

Sample (adjusted): 1988Q1 2007Q2
Included observations: 78
Trend assumption: Linear deterministic trend (restricted)
Series: *m1 p y*
Exogenous series: dummy94 dummy2001 d_q2 d_q3 d_q4
Lags interval (in first differences): 1 to 3

Null hypothesis	r=0	r≤1	r≤2
Eigenvalue	0.35	0.19	0.08
λ trace	56.21	22.29	6.14
5% critical value	42.92	25.87	12.52
Prob.	0.00	0.13	0.44
λ max	33.92	16.15	6.14
5% critical value	25.82	19.39	12.52
Prob.	0.00	0.14	0.44

Unrestricted Co-integrating Coefficients

m1	*p*	*y*	*trend*
4.568617	-5.050151	-13.53810	0.208312
12.04453	-7.345213	10.68671	-0.746479
-2.892050	-0.396325	-13.62585	0.504324

Unrestricted Adjustment Coefficients (alpha)

D(*m1*)	-0.017866	-0.017284	-0.003235
D(*p*)	-0.027723	0.002961	0.009664
D(*y*)	0.013637	-0.007930	0.003300

1 Co-integrating Equation (t-stat. in parantheses): Log likelihood 416.3133

m1	*p*	*y*	*trend*	*C*
1.000000	-1.105401	-2.963282	0.045596	20.13076
	(-8.19661)	(-2.98353)	(2.06276)	

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

D(*m1*)	D(*p*)	D(*y*)
-0.081622	-0.126655	0.062302
(-2.91962)	(-4.20788)	(3.74650)

Multivariate Statistics for Testing Stationarity

	m1	*p*	*y*
$\chi^2(2)$	16.16876	13.77842	14.28772

Homogeneity and Symmetry Restrictions on Co-integrating Coefficients
b(1,1)=1, b(1,2)=-1, $\chi^2(1)$ = 0.127419 Prob. 0.721123
b(1,1)=1, b(1,2)=-1, b(1,3)=-1, $\chi^2(2)$ = 3.226091 Prob. 0.199280
VEC Residual Serial Correlation LM Test (Probs. from chi-square with 9 df.)
LM(4) = 4.985335 (Prob. $\chi^2(9)$ 0.8356)
VEC Residual Normality Test
Skewness $\chi^2(3)$ = 5.277036 (Prob. 0.1526)
Kurtosis $\chi^2(3)$ = 3.297656 (Prob. 0.3480)
Jarque-Bera $\chi^2(6)$ = 8.574692 (Prob. 0.1989)

Table 4: Co-integration Test (using M2 monetary aggregate)

Sample (adjusted): 1988Q2 2007Q2
Included observations: 77
Trend assumption: Linear deterministic trend (restricted)
Series: *m2 p y*
Exogenous series: dummy94 dummy2001 d_q2 d_q3 d_q4
Lags interval (in first differences): 1 to 4

Null hypothesis	r=0	r≤1	r≤2
Eigenvalue	0.34	0.18	0.11
λ trace	55.83	23.67	8.70
5% critical value	42.92	25.87	12.52
Prob.	0.00	0.09	0.20
λ max	32.15	14.98	8.70
5% critical value	25.82	19.39	12.52
Prob.	0.01	0.19	0.20

Unrestricted Co-integrating Coefficients

m2	*p*	*y*	*trend*
2.849190	-4.501268	-22.42763	0.417110
-9.635733	7.845554	-6.026598	0.405759
3.165284	-3.586059	-18.99420	0.129815

Unrestricted Adjustment Coefficients (alpha)

D(*m2*)	-0.016324	0.010182	-0.007767
D(*p*)	-0.023764	0.005735	0.010747
D(*y*)	0.014636	0.007309	0.001515

1 Co-integrating Equation (t-stat. in parantheses): Log likelihood 439.7728

m2	*p*	*y*	*trend*	*C*
1.000000	-1.579841	-7.871581	0.146396	66.90438
	(-6.54612)	(-4.13410)	(3.53337)	

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

D(*m2*)	D(*p*)	D(*y*)
-0.046510	-0.067707	0.041699
(-3.04624)	(-3.72752)	(4.10581)

Multivariate Statistics for Testing Stationarity

	m2	*p*	*y*
$\chi^2(2)$	14.54302	14.07429	11.49768

Homogeneity and Symmetry Restrictions on Co-integrating Coefficients
b(1,1)=1, b(1,2)=-1, $\chi 2(1)$ = 5.111917 Prob. 0.023762
b(1,1)=1, b(1,2)=-1, b(1,3)=-1, $\chi^2(2)$ = 9.099327 Prob. 0.010571
VEC Residual Serial Correlation LM Test (Probs. from chi-square with 9 df.)
LM(4) = 7.813568 (Prob. $\chi^2(9)$ 0.5530)
VEC Residual Normality Test
Skewness $\chi^2(3)$ = 4.098313 (Prob. 0.2510)
Kurtosis $\chi^2(3)$ = 5.990072 (Prob. 0.1121)
Jarque-Bera $\chi^2(6)$ = 10.08839 (Prob. 0.1210)

we support a case of near-proportionality of money and prices but now not in a one-to-one way. Thus, these results yield a strong support to the *ex-post* stationary characteristic of the velocity of money in a quantity theoretical stable functional relationship.

However, we are unable to find both money supply variables as weakly exogenous in the long-run variable space. In both Tab. 3 and Tab. 4, all adjustment coefficients indicating feedback effects of disturbances from the steady-state functional forms and carrying the long-run knowledge from co-integrating vectors into the VEC models are found highly different from zero in a statistically significant way. Such a finding requires that VEC models upon all these endogenous variables can be constructed through error-correction mechanism. Following Ozmen (2003), no variable alone can be interpreted as the uni-directional forcing variable for the long-run evolution of the other variables, and this imposes them an endogenous characteristic in the QTM long-run variable space. Ozmen attributes such a result to that this would contradict the QTM assumption that money is the sole forcing variable in the multivariate co-integrating system and he gives support to an endogenous money creation framework conditioned upon long-run courses of prices and real income. Thus, rejecting the weak exogeneity of both real income and money supplies considering a positive relationship does not support the neutrality hypothesis embedded in the quantity relationship. For the design of monetary policy, a possible explanation can be brought out such that monetary authority seems to follow an accommodative monetary policy inside the period given the endogenous characteristics of the monetary variables. These all would weaken the discretionary policy role of money in the conduct of future stabilization policies.

Having established the main theoretical model and tested assumptions on which the theory is constructed, we now try to test the (super)neutrality of money. Following Grauwe and Polan (2005), for the (super)neutrality condition to hold, a permanent increase in the growth rate of money must leave output unaffected in the long-run. If there is a positive effect of money growth on output, it only holds in the short run. To test this proposition, we estimate the following equation:

$$\Delta y = \alpha + \delta ec_{-1} + \alpha\Delta m + \beta\Delta p + \varepsilon \qquad (19)$$

where, Δy is the growth rate of real output, Δm the growth rate of money supply and Δp the growth rate of prices all expressed in log differences, and ε is again $N(0,\sigma^2)$ white-noise error term. The OLS results including stationary knowledge of long-run relationship in co-integration analysis as one period lagged error correction term (ec_{-1}) are given below (using White HCSE&Covariance):

Table 5: OLS Estimation Results for (Super)Neutrality of Money

Var.	Coefficient	Std. error	*t*-statistic	Wald tests (α = 1) (p-value = 0.0012)
C	-0.171584	0.046315	-3.704739	
ec_{-1}	0.260340	0.057921	4.494754	
$\Delta m1$	2.168818	0.345938	6.269375	
Δp	-0.585784	0.285842	-2.049326	
Adj. R^2	0.310033		D-W stat.	1.951524
S.E. of reg.	0.212736		F-stat.	12.38340

Table 6: OLS Estimation Results for (Super)Neutrality of Money

Var.	Coefficient	Std. error	*t*-statistic	Wald tests (α = 1) (p-value = 0.6619)
C	-0.075242	0.059131	-1.272462	
ec_{-1}	0.223826	0.072368	3.092882	
$\Delta m2$	1.191054	0.435083	2.737535	
Δp	-0.541837	0.348105	-1.556535	
Adj. R^2	0.095488		D-W stat.	1.922765
S.E. of reg.	0.243575		F-stat.	3.674417

Results in Tab. 5 and Tab. 6 reveal that we reject the (super)neutrality condition for both M1 and M2 money supply measures. Changes in the growth rate of money supply lead to a significant increase in the real output growth rate. We find through the lagged error correction term that excess money in nominal terms leads to an increase in real income growth rate. Besides, there exists a negative relationship between real income growth and changes in the price level, i.e. domestic inflation, though this relationship has not a statistical meaning in acceptable significance levels in Tab. 6.

4.5. Stability Tests

If the conditional economic models have been found dependent on specific policy actions and institutional structures of the economy though they have been estimated by using most recent or popular econometric techniques, substantial changes in policies or the institutional structure, in this case, may lead reserachers unwarranted estimation results and nullify the best econometric models even when the estimates seem to have desired statistical prerequisites (Stanley, 2000). In such circumstances subject to the well-known Lucas critique through Lucas (1981), the theories and policies *ex-ante* assumed for estimation purposes would have been undermined leading to the invalidated estimations and policy proposals.

Establishing co-integration in the variable space with appropriate signs as a long-run steady-state economic relationship may be interpreted as a sign of stable functional relationship. However, evidence of co-integration should not be taken as evidence in favor of constancy of estimated coefficients in co-integrating space, and the estimated functional form can be found in this case subject to structural breaks and parameter instabilities, as well. Hence, possible break points inside the period as for the model specification should be searched for elaborately, otherwise "... not only dynamic misspecifications but also an invalid conditioning and a change in the relevant variable space ... *due to a policy regime change* should be taken as complementary explanations for parameter instabilities" (Ozmen, 1996: 272)[3]. Above, we find that the weak exogeneity condition can be rejected for all the variables, because the adjustment coefficients of each variable in both the model using M1 and the model using M2 monetary aggregate have been found highly significant in a statistical sense. Therefore, in this paper, we will focus upon the model stability tests to see whether the estimated model exhibits structural breaks inside the period examined.[4]

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[3] Italics have been changed by ourselves.

[4] This also requires testing the superexogeneity of the variables of interest, which can be implemented by constructing marginal models. However, since we reject the null hypothesis of weak exogeneity for all the variables in this paper, we only deal with system stability tests. But, different modeling approaches especially on

In Fig. 1 and Fig. 2, we first present the plot of recursive residuals about a zero line for the error correction models derived from the co-integrating relationships estimated in Tab. 3 and Tab. 4 above. Considering ±2 standard error bands, residuals outside the standard error bands suggest instability in the parameters of the equation. We can easily notice that for the model using M1 money supply the first period of 1991, the 1994 crisis period and the subsequent periods and the period of 2000 stabilization program witness parameter instabilities, which may be attributed some changes in the monetary policy dealing with the course of narrowly defined monetary aggregates. For the model using M2 monetary aggregate, potential instabilities occur for the crisis-following 1995 and 2002 periods. 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 percent. 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 percent 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. We see that evidence against the parameter constancy verifies the recursive residual estimates obtained above. As with the CUSUM of squares test, movement outside the critical lines is suggestive of parameter or variance instability. We find an outstanding evidence that the 2001 crisis period had been subject to the major parameter instabilities for the model using m1 monetary aggregate. 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 such that major instabilies occur for the pre-2000 period. We should note that the recursive tests for the model using M2 monetary aggregate yield higly similar results to the model using M1 monetary aggregate. Furthermore, the CUSUM of squares test now catchs up the parameter instability for the post-1994 economic crisis period.

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monetary aggregates and relationships in the Turkish economy can also be implemented in future researches. See Engle, Hendry and Richard (1983), Hendry and Ericsson (1991), Favero and Hendry (1992), Bårdsen (1992), Engle and Hendry (1993), Metin (1995), Ghartey (1998) and Cheong (2003) for reconsiderations and applications of this phenomenon in economics literature.

Figure 1: Recursive Estimates (m1 money supply)

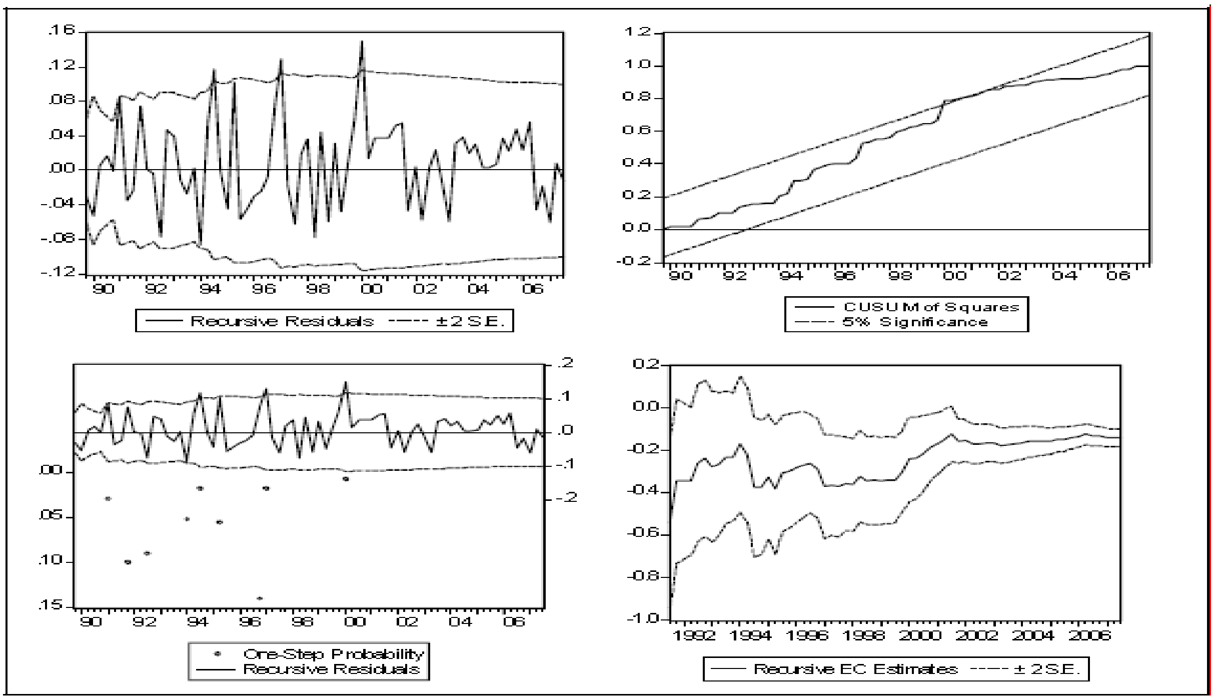
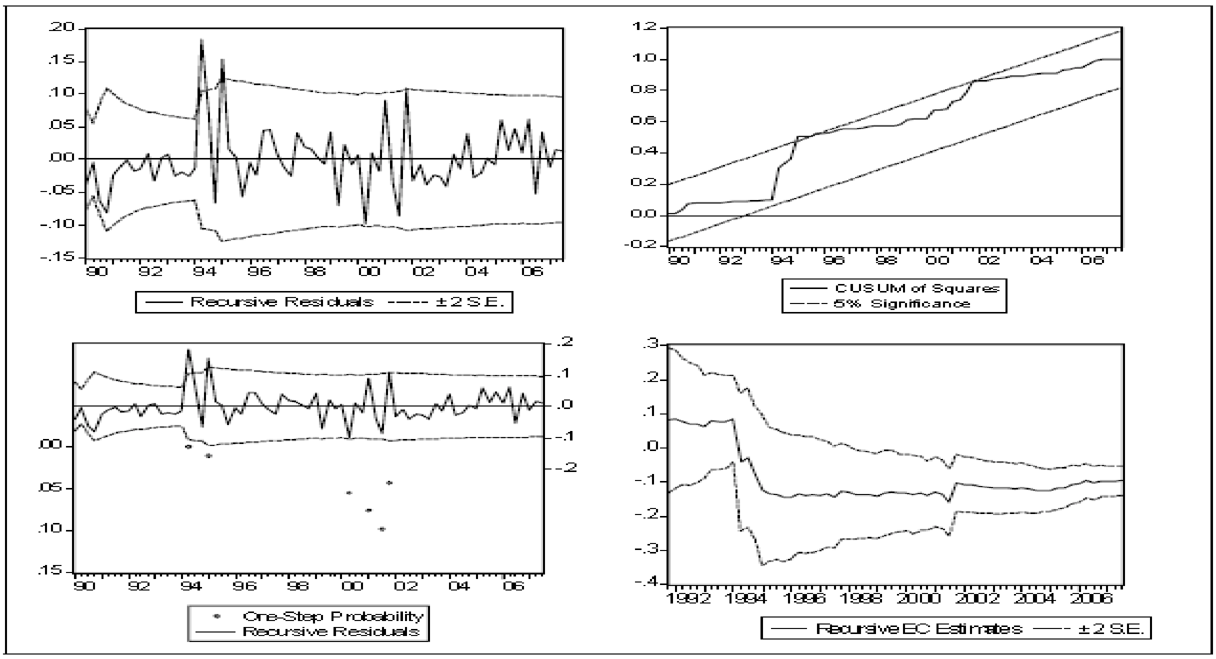


Figure 2: Recursive Estimates (m2 money supply)



5. CONCLUDING REMARKS

The quantity theory of money (QTM) is one of the fundamental building blocks of economics theory and relates mainly prolonged increases in prices to the increases in nominal quantity of money. Based on *a priori* assumption of stability of the functional relations that affect the quantity of money demanded, the QTM assumes that variations in the velocity of money can be foreseen and explained by the economic agents considering a stationary economic relationship for the various phases of business cycles.

In this paper, we examine the validity of stability of long-run relationships between monetary aggregates, prices and real output level in a quantity theoretical perspective for the Turkish economy. Using some contemporaneous econometric techniques, our estimation results exhibit that stationary characteristics of the velocitities of narrowly and broadly defined monetary aggregates cannot be rejected. However, we cannot find both money supply variables as weakly exogenous in the long-run variable space. This requires that money should be taken endogenous for the long-run evolution of prices and real income, thus money cannot be considered the only forcing variable in the multivariate co-integrating system. For the design of monetary policy, a possible explanation can be derived such that given the endogenous characteristics of the monetary variables, monetary authority seems to follow an accommodative monetary policy inside the period. These all would weaken the discretionary policy role of money in the conduct of future stabilization policies. Our estimation results reveal that the changes in the growth rate of M1 and M2 money supplies lead to significant increases in the real output growth rate leading us to reject the (super)neutrality condition of money. Finally, some parameter instabilities and the structural breaks have been attributed to the estimated model especially for the 1994 and 2001 economic crisis periods in the Turkish economy, which require future researches to examine these issues more elaborately.

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