The search for co-integration between money, prices and income: low frequency evidence from the Turkish economy

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The Search for Co-Integration Between Money, Prices and Income: Low Frequency Evidence From the Turkish Economy

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Summary: In this paper, we aim to test the empirical validity of the QTM relationship for the Turkish economy. Using some contemporaneous time series estimation techniques, our estimation results reveal that stationarity characteristics of the velocities of currency in circulation and the broad money aggregate in the economy cannot be rejected through a quantity theoretical co-integrating long-term variable space. We find that there exists an about one-to-one proportionality between money and prices and money and real income, and that the exogeneity of money cannot be rejected for the currency in circulation in the economy. But, the exception here comes from the broad monetary aggregate used in the QTM equation such that money seems to be endogenous as for the long-term variable space.

Key words: Money, Prices, Income, Quantity Theory of Money, Co-integration, Long-span Data, Turkish Economy

JEL: C32, E51, E52, E61

Introduction

Over the past hundred years, the extent and direction of causality and the stability of empirical requlartities among money, income and prices have drawn many economists' attention to the determinants of functional relationships that constitute the fundamental building blocks of the capitalist system. The predictability of the long-run courses of nominal income and prices, and the identification of the role of money as a bridge for the interactions between these macroeconomic aggregates have long been perceived as a prereqisite for the use of stabilization tools in the conduct of monetary policy, given that inferences dealing with monetary policy will meet the stylized facts of the economy only if they succeed in constructing foresights consistent with the behavioral preferences of the economic agents.

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Otherwise, discretionary monetary policies are able to only partially correct monetary disequilibrium stemmed from current macroeconomic framework as well as being not fully justified in a theoretical sense.

In this sense, one of the most essential contributions that relate the course of the monetary aggregates to that of money and income, going back to the initial stages of capitalism as discussed by David Hume (1970), is the *Quantity Theory of Money* (henceforth, QTM) which tries to mainly theorize the role of money in assessing the business cycles characteristics and the *steady-state* long-run course of the aggregate transactions volume. Resurrecting the interest upon the QTM, Friedman (1956) considers the QTM as a stable functional relationship that affects the quantity of money demanded, and such a consideration in turn leads to the additional implication that the causes of variations in the velocity of money can be foreseen and explained by economic agents. Together with a dichotomy assumption that in a *hypothesized* long-run period the volume of real output is likely to be mainly determined by real factors, the tendency for equilibrium in the money market forces the *ex-ante* demand for money balances to have been equalized to the actual supply, and this brings out the importance of money supply as a major determinant of nominal income.

The QTM can be described by the well known exchange identity:

$$M\,V_T = P\,T \tag{1}$$

where M is the money supply, VT 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. In line with the approach of Pigou (1917) and considering the importance of money demand relationship for the implications related to the QTM, Eq. 3 can also be re-written as follows:

$$M / P = k\, Y \tag{4}$$

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 also be consisted of a set of opportunity costs to hold money other than the scale-real income variable. 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, the QTM tends to have been subject to an unstable functional form that destabilizes the implications dealing with the stable velocity of money.

Based on these theoretical fundamentals, some other extensions of the theory can be derived more explicitly. Assume the QTM in terms of the growth rates:

$$m + v = p + y \tag{5}$$

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 the stationary velocity of money and unaffected real output level in the long-run following the permanent changes in the growth rate of money supply. That is, in a more elaborately way to say, real output and velocity changes must be *orthogonal* to the growth rate of the money stock considered (Grauwe and Polan, 2005). Considering all these assumptions, for empirical 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 the velocity of money (v) has been subject to a stationary long-run process.

Among many other papers, Fisher and Seater (1993), King and Watson (1997) and Bullard (1999) examine some theoretical underpinnigs of the QTM relationship. Serletis and Krause (1996) and Serlestis and Koustas (1998) using a low frequency data from ten developed countries over one hundred year give in general support for the long-run neutrality proposition. Karfakis (2002; 2004) and Ozmen (2003) examine the validity of the QTM relationship for the case of Greece and find contradictory results especially for the exogeneity / endogeneity characteristic of the money considered. Herwartz and Reimers (2006) in a panel based paper also try to 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. Finally, a recent paper by Aslan and Korap (2007) upon the Turkish economy supports the stationary characteristics of narrowly and broadly defined monetary aggregates for the post-1987 period till the end-2006, but also find that endogenous characteristics of the monetary aggregates for the long-run evolution of prices and real income cannot be rejected.

In this paper, our aim is to conduct an empirical model using long-span historical data to test the empirical validity of the QTM relationship for the Turkish economy. To this end, the contemporaneous time series techniques have been applied to extract the necessary knowledge of the QTM from the actual data. For this purpose, the next section deals with the preliminary data issues and the third section describes estimation methodology. The results of the empirical model are presented in the fourth section. The last section summarizes results to conclude the paper.

1.Preliminary Data Issues

1.1. Data Definitions

For empirical purposes, the data used in their natural logarithms cover the investigation period 1950-2006 with low frequency annual observations and are taken from the Statistical Indicators (1923-2007) published by the Turkish Statistical Institute (2008). In a quantity theoretical approach, the choice of monetary variable has been of a special importance. Following Lucas (1980), thus, two alternative variable specifiations which correspond to be theoretically termed ‘money’ for different monetary aggregates have been considered, represented by either currency in circulation (*CC*) as a narrow money definiton or broad money (*M2*) as a sum of *CC* plus sight and time deposits denominated in the domestic currency in the banking system. Somewhat supporting the financial development in the economy, note that the proportion of *CC* to *M2* takes a value between 40%-50% in 1950s, 30%-40% in 1960s, 20%-30% in 1970s, 10%-20% in whole 1980s and 1990s and finally about 10% or lower for the post-2000 period. The gross national product (GNP) deflator (*PRI*) 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 (*INC*) has been calculated by dividing nominal GNP to the deflator values. As exogenous variables, we also use, in a co-integrating framework below, a shift dummy variable d80 which takes a zero value before the year 1980 and a unity value after 1980 to represent the enormous economic / financial change in the economy from an inward-looking import-substitution policy to an export-based openness to trade framework, and two impulse dummy variables, d94 and d01, that take a unity value for the economic / financial crisis years 1994 and 2001 and zero otherwise as a proxy of structural diversifications of the Turkish economy.

1.2. Unit Root Tests

Spurious regression problem analyzed by Granger and Newbold (1974) indicates that using nonstationary time series steadily diverging from long-run mean causes to unreliable correlations within the regression analysis leading to unbounded variance process. This is particularly likely to be happened when the adjustment determination coefficient under the impact of correlated trends is found highly larger than the regression Durbin-Watson statistic which can also be resulted from non-stationary residuals. However, for the mean, variance and covariance of a time series to be constant over time, conditional probability distributions of the series must be invariant with respect to the time, and if only so the conventional procedures of OLS regressions can be applied using a stationary process for the variables. Dickey and Fuller (1981) provide one of the commonly used test methods known as augmented Dickey-Fuller (ADF) test of detecting whether the time series data are of stationary form. This can be formulated for any y_t variable as follows:

$$\Delta y_t = \alpha + \beta t + (\rho - 1) y_{t-1} + \sum_{i=1}^{k} \eta_i \Delta y_{t-i} + \varepsilon_t \qquad (6)$$

of which the null hypothesis is the presence of a unit root (ρ=1) against the alternative (trend)stationary hypothesis. For *yt* to be stationary, (ρ-1) should be negative and significantly different from zero. Moreover, while the assumption that y_t follows an autoregressive (AR) process may seem restrictive, Said and Dickey (1984) demonstrate that the ADF test is asymptotically valid in the presence of a moving average (MA) component provided that sufficient lagged difference terms are included in the test regression. The estimated ADF statistics are compared with the simulated MacKinnon (1996) critical values which employ a set of simulations to derive asymptotic results and to simulate critical values for arbitrary sample sizes. For the case of stationarity, we expect that these statistics must be larger than the critical values in absolute value and have a minus sign.

However, conventional unit root tests such as the most widely used ADF estimation procedure tend to be strongly criticized in the contemporaneous economics literature when they have been subject to structural breaks which yield biased estimations. These tests assume that variables can be characterized as a random walk process which requires differencing to achieve a stationary time series. 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 for 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 be coincided with dates chosen exogenously. To address this issue, several methodologies have been suggested to allow for the determination of the date of structural breaks endogenously. Considering these issues, in our paper, we follow the widely used Zivot & Andrews (1992) (henceforth ZA) methodology allowing the data to indicate breakpoints endogenously rather than imposing a breakpoint from outside the system. 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:

$$y_t = \mu + y_{t-1} + \varepsilon_t \tag{7}$$

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 $\lambda = TB / T$, $\lambda \quad |0.15, 0.85|$, by following the augmented regressions:

Model A:

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

Model B:

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

Model C:

$$y_t = \mu + \beta t + \theta DU_t(\lambda) + \gamma DT_t(\lambda) + \alpha y_{t-1} + \sum_{j=1}^{k} c_i \Delta y_{t-j} + \varepsilon_t \tag{10}$$

where *DUt* and *DTt* are sustained dummy variables capturing a mean shift and a trend shift occuring at the break date respectively, i.e., *DUt*(λ) = 1 if $t > T\lambda$, and 0 otherwise; *DTt**(λ) = $t - T\lambda$ 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 distribued (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. We report the ADF and ZA unit root test results in Tab. 1 and Tab. 2 below.

Table 1. ADF Unit Root Test

Variables	in levels		in first differences	
	τ_c^{ADF}	τ_c^{ADF}	τ_c^{ADF}	τ_c^{ADF}
CC	-1.42 (3)	-2.50 (3)	-7.31 (1)*	-7.46 (1)*
M2	1.47 (1)	-1.52 (1)	-1.98 (1)	-3.69 (0)**
INC	-1.64 (0)	-1.61 (0)	-8.62 (0)*	-8.53 (0)*
PRI	-0.93 (3)	-2.07 (1)	-6.88 (1)*	-7.04 (1)*

Notes: τ_c and τ_c are the test statistics for the ADF testswith allowance for only constant and constant&trend terms in the unit root tests, respectively. 1% and 5% critical values are τ_c = -3.56 (1%), τ_c = -2.92 (5%),τ_t = -4.14 (1%) and τ_c = -3.49 (5%). * and ** denote the rejection of the unit root null hypothesis for the 1% and 5% levels, respectively. The numbers in parentheses are the lags used for the adf test, which are augmented up to a maximum of 8 lags. The choice of optimum lag for the ADF test was decided on the basis of minimizing the Schwarz information criterion. ADF unit root test procedure has been implemented in EViews 6.0.

Table 2. ZA Unit Root Test

	Intercept			Trend			Both		
	k	min *t*	*TB*	*k*	min *t*	*TB*	*k*	min *t*	*TB*
CC	2	-3.525	(1990)	2	-3.275	(1980)	2	-3.384	(1979)
M2	2	-2.287	(1997)	0	-3.678	(1979)	0	-3.378	(1980)
INC	0	-4.554	(1970)	0	-3.199	(1978)	0	-3.903	(1978)
PRI	1	-3.703	(1989)	1	-3.347	(1977)	0	-3.174	(1978)

Notes: Estimations with 0.15 trimmed. min *t* is the minimum *t*-statistic calculated.Critical values – intercept: -5.43 (1%), -4.80 (5%); trend: -4.93 (1%), -4.42 (5%); both: -5.57 (1%), -5.08 (5%). ZA unit root test procedure has been implemented in STATA 9.0.

The unit root test results from the ADF equation indicate that the unit root null hypothesis cannot be rejected for all the variables in their levels, but differencing provides stationarity. Note that the differenced form of the M2 money supply tends to be trend-stationary. Therefore, we infer that all the variables have an I(1) characteristic due to the ADF test results. When we consider the ZA unit root test results in Tab. 2 allowing endogenous break in the time series used, we find that no change occurs in the non-stationary characteristics of the variables.

2.Estimation Methodology

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

$$z_t = \Pi_1 z_{t-1} + \Pi_2 z_{t-2} + \ldots + \Pi_k z_{t-k} + \varepsilon_t \tag{11}$$

where ε_t is assumed to follow i.i.d. process with a zero mean and normally distributed N(0, σ^2) error structure and z is (nx1)and the Π_i is (nxn) matrix of parameters. Gonzalo (1994) indicates that Johansen multivariate co-integration methodology performs better than other estimation methods even when the errors are non-normalt distributed. Eq. 11 can be written leading to a vector error correction (VEC) model of the form:

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + \Gamma_2 \Delta z_{t-2} + \ldots + \Gamma_k \Delta z_{t-k+1} + \Pi z_{t-k} + \varepsilon_t \tag{12}$$

$$\Gamma_i = -I + \Pi_1 + \ldots + \Pi_i (i = 1, 2, \ldots, k-1) \tag{13}$$

$$\Pi = I - \Pi_1 - \ldots - \Pi_k \tag{14}$$

Eq. 12 can be arrived by subtracting z_{t-1} from both sides of Eq. 11 and collecting terms on z_{t-1} and then adding -(Π_1 - 1)X_{t-1} + (Π_1 - 1)X_{t-1}. Repeating this process and collecting of terms would yield Eq. 12. This specification of the system of variables carries on the knowledge of both the short- and the long-run adjustment to changes in z_t, via the estimates of Γ_i and Π.

Following Harris and Sollis (2003), $\Pi = \alpha\beta'$ where α measures the speed of adjustment coefficient of particular variables to a disturbance in the long-run equilibrium relationship and can be interpreted as a matrix of error correction terms, and β is a matrix of long-term coefficients such that $\beta' z_{t-k}$ embedded in Eq. 12 represents up to (n-1) co-integrating relations in the multivariate model which ensure that z_t converge to their long-term steady-state solutions.

For the lag length of unrestricted VARs, 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, all the criterions suggest to use lag length 2 for the model using *CC* as a monetary aggregate. For the model using broad money balances *M2* LR, FPE, SC and HQ criterions suggest 2 lag orders to be considered, while the minimized AIC statistics indicate 3 lag orders as the appropriate selection. Thus, we choose the lag length 2 for both unrestricted VAR models. 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 $r+1$ cointegrating relations and trace for the null hypothesis of r co-integrating relations against the alternative of n co-integrating relations, for $r = 0, 1, \ldots, n$-1 where n is the number of endogenous variables.

3. Results

We now report the results of Johansen co-integration test using max-eigen and trace tests based on critical values taken fom Osterwald-Lenum (1992). Following Johansen (1992), for the co-integration test we restrict intercept and trend factors into the long run variable space in line with the Pantula principle, but do not assume a quadratic deterministic trend lying in both the co-integrating model and the dynamic vector error correction model. The rank tests are presented in Tab. 3 and Tab. 4. From the results in Tab. 3, both LR tests used for the *CC* model approve the existence of one potential stationary relationship in the long-term variable space as a co-integrating vector. For the *M2* model, we find in Tab. 4 that trace test indicates one co-integrating relation between the variables of interest while no co-integrating vectors can be detected by the max-eigen test considering 5% critical values. Following these

Table 3. Co-integration Rank Test Results for the *CC* Model

Null hypothesis	r=0	$r\leq1$	$r\leq2$
Eigenvalue	0.3629	0.2292	0.1598
λ-trace	48.681*	23.889	9.5748
5% cv	42.915	25.872	12.518
λ-max	26.792*	14.314	9.5748
5% cv	25.823	19.387	12.518

Notes: Both Trace and Max-eigen tests indicate 1 co-integrating eqn at the 5% level. An asterisk denotes rejection of the hypothesis at the 5% level.

Table 4. Co-integration Rank Test Results for the *M2* Model

Null hypothesis	r=0	$r\leq1$	$r\leq2$
Eigenvalue	0.3777	0.2790	0.1377
λ-trace	44.629*	22.335	6.9635
5% cv	42.915	25.872	12.518
λ-max	22.294	15.372	6.9635
5% cv	25.823	19.387	12.518

Notes: Trace test indicates 1 co-integrating eqn while Max-eigen test indicates no co-integration atthe 5% level. An asterisk denotes rejection of the hypothesis at the 5% level.

findings, therefore, we accept that for both the *CC* and *M2* models one potential co-integrating vector is likely to be lying in the variable space. However, it is possible that some structural breaks may be attributed to the co-integrating relationship especially for a country such as Turkey. Following the suggestions of an anonymous referee, on this issue, in order to test the existence of a co-integrating relationship subject to structural breaks, we also employ the medhodology suggested by Johansen et al. (2000) which can be used to specify up to two structural beaks either in levels or in levels and trend jointly. Here we tend to test the sensitivity of the rank test results obtained above to some exogeneous breaks in levels and trend jointly, allowing trend shift restricted to error correction term and level shift unrestricted in the model. These results assuming only one-break occurred in the military intervention year 1980 and two-breaks coincided with 1994 and 2001 economic / financial crises are reported

Table 5. The Rank Tests for the *CC* Model with Exogenous Breaks

Restricted Dummies 1980 Military Intervention Period

Trend and Intercept Included

Response Surface Computed

r	LR	*p*-val	90%	95%	99%
0	47.46	0.0446	43.40	46.62	53.07
1	18.08	0.4912	25.83	28.41	33.68
2	6.39	0.5359	12.08	14.00	18.08

Restricted Dummies 1994 and 2001 Economic / Financial Crisis Periods

Trend and Intercept Included

Response Surface Computed

r	LR	*p*-val	90%	95%	99%
0	66.22	0.0071	54.15	57.73	64.86
1	28.18	0.2833	33.41	36.32	42.21
2	7.33	0.7431	16.24	18.46	23.10

Table 6. The Rank Tests for the *M2* Model with Exogenous Breaks

Restricted Dummies 1980 Military Intervention Period

Trend and Intercept Included

Response Surface Computed

r	LR	*p*-val	90%	95%	99%
0	60.24	0.0412	55.46	59.09	66.32
1	22.04	0.6994	34.46	37.42	43.40
2	9.83	0.5677	16.79	18.93	23.38

Restricted Dummies 1994 and 2001 Economic / Financial Crisis Periods

Trend and Intercept Included

Response Surface Computed

r	LR	*p*-val	90%	95%	99%
0	59.06	0.0945	58.72	62.58	70.26
1	30.11	0.2956	36.09	39.26	45.70
2	8.57	0.6661	17.62	20.11	25.36

In Tab. 5 and Tab. 6. Note that the critical values as well as the *p*-values are taken from the Johansen trace tests and are obtained by computing the respective response surface estimates in JMulTi 4. In Tab. 5 and Tab. 6, we see that the null hypothesis of one co-integrating vector lying in the long-run variable space cannot be rejected even if some exogeneous known breaks have been assigned to the data.

Having verified the presence of one co-integrating relationship using data from the long-run variable space, to see the properties of these vectors we give the unrestricted co-integrating and relevant adjustment coefficients below:

Table 7. Unrestricted Coefficients for the *CC* Model

Unrestricted Co-integrating Coefficients			
CC	*PRI*	*INC*	*TREND*
3.7419	-4.2374	-2.3592	0.0418
6.2820	-6.3257	-10.293	0.3497
-4.4200	2.9335	-10.563	1.1624
Unrestricted Adjustment Coefficients (‘D’ indicates difference operator)			
D(*CC*)	-0.0154	-0.0399	0.0031
D(*PRI*)	0.0316	-0.0237	0.0014
D(*INC*)	-0.0017	0.0120	0.0230

Table 8. Unrestricted Coefficients for the *M2* Model

Unrestricted Co-integrating Coefficients			
M2	*PRI*	*INC*	*TREND*
10.381	-10.423	-11.979	0.001
-3.2045	2.2879	-10.398	1.046
-1.649	2.739	6.809	-0.607
Unrestricted Adjustment Coefficients (‘D’ indicates difference operator)			
D(*M2*)	-0.0527	0.0069	0.0038
D(*PRI*)	-0.0145	-0.0207	-0.0206
D(*INC*)	-0.0048	0.0244	-0.0104

When we examine the unrestricted co-integrating coefficients for both models, we see that the first vector with the largest eigenvalue seems to be a theoretically plausible QTM vector. Since the Johansen methodology only gives us the unrestricted coefficients that tend to converge to an econometrically identified stationary relationship in a co-integrating vector, some normalizations are needed to be carried out to give the variables economical meaning. Thus, rewriting the normalized equations for both monetary aggregates under the assumption of $r = 1$ yields below (t-stats. are given in parentheses):

$$
\begin{array}{llll}
B_{CC}z_t = CC - 1.132PRI & - 0.631INC & + 0.011TREND & - 1.71 \\
\qquad\qquad\quad (-14.22) & (-1.808) & (1.878) &
\end{array}
\tag{15}
$$

$$
\begin{array}{llll}
B_{M2}z_t = M2 - 1.004PRI & - 1.154INC & + 0.001TREND & + 3.53 \\
\qquad\qquad\quad (-29.28) & (-3.388) & (0.004) &
\end{array}
\tag{16}
$$

Results in Eq. (15) and Eq. (16) give some support to the expected long-run characteristics of the QTM relationship as for the statistical significance and signs of the variables. For both models the price elasticity takes highly close values to unity. To further test price homogeneity, in this sense, we apply to the homogeneity and symmetry restrictions only for the variable *PRI* and find $\chi 2(1) = 1.7178$ (prob. 0.1899) for the *CC* model and $\chi 2(1) = 0.0069$ (prob. 0.9334) for the *M2* model. When these restrictions have been tested for the *INC* variable, we estimate $\chi 2(1) = 0.1277$ (prob. 0.7209) for the *CC* model and $\chi 2(1) = 0.0823$ (prob. 0.7742) for the *M2* model. Finally, we test symmetry and homogeneity restrictions to constitute (-1 1 1) relationship for both price and real income elasticities. Our findings indicate that these restrictions together cannot be rejected for the *M2* model variable space that yields $\chi 2(2) = 0.2272$ (prob. 0.8926), but for the *CC* model we are unable to obtain such a result through $\chi 2(2) = 8.2727$ (prob. 0.0160). Thus, these results give strong support to the QTM assumptions that there exists an about one-to-one proportionality between money and prices and money and real income. We can infer here that the *ex-post* stationary characteristic of the velocity of money should not be rejected in line with a quantity theoretical stable functional relationship. The estimated models also fit well with the diagnostics such that no vector error correction serial correlation problem can be found through LM(1) = 5.9067 (prob. 0.7492) for the *CC* model and LM(1) = 3.6927 (0.9305) for the *M2* model. As a next step, we examine the properties of the adjustment coefficients for each estimated cointegrating model equations. The results are reported in Tab. 9 and Tab. 10 below.

Table 9. Adjustment Coefficients of the Normalized *CC* Model [*t*-stats. in ()]

D(*CC*)	D(*PRI*)	D(*INC*)
-0.0575	0.1185	-0.0063
(-1.1823)	(3.2309)	(-0.1785)

Tab. 10 Adjustment Coefficients of the Normalized *M2* Model [*t*-stats. in ()]

D(*M2*)	D(*PRI*)	D(*INC*)
-0.5471	-0.1507	0.0503
(-4.0499)	(-1.2680)	(0.5149)

In Tab. 9, we observe that the only significant feedback effect of disturbances from the *steady-state* functional form occurs upon the price variable. This is consistent with the exogeneity status of the money in the QTM relationship. Thus, possible vector error correction model for this co-integrating relationship must be constructed upon the price variable. However, for the *M2* model, the exogeneity of money has been rejected, while weak exogeneity of the price and real income variables cannot be rejected. Somewhat supporting the findings of Ozmen (2003) upon the Greece economy, this finding contradicts the QTM assumption that money is the sole forcing variable in the multivariate cointegrating system. All in all, these results must be elaborately considered to appreciate the basic characteristics of the long-term course of the Turkish economy in the sense that the QTM relationship can still provide important knowledge for the relationships between money, prices and income. The exception for the QTM assumptions is that money seems to be endogenous as for the long-term variable space when the broad monetary aggregates have been used in the QTM equation. These latter findings are also somewhat similar to the recent findings of Aslan and Korap (2007) upon the Turkish economy with higher frequency data for the post-1987 period, that give support to the stationary characteristics of the narrowly- and broadly-defined monetary aggregates within a quantity theoretical framework with the only exception that monetary aggregates have been estimated endogenous for the long-term evolution of prices and real income, leading to the inference that money cannot be considered the only forcing variable in the multivariate QTM variable space. Of course, these issues of interest further require one-country time series and multi-national panel studies to control the validity of the QTM assumptions. On this point, for instance, testing the neutrality of money for the Turkish economy, that has not been to the great extent emphasized in this paper, would be highly complementary for the estimation results obtained in this paper.

Concluding Remarks

Over the past hundred years, the extent and direction of causality and the stability of empirical reqularities among money, income and prices have drawn many economists' attention to the determinants of functional relationships that constitute the fundamental building blocks of the capitalist system. In this sense, one of the most essential contributions that relate the course of the monetary aggregates to that of money and income is the *Quantity Theory of Money* (QTM) which tries to mainly theorize the role of money in assessing the business cycles characteristics and the *steady-state* long-run course of the aggregate transactions volume. In this paper, we aim to test the empirical validity of the QTM relationship for the Turkish economy. Using some contemporaneous time series estimation techniques, our estimation results reveal that stationarity characteristics of the velocities of currency in circulation and the broad money aggregate in the economy cannot be rejected through a quantity theoretical cointegrating long-term variable space. We find that there exists an about one-to-one proportionality between money and prices and money and real income, and that exogeneity of money cannot be rejected for the currency in circulation in the economy. But, the exception here comes from the broad monetary aggregate used in the QTM equation such that money seems to be endogenous as for the long-term variable space. We must state that these issues of interest need to be supported by further investigations, and thus one-country time series as well as cross-country panel evidences on more detailed QTM assumptions would help researchers appreciate the generality of the empirical results obtained in this paper. In addition, it will also be complementary to test the sensitivity of our findings by assuming some endogenously determined structural breaks in line with the developments in time series data estimation techniques.

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References

Aslan, O. and L. Korap. 2007**.** "Testing Quantity Theory of Money for the Turkish Economy." *Journal of BRSA Banking and Financial Markets*, 1(2): 93-109.

Bullard, J. 1999. "Testing Long-Run Monetary Neutrality Propositions: Lessons from the Recent Research." *FRB of St. Louis Review*, November/December, 57-77.

Dickey, D.A. and W.A. Fuller. 1981. "Likelihood Ratio Statistics for Autoregressive Time Series with Unit Roots." *Econometrica*, 49: 1057-072.

Dwyer, G.P. and R.W. Hafer. 1999. "Are Money Growth and Inflation Still Related?." *FRB of Atlanta Economic Review*, Second Quarter, 32-43.

Engle, R.F. and C.W.J. Granger. 1987. "Co-integration and Error Correction: Representation, Estimation, and Testing." *Econometrica*, 55: 251-76.

Fisher, M.E. and J.J. Seater. 1993. "Long-Run Neutrality and Superneutrality in an ARIMA Framework." *American Economic Review*, June, 83: 402-15.

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

Gonzalo, J. 1994. "Five Alternative Methods of Estimating Long-run Equilibrium Relationships." *Journal of Econometrics*, 60: 203-33.

Granger, C.W.J. and P. Newbold. 1974. "Spurious Regressions in Economic." *Journal of Econometrics*, 2(2): 111-20.

Grauwe, P.D. and M. Polan. 2005. "Is Inflation Always and Everywhere a Monetary Phenomenon?." *Scand. J. of Economics*, 107(2): 239-59.

Harris, R.I.D and Sollis, R. 2003. *Applied Time Series Modelling and Forecasting*. Wiley.

Herwartz, H. and H.-E. Reimers. 2006. "Long-Run Links among Money, Prices and Output: Worldwide Evidence." *German Economic Review*, 7: 65-86.

Hume, D. 1970. "Of Money." In *Writings on Economics*, ed. E. Rotwein, University of Wisconsin Press. Reprinted in selected essays from Political Discourses, 1752.

Johansen, S. 1988. "Statistical Analysis of Cointegration Vectors." *Journal of Economic Dynamics and Control*, 12: 231-54.

Johansen, S. 1992. "Testing for Weak Exogeneity and the Order of Cointegration in the U.K. Money Demand Data." *Journal of Policy Modeling*, 14: 313-34.

Johansen, S. and K. Juselius. 1990. "Maximum Likelihood Estimation and Inference on Cointegration with Applications to the Demand for Money." *Oxford Bulletin of Economics and Statistics*, 52: 169-210.

Johansen, S., Mosconi, R. and Nielsen, B. 2000. "Cointegration Analysis in the Presence of Structural Breaks in the Deterministic Trend." *Econometrics Journal*, 3: 216-49.

Karfakis, C. 2002. "Testing the Quantity Theory of Money in Greece." *Applied Economics*, 34: 583-87.

Karfakis, C. 2004. "Testing the Quantity Theory of Money in Greece: Reply to Ozmen." *Applied Economics Letters*, 11: 541-43.

King, R.G. and M.W. Watson. 1997. "Testing Long-Run Neutrality." *FRB of Richmond Economic Quarterly*, 83(3): 69-101.

Lucas, R.E., Jr. 1980. “Two Illustrations of the Quantity Theory of Money.” *American Economic Review*, 70(5): 1005-014.

MacKinnon, J.G. (1996). “Numerical Distribution Functions for Unit Root and Cointegration Tests.” *Journal of Applied Econometrics*, 11: 601-18.

Mishkin, F.S. 1997. *The Economics of Money, Banking and Financial Markets*. 5. Ed., Addison-Wesley.

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: 461-72.

Ozmen, E. 2003. “Testing the Quantity Theory of Money in Greece.” *Applied Economics Letters*, 10: 971-74.

Perron, P. 1989. “The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis.” *Econometrica*, 57: 1361-401.

Pigou, A.C. 1917. “The Value of Money.” *Quarterly Journal of Economics*, 32: 38-65.

Said, S.E. ve Dickey, D.A. 1984. “Testing for Unit Roots in Autoregressive Moving Average Models of Unknown Order.” *Biometrika*, 71: 599-607.

Serletis, A. and Z. Koustas. 1998. “International Evidence on the Neutrality of Money.” *Journal of Money, Credit, and Banking*, 30(1): 1-25.

Serletis, A. and D. Krause. 1996. “Empirical Evidence on the Long-Run Neutrality Hypothesis Using Low-Frequency International Data.” *Economics Letters*, 50: 323-27.

Turkish Statistical Institute. 2008. *Statistical Indicators 1923-2007*. Ankara, Turkey.

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: 251-70.