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MODELING BASE MONEY DEMAND AND INFLATION FOR THE TURKISH ECONOMY

TÜRKİYE EKONOMİSİ ÜZERİNE PARASAL TABAN VE ENFLASYON MODELİ

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ABSTRACT: In this paper, a reserve money demand model is tried to be constructed for the Turkish economy. Using contemporaneous multivariate co-integration methodology for the investigation period 1987Q1-2007Q3 of the quarterly observations, we find that the real income elasticity of money demand is highly greater than unity which means that there exists an ongoing monetization process with regard to the increases in the real income in the economy. The most important alternative cost against the real money holdings seems to be the expected depreciation rate of the domestic currency against the exchange rate. Such a finding reveals the importance of currency substitution phenomenon dominated in the economy when the economic agents determine the motives of demand for monetary balances. Furthermore, a critical finding estimated in the paper is that domestic inflation has a weakly exogenous characteristic in the money demand variable space which requires no dynamic error correction model constructed on domestic inflation as a function of the excess money demand taken place under the money market disequilibrium conditions.

Keywords: Money demand ; Inflation ; Currency substitution ; Co-integration

ÖZET: *Bu çalışmada, Türkiye ekonomisi için bir rezerv para talebi modeli oluşturulmaya çalışılmaktadır. Üçer aylık gözlemleri dikkate alan 1987Q1-2007Q3 inceleme dönemi için çağdaş çok değişkenli eş-bütünleşim yöntemi kullanılarak elde ettiğimiz bulgular para talebinin reel gelir esnekliğinin birim değerden oldukça yüksek olduğunu göstermekte ve ekonomide reel gelir sürecindeki artışla ilgili olarak süregelmekte olan parasallaşma olgusunun varlığını ifade etmektedir. Reel para tutumları karşısındaki en önemli almaşık maliyet unsuru yerli paranın döviz kuru karşısındaki beklenen değer kaybı olarak gözükmektedir. Bu tür bir bulguysa iktisadi birimler parasal büyüklük tutumlarıyla ilgili güdülerini belirlerken ekonomide yerleşik para ikamesi olgusunun önemini ortya koymaktadır. Ayrıca, çalışmada elde edilen kritik bir bulgu para piyasası dengesizlik koşulları altında meydana gelen para talebi fazlası üzerine koşullandırılmış dinamik bir hata düzeltme modelini gereksiz kılacak bir şekilde yurtiçi enflasyonun para talebi değişken uzayında zayıf dışsal bir yapıya sahip olmasıdır.*

***Anahtar kelimeler**: Para talebi* ; *Enflasyon* ; *Para ikamesi* ; *Eş-bütünleşim*

1. Introduction

The Turkish economy had been subject to a chronic two-digits inflationary framework over a two decades period until the early-2000s and such an economic framework determined how the decisions of economic agents were constructed in many fields of daily living. By the beginning of 2000, an anti-inflationary stabilization program based on a crawling-peg regime had been tried to be implemented by the policy makers. Although seemed to be successful in bringing inflation down as one-half of the initial level for the first 10 months realization, the subsequent two economic crisis periods led the inflation stabilization program to be failed and the economy witnessed a great slump in real GNP. Among many others, Dornbusch (2001), Eichengreen (2001), Uygur (2001), Alper (2001), Ertugrul and Yeldan (2002), Akyuz and Boratav (2003) and Ekinci and Erturk (2007) criticize the reasons behind the Turkish-2000 stabilization program and examine the developments leading to the collapse of the program. Following such developments, the Turkish economy has still been trying to establish an inflation targeting (IT) framework supported by free-floating exchange rate system and in this way aims at providing forward looking nature of the policy stance as a main characteristic of the IT (Leigh and Rossi, 2002).

There exists a large literature constructed upon the reasons of the Turkish inflation and many papers try to reveal the consequences of different stabilization programs against the domestic inflation. In this respect, Alper and Ucer (1998), Akyurek (1999) and Erlat (2001) point out the importance of inflationary stickiness and expectations phenomenon with a long-memory in Turkish inflation rates. Ozmen (1998) and Koru and Ozmen (2003) find that, in the long-run, inflation appears to determine currency growth and that inflation seems not to be the result of an active monetary policy aiming to maximize seigniorage revenues. Neyapti (1998) also emphasizes the importance of inertia phenomenon on the domestic inflationary framework and indicates that targeting net domestic assets in fighting inflation may not be appropriate for the Turkish economy and suggests to use interest rate policy tool for this purpose. Likewise, Metin-Ozcan, Berument and Neyapti (2004) state that strong inertial nature identifies one of the salient characteristics of the Turkish inflation. Us (2004) attributes the relatively high and inertial nature of the Turkish inflation mainly to the increases in public sector prices and the depreciation of domestic currency and indicates that high prices have not been as a result of expansionary monetary policy, leading to the inference that inertial nature of the Turkish inflation is not a monetary phenomenon. Baydur and Suslu (2004) conclude interestingly that the Central Bank of the Republic of Turkey (CBRT) assisted in the rise of inflation by implementing tight monetary policy from 1987 to 1997 and that it contributed to the fall of inflation by following relatively loose monetary policy after 1997. Besides, they estimate that the CBRT does not have a monopolistic power in controlling inflation rate. Altinkemer (2004) does not support the possibility of monetary targeting for Turkey as well, due to the joint endogeneity characteristics of inflation and real base money. Saatcioglu (2005) estimates that cost-push rather than demand-pull factors led by e.g. exchange rate depreciations and public sector pricing behavior are responsible for the domestic inflationary framework. He concludes that the monetary authority seems obliged to realize accommodative monetary policy because of a chronic inflationary environment and in turn he proposes not to target monetary variables in a stabilization effort in fighting inflation.

In this paper, our aim is to examine whether targeting base money aggregate under the liability of the monetary authority can be considered as an appropriate policy tool to fight domestic inflation in the Turkish economy. For this purpose, a base money demand model as a function of a set of alternative cost variables as well as of a scale-real income variable is contructed to reveal the main characteristics of the demand for base money balances in the economy and to bring out whether information content of disturbances from the steady-state money demand function can be modelled to explain the changes in the domestic inflation. The data and methodological issues are presented in the next section. Section 3 is devoted to applying an empirical base money demand model for the Turkish economy. The last section summarizes results to conclude the paper.

2. Data and Methodology

2.1. Preliminary Data Issues

In this section, a base money demand model is constructed for the investigation period 1987Q1-2007Q3 with quarterly observations. The monetary variable considered is the reserve money aggregate under the liability of the monetary authority. Reserve money aggregate (*res*) is the sum of currency issued, deposits of banking sector as required and free deposits, extrabudgetary funds and deposits of non-bank sector. The real gross domestic product data at constant 1987 prices are used for the real income variable (y). The variables chosen to represent alternative costs to hold base money balances are the annualized quarterly inflation based on GDP-deflator (p), which is calculated in a four-period lagged differenced form in natural logarithms, expected exchange rate depreciation (e), and 12-months weighted time deposit rate (*rtd*). For the expected real exchange rate depreciation, we follow Goldfajn and Valdes (1999) and Civcir (2000) and estimate a regression of trade weighted real exchange rate series based on producer price indices published by the Central Bank of the Republic of Turkey (CBRT), for which an increase means appreciation of the domestic currency, onto a constant and trend. Then the deviation of the actual series form the predicted series is calculated for real exchange rate misalignment and is assumed to represent expected depreciation of domestic currency against the exchange rate.

All the data used are in their natural logarithms with their seasonally unadjusted values except the 12-months weighted time deposit rate and expected exchange rate depreciation data which are considered in their linear forms, and are taken from the electronic data delivery system of the CBRT. 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 included into the model construction as exogenous variables. Under the assumption of no money illusion, the demand for money is used as a demand for real money balances. In this paper, the GDP deflator is used to deflate the money supply.

Spurious regression problem analysed by Granger and Newbold (1974) indicates that non-stationary time series steadily diverging from long-run mean will give biased standard errors with an 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 considered. Dickey and Fuller (1979) provide one of the commonly used test methods known as augmented Dickey-Fuller (ADF) test to examine the non-stationary characteristics of the variables. This test can be formulated as follows:

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

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 η_i. The null hypothesis of the ADF test is the presence of a unit root (ρ = 1) against the alternative stationary hypothesis. We compare the estimated ADF statistics with the simulated MacKinnon (1991, 1996) critical values. For y_t to be stationary, (ρ - 1) should be negative and significantly different from zero:

Table 1. Unit Root Tests

Variable	τ_c	τ_t
res	-2.35 (1)	-3.09 (1)
Δ*res*	-8.45* (2)	-6.49* (3)
y	-0.08 (8)	-2.35 (8)
Δ*y*	-3.08 (7)*	-3.56 (7)*
p	-0.38 (4)	-1.61 (4)
Δ*p*	-6.87 (3)*	-6.87 (3)*
e	-2.85 (0)	-2.92 (0)
Δ*e*	-10.05 (0)*	-10.03 (0)*
rtd	-1.20 (0)	-2.10 (0)
Δ*rtd*	-9.07 (0)*	-9.15 (0)*
5% critical values	**-2.90**	**-3.47**

Above, τ_c and τ_t are the test statistics with allowance for only constant and constant&trend tems in the unit root tests, respectively. 'Δ' denotes the first difference operator, while '*' means that the data are of stationary form. The numbers in parantheses are the lags used for the unit root test and augmented up to a maximum of 10 lags. The choice of optimum lag for the ADF test was decided on the basis of minimizing the Schwarz information criterion. The test statistics indicate that null hypothesis of a unit root cannot be rejected for all the variables in the level form, but that differencing provides stationarity. Moreover, multivariate statistics for testing stationarity obtained from co-integration methodology given below verify these findings.

2.2. Econometric Methodology

We now test for a long-run stationary relationship derived from a money demand variable space and for this purpose the multivariate co-integration and vector error correction (VEC) techniques proposed by Johansen (1988) and Johansen and Juselius (1990) are used. 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:

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

where ε_t is the $N(0, \sigma^2)$ disturbance term, assuming an expected value with a normally distribued zero-mean and constant variance, and z is (nx1) and the Π_i an (nxn) matrix of parameters. Eq. 2 can be rewritten leading to a vector error correction (VEC) model of the form:

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + \Gamma_2 \Delta z_{t-2} + \ldots + \Gamma_{k-1} \Delta z_{t-k+1} + \Pi z_{t-k} + \varepsilon_t \tag{3}$$

where:

$$\Gamma_i = -I + \Pi_1 + \ldots + \Pi_i \ (i = 1, 2, \ldots, k-1) \text{ and } \Pi = I - \Pi_1 - \Pi_2 - \ldots - \Pi_k \tag{4}$$

Eq. 3 can be arrived by subtracting z_{t-1} from both sides of Eq. 2 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. 3 (Hafer and Kutan, 1994). 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 (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. 3 represents up to (n-1) co-integrating relations in the multivariate model which ensure that z_t converge to their long-run steady-state solutions. Note that all terms in Eq. 3 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.

Dealing with the rank conditions, three 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) leading to that Π would be an (nxn) matrix of zeros. In this case, a VAR model consisted of a set of variables in first differences thus carrying no long-run knowledge of any stationary relationship can 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. What is of special interest here is the possibility that there exist r co-integrating vectors in $\beta' z_{t-k} \sim$ I(0) and (n-r) common stochastic trends when Π has reduced rank, i.e., $0 < r \leq (n-1)$. That is, first r columns of β are the linearly independent combinations of the endogenous variables settled in vector z_t, which represent stationary relationships. Whereas, the latter (n-r) columns constitute the non-stationary vectors of I(1) common trends, which require that the last (n-r) columns of α take insignificant values highly close to zero, impeding feedback effects of deviations from 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 dynamics are unknown. Further, this method does not suffer from problems associated with normalisation (Johansen, 1995).

We estimate the long run co-integrating relationships between the variables by using two likelihood test statistics known as maximum eigenvalue for the null hypothesis of r versus the alternative of r+1 co-integrating relationships and trace for the null hypothesis of r co-integrating relations against the alternative of n co-integrating relations, for r = 0,1, ... , n-1 where n is the number of endogenous variables. For the lag length of unrestricted VAR model, the sequential modified LR statistics employing small sample modification and minimized Akaike information criterion (AIC) are considered to select the appropriate model between different lag specifications. For all the models, both LR and AIC statistics suggest to use the maximum lag order so that VAR(5) model is estimated. We also test unit income homogeneity restriction as was used generally in standard money demand models, which constructs a proportional stationary relationship between the long-run courses of real monetary balances and real income in line with a quantity theoretical perspective. Following Johansen (1992) and Harris (1995), 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 eror correction model.

3. Results

The results of Johansen co-integration test are reported in Tab. 2 below 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). The latter 'D' at the beginning of the variables indicates the first difference operator:

Both rank test statistics indicate that there exist two potential co-integrating vectors lying in the long-run variable space. When the unrestricted co-integrating coefficients are examined, the first row with the largest eigenvalue seems to be a standard money demand vector, since all the variables have expected and statistically significant normalized signs with regard to the real base money balances. Therefore, we assume that the first vector respresents the base money demand relationship that we examine in this paper.[1] Estimation results reveal that the real income elasticity of the real money balances is highly above the unity that indicates an increasing ongoing monetization process in the economy for the period under investigation. Moreover, the unit income elasticity homogeneity restriction which requires a proportional relationship between real base money balances and real income through a quantity theoretical perspective is rejected in line with the LR test results. The main alternative cost variable against holding money balances seems to be the variable that represents the expected depreciation of the domestic currency against the exchange rate, which is the most significant alternative cost and has also the largest elasticity among the alternative costs, and this result brings out the importance of currency phenomenon when the economic agents determine the motives that determine demand for money. As for the co-integrating model adjustment coefficient of the real money balances, we find that nearly 6.3% of the adjustment in the money demand disequilibrium conditions to the long-run equilibrium is realized within one-period.

A critical finding in Tab. 2 is that domestic inflation is weakly exogenous in the money demand variable space since the adjustment coefficient of domestic inflation is found statistically insignificant. The weakly exogenous characteristic of inflation implies that money demand equations should not be appreciated as price equations (MacKinnon and Milbourne, 1988). 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 means explicitly that the main factors leading to the domestic inflation are determined out of the money demand variable space used in this paper.

[1] An alternative methodology might be estimating a linear combination of the two vectors, which also represents a stationary relatonship, but we do not follow such a methodology, since the first vector alone satisfies the *a priori* expectations with regard to a standard money demand equation.

Table 2. Money Demand Co-integrating Model

Null hypothesis	r=0	r≤1	r≤2	r≤3	r≤4
Eigenvalue	0.90	0.47	0.18	0.12	0.05
λ-trace	249.45	77.01	28.21	13.24	3.68
5% cv	88.80	63.88	42.92	25.87	12.52
prob.	0.00	0.00	0.61	0.72	0.79
λ-max	172.44	48.80	14.97	9.56	3.68
5% cv	38.33	32.12	25.82	19.39	12.52
prob.	0.00	0.00	0.64	0.67	0.79
Unrestricted Co-integrating Coefficients					
res	*y*	*p*	*e*	*rtd*	*trend*
0.9052	-7.8859	0.8696	3.8686	0.8996	0.0797
-2.4341	-0.5704	-7.9247	-2.5285	1.6375	-0.0302
-2.1124	1.6375	-2.0416	15.597	3.5172	0.0145
4.9955	-0.1896	6.0038	-0.39263	2.6539	0.0261
-3.6913	0.1005	-3.2259	-3.1283	-0.3422	0.0342
Unrestricted Adjustment Coefficients					
res	-0.0699	-0.1968	0.0071	-0.00334	0.0284
y	0.2357	0.0006	-0.0061	-0.0078	0.0051
p	-0.0359	0.2134	-0.0058	-0.0108	-0.0137
e	-0.0119	0.0027	-0.0232	0.0075	0.0061
rtd	-0.0380	-0.0350	-0.0006	-0.0630	-0.0250
1 Co-integrating Equation					
Normalized co-integrating coefficients (std. errors in parantheses)					
res	*y*	*p*	*e*	*rtd*	*trend*
1.0000	-8.7115	0.9607	4.2736	0.9937	0.0881
	(0.3589)	(0.2395)	(0.7662)	(0.2155)	(0.0044)
Adjustment coefficients (std. errors in parantheses)					
D(*res*)	D(*y*)	D(*p*)	D(*e*)	D(*rtd*)	
-0.0632	0.2133	-0.0352	-0.0108	-0.0344	
(0.0314)	(0.0096)	(0.0347)	(0.0072)	(0.0243)	
Multivariate Statistics for Testing Stationarity					
	res	*y*	*p*	*e*	*rtd*
χ^2 (4)	161.33	31.39	151.65	163.43	163.50
Unit Income Homogeneity Restriction					
b(1,2) = -1 χ^2(1)=101.19					
Vector Diagnostic Tests					
Vector Error Correction (VEC) Residual Serial Correlation LM tests					
Lags	LM-Stat	prob.			
4	17.6582	0.8566			
VEC Residual Normality Tests					
Skewness χ^2 (5)	64.5148	prob.	0.0000		
Kurtosis χ^2(5)	329.7196	prob.	0.0000		
Jarque-Bera (10)	394.2344	prob.	0.0000		
VEC Residual Heteroskedasticity Tests					
χ^2 (780) = 744.1614					

However, as Civcir (2000) states, excess money derived from a standard money demand equation should have a positive significant effect on the inflation. Therefore, base money aggregate under the control of monetary authority should not be considered a forcing factor for the long-run evolution of domestic inflation, and for the design of monetary policy such an inference in turn would weaken the discretionary policy role of base money in the conduct of stabilization policies against inflation within the period examined. We can conclude that stabilization efforts based on monetary targeting using base money aggregate under the liability of the CBRT may not be consistent with *ex-ante* policy purposes to lower the inflationary inertia phenomenon settled in the economy. Whereas, any stabilization effort based on monetary targeting against inflationary framework requires that a stable money demand functional form can be constructed to monitor the effects of base money growth on the changes in domestic inflationary framework and that long run knowledge resulted from money demand equation can be included into the determination of inflation.[2] But the latter inferences cannot be fulfilled for the Turkish economy leading to that monetary targeting on the policy variable base money aggregate cannot be the appropriate policy choice to fight domestic inflation.

Finally, our co-integrating model has good diagnostics except the violation of the VEC normality condition. However, we omit this problem through Gonzalo (1994).

4. Concluding Remarks

One of the predominant characteristics of the Turkish economy over a two decades period is the high inflationary framework settled in the economy, and such a case constitutes an important benchmark for economic agents in constructing their expectations as to the future periods. Following the collapse of anti-inflationary stabilization program witnessed in the early-2000s, the Turkish economy has still been trying to establish an inflation targeting framework supported by free-floating exchange rate system.

In this paper, a base money demand model is tried to be constructed for the Turkish economy to examine whether targeting base money aggregate under the liability of the monetary authority can be considered an appropriate policy tool to fight domestic inflation in the Turkish economy. Using contemporaneous multivariate co-integration methodology for the investigation period 1987Q1-2007Q3 of the quarterly observations, we find that the real income elasticity of money demand is highly greater than unity which means that there exists an ongoing monetization process with regard to the increases in the real income in the economy. The most important alternative cost against the real money holdings seems to be the expected depreciation rate of the domestic currency against the exchange rate. Such a finding reveals the importance of currency substitution phenomenon dominated in the economy when the economic agents determine the motives of demand for monetary balances. Furthermore, a critical finding estimated in the paper is that domestic inflation has a weakly exogenous characteristic in the money demand variable space which requires no dynamic error correction model constructed on domestic inflation as a function of the excess money demand taken place under the money market disequilibrium conditions. Of course, future papers upon the long-run stationary relationships between narrowly / broadly defined monetary aggregates and domestic inflation considering also a large set of alternative costs and policy variables and the studies revealing the extent to which the structural breaks and the parameter instabilities are occured in the money demand variable space will be complementary in order to verify the estimation results obtained in this paper.

[2] We do not report estimation results for the stability tests of the money demand equation to save space. But, these results are available from the authors upon request.

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