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THE ENDOGENOUS MONEY GROWTH: AN OUTCOME OF HIGH BUDGET DEFICITS IN TURKEY*

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Abstract:

This study is an attempt to reveal the characteristics of the adjustment dynamics of the monetary transmission process in the Turkish economy. The estimations based on the cointegration analysis of time series confirm the presence of fiscal dominance over monetary dynamics resulting from continuously increasing borrowing requirements of the public sector in the period 1985-2001. The empirical findings imply that the monetary transmission process is characterised by the monetary accommodation of high inflation and budget deficits, the outcome of which is an endogenous credit-money expansion in the economy.

Özet:

Türkiye'deki Yüksek Bütçe Açıklarının Bir Sonucu: İçsel Para Büyümesi

Bu çalışma, Türkiye ekonomisinde parasal aktarım sürecininin uyum dinamiğinin özelliklerini ortaya çıkarma amacına yönelik bir denemedir. Zaman serilerinde eş-bütünleşme analizine dayanan tahminler, 1985-2001 döneminde, sürekli artan kamu iç borçlanma gereğinin bir sonucu olarak ortaya çıkan parasal dinamikler üzerindeki mali egemenliğin varlığını doğrulamaktadır. Ampirik bulgular, parasal aktarım sürecinde, para arzının yüksek enflasyon ve bütçe açıklarına eşlik etmesinin bir sonucu olarak, kredi-para arzının ekonomide içsel genişlediğini ortaya koymaktadır.

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Keywords: Endogenous credit-money, budget deficits, cointegration, Turkey.
Anahtar Sözcükler: İçsel kredi-para, bütçe açıkları, eş-bütünleşme, Türkiye.

INTRODUCTION

The persistently increasing rates of inflation and growing budget deficits have weakened the success of a monetary policy for years in Turkey. Average rates of inflation and nominal interest rates stuck above 50 % accompanied by a debt-GNP ratio doubling every five years since 1985, which appears to cause continuing monetary accommodation so that the debt rollover is maintained without increasing the budget costs of domestic borrowing. The money supply growth after the mid 1980s due to the accommodative behaviour of the monetary authority cannot be explained by the direct monetisation of the deficits by the central bank resources. Rather, it is characterised by indirect monetisation, which occurs through the monetary authority's draining reserves towards commercial banks to enable debt sales to the banking system. The outcome has been considerable expansion in the assets and liabilities of commercial banks, which in fact means credit-money growth in the economy. In this sense, money market activities have been extensively dominated by fiscal dynamics from the mid 1980s onward. The ratio of the domestic debt (the total stock of treasury bills and government bonds) to M2Y, which is considered to be a measure of the fiscal pressure on financial markets, reveals the gradually increasing dominance of the debt finance dynamics between 1985 and 2001. In 2001, this ratio peaks at 115%, almost having doubled with respect to the ratio of the previous year, signalling that the end is near in the debt rollover process.

Under the economic profile presented above, it is worth investigating the characteristics of the macroeconomic transmission process in the Turkish economy in order to reveal the dynamics behind the change in key macroeconomic variables. This may be essential for choosing the optimal policy design in achieving macroeconomic stability. Precise knowledge about the transmission process allows policy makers to gain accuracy in forecasting the reaction of the policy instruments to particular shocks, thereby allowing the formation of good working monetary and fiscal policy rules that increase the credibility of policy makers and future policies.

This study is an attempt to empirically investigate the characteristics of the monetary transmission mechanism in the Turkish economy. The plan of the study is as follows: Following this introduction part, in Part II there is an overview of the link between the money supply process and the budget deficits, along with a review of the relevant literature. Part III includes an empirical investigation to test the endogenous credit-money hypothesis by modelling a simple macroeconomic transmission process. The empirical investigations, which exploit a quarterly data set covering the period 1985: 1-2001: 4, are based on the time series evidence provided by Johansen's cointegration

approach. The last part concludes the study with the policy implications of the findings.

1. MONEY SUPPLY UNDER HIGH BUDGET DEFICITS

In an environment where the macroeconomic transmission process is dominated substantially by the effects of persistent high budget deficits, the central bank's main concern becomes providing government finance at lower costs and maintaining the stability of financial markets. In such conditions, the monetary authority has a passive role and cannot exert any restrictive power on monetary aggregates. The money supply becomes endogenously determined.

The endogeneity of the money creation process is enhanced as commercial banks become the major customers of the government debt instruments. This is nothing but the monetisation of deficits through banks' creation of deposit liabilities against their government-bond acquisitions, instead of a direct monetisation by the monetary authority. In this regard, Goodhart (1995: 251) argues that under massive government deficits, unless the public debt can be sold to non-bank borrowers, there would be "…excessive build-up of bank liquidity, and multiple expansion of loans and deposits" as the banking system finances the public sector debt. Also, Moore (1985: 15) states precisely,

> "*Whenever the increase in the money stock is a by-product of increased borrowing from the banking system, whether by the public or the private sector, the increase in the supply of money is a consequence of increased loan expenditure, not the cause of it, which the central bank can influence only indirectly by changing short-term interest rates. Both the high-powered base and the money stock are then in fact endogenous.*"

The money creation process under high budget deficits can as well be characterised as an endogenous credit-money expansion rather than a monetary expansion to maximize seignorage revenue. This implies that high fiscal deficits, which are financed substantially by the banking system, reduce the likelihood of conducting any restrictive monetary policy. Thus, it may be argued that under such conditions the attempt to control monetary aggregates cannot be effective as a monetary policy action. Money growth can accommodate the inflation as well as inflation can be the outcome of the excessive money growth.

The empirical literature investigating the relationships among deficits, money growth and inflation presents no consensus on the direction and the

significance of the causality between the variables in question. Hamburger and Zwick (1981) and Allen and Smith (1983) find evidence of fiscal influences on money supply growth in the U.S. economy while McMillin and Beard (1982), Joines (1985) and King and Plosser (1985) reject this evidence. The multi-country analyses on this subject performed by Barnhart and Darrat (1988), Haan and Zelhorst (1990), Karras (1994) and Sikken and Haan (1998) report no evidence of monetary accommodation of budget deficits. Hondroyiannis and Papapetrou (1997) argue that budget deficits in Greece lead inflation indirectly through the money growth resulting from high budget deficits. The recent research conducted by Favero and Spinelli (1999) and Fratianni and Spinelli (2001) explore fiscal dominance in Italy from a historical perspective. The findings are in favour of the fiscal dominance hypothesis as deficits are found to affect money supply growth significantly and inflation is found to be influenced by both money supply and government deficits.

In the context of studies on the Turkish economy, recognising the differences in their sample periods and data frequencies, there are conflicting results in identifying the relationships among budget deficits, money supply growth and inflation. Ülengin (1995) finds reserve money growth is inflationary and affected by budget deficits. Metin (1998) argues that budget deficits increase inflation immediately while the monetisation of deficits increases inflation with a delay. Using the cointegration analysis under the assumption of the long-run money neutrality, Akçay, Alper and Özmucur (1996) find a significant impact of budget deficits on inflation, whereas in the same study, by using unrestricted Vector Autoregression estimation, they find that inflation is affected neither by budget deficits nor by money supply growth. In a later study by Akçay, Alper and Özmucur (2001), having found no permanent impact of budget deficits on inflation, they suggest that the change in the PSBR is suggested to be a better indicator of the effects of high budget deficits on inflation. Exploiting the weak-exogeneity testing facility of Johansen's cointegration analysis Özmen and Koru (2000) conclude that the money supply growth and inflation are jointly determined variables, and that there is no evidence for the direct effect of budget deficits on inflation. However, it is found that budget deficits exogenously determine the credit-money expansion, which is consistent with the outcome of the reserve accommodation that enables debt sales to commercial banks. Another empirical study by Özmen (1998: 551) highlights that inflation in Turkey is not the result of monetary expansions to maximize the seignorage revenue; rather monetary accommodation "…has allowed the government to keep real seignorage revenue relatively constant in the face of accelerating inflation."

2. THE EMPIRICAL ANALYSIS

2.1. The Methodology

An I(1) cointegration analysis is going to be used in the empirical investigations, which is based on the maximum likelihood approach proposed by Johansen (1988, 1995) and Johansen and Juselius (1990). In this framework, a cointegrated-VAR with k-lags is denoted in a vector error-correction mechanism (VECM) form for the endogenous variables as:

$$\Delta Z_t = \Pi Z_{t-k} + \sum_{i=1}^{k-1} \Gamma_i \Delta Z_{t-i} + \mu + \psi D_t + \varepsilon_t, \qquad t = 1,....,T; \quad i = 1,...., k-1$$

where $Z_t = (z_1, z_2,..., z_n)'$ is an n×1 vector of the n endogenous variables. Γ and Π are parameter matrices that reflect the short-run and long-run dynamics respectively. μ is a vector of constants, D_t denotes centred seasonal dummies and ε_t is an independent identically distributed error term. $\Pi_{n\times n}$ is the long-run coefficient matrix, which can be decomposed into r distinct cointegrating vectors of $\beta'_{r\times n}$ and an adjustment matrix $\alpha_{n\times r}$ so that it can be denoted as $\Pi = \alpha\beta'$. The number of r linearly independent columns in Π corresponds to the number of the cointegrating vectors, which is known as the cointegrating rank. The elements of an α-matrix indicate the speed of adjustment of a variable in case of a disturbance to the long-run equilibrium relation, while the elements of a β-matrix indicate the long-run responses of the variables in the equilibrium relation. Note that in the cointegration analysis a cointegration vector represents the long-run equilibrium relation among the variables in question.

2.2. Data and the Definition of the Variables

The data set exploited in the cointegration analysis covers a quarterly sample period of 1985:1-2001:4. The list of the definitions of the endogenous variables exploited in the cointegration analysis is as follows:

M : M2Y (TL-billions)
P : Consumer Price Index (1987=100)
y : Real GNP (with 1987 buyers prices) (TL-billions)
R : 3-month deposit interest rate (weighted monthly average)
DEF: Consolidated Budget Cash Balance (monthly average) (TL- billions)

The variables M and P are utilized in the form of rate of growth, i.e., the first-difference of the variable in natural logarithm (ln). In this regard, **ΔM** denotes the rate of growth of the money stock while **ΔP** is the rate of consumer price inflation. The consolidated budget cash balance (DEF) is used as a ratio of nominal GNP, denoted by **d.** However, because the variable DEF is a cumulative sum of the monthly budget cash balances calculated on a yearly basis, this variable is recalculated as a cash balance realization between two consecutive periods of a year rather than their cumulative sums. Finally, the interest rate variable (R) is transformed as ln(1+Rate/400) and denoted by "**r**". The annual interest rate R is divided by 400 in order to make the estimated coefficients comparable with logarithmic quarterly changes while taking the natural logarithm of (1+R/400) is nothing but an approximation to R/400.

All data are obtained from the database of the Central Bank of the Republic of Turkey (CBRT).

2.3. The Empirical Model

In the empirical analysis, in order not to extend the size of the model due to the rather small sample size, variables such as exchange rate and foreign interest rates are not included in the analysis. As stated by Juselius (1996: 796), in general, the long-run results are still valid despite the absence of some variables, "because the cointegration property is invariant under changes in the information set…"

Table 1. ADF Unit Root Test Statistics

	Levels ADF (k)		1st differences ADF (k)
	without trend	with trend	without trend
ΔM	-2.892 (3)	**-7.024** (0)	**-9.087** (1)
ΔP	-2.762 (3)	-2.641 (3)	**-8.797** (1)
y	-1.765 (4)	-1.846 (4)	**-3.717** (5)
r	-2.895 (0)	-3.446 (0)	**-7.335** (1)
d	-0.252 (3)	-2.797 (3)	**-4.804** (4)

Notes: Bold values of the computed ADF t-statistics indicate statistical significance at the 5 % level, based on the critical values of MacKinnon (1991). The figures in parentheses denoted by k are the significant lag lengths at which there is no serial correlation.

As a preliminary step to the estimations, the integration order of the variables **ΔM**, **ΔP**, **y**, **r** and **d** is investigated by testing the *non-stationarity* hypothesis with the Augmented Dickey-Fuller (1981) (ADF) test. The test results in Table 1 indicate that each of the five variables in question is generated by an I(1) process except for **ΔM** of which the data generation process includes a deterministic trend. However, regarding the small sample problems of the unit root tests[1], univariate integration order relevant for a valid I(1) cointegration analysis should better be decided by a multivariate *stationarity test*[2].

2.3.1. The Unrestricted VAR Estimations: The model to begin with can be represented by a vector of endogenous variables such as $Z_t = (\Delta M, \Delta P, y, r, d)'$. The modelling process begins with an unrestricted VAR(k) system with a constant, three centred seasonal dummies and a deterministic trend. To determine the appropriate lag length of the VAR(k) model, a sequential test of system reduction from VAR(5) to VAR(0) is performed[3]. At the 5 % significance level, the reduction from VAR(5) to VAR(4) and from VAR(4) to VAR(3) is not rejected according to the LR statistics F(25,109)=0.59 and F(25,127)=1.47 respectively, while the reduction from VAR(3) to VAR(2) is rejected with a test statistic of F(25,146)=1.92. However, a lag length k=3 has no economic meaning in quarterly data. Therefore, the reduction process is continued from VAR(2) to VAR(1), which is not rejected with a statistic F(25,164)=1.53, whereas the reduction from VAR(1) to VAR(0) is rejected strongly with a statistic F(25,183)=9.07. This means the VAR(1) model is data-acceptable. The validity of the VAR(1) model is also confirmed by the Schwarz (SC) and Hannan-Quinn (HQ) criteria[4]. However, the residual diagnostics of the system alert to both heteroscedasticity and serial correlation in the residuals of the VAR(1) model[5]. As a solution to this problem, the lag length is increased by one (from k=1 to k=2), which seems to result in plausible residual diagnostics as can be observed in Table 2.

Table 2. The Univariate and Multivariate Misspecification Tests

	ΔM	ΔP	y	r	d	System
$F_{AR4}(4,47)$	2.43	0.90	0.94	2.15	0.13	
$F_{ARCH4}(4,43)$	0.14	0.07	0.23	0.58	0.42	
$F_{HET}(22,28)$	0.42	0.39	0.51	0.69	1.19	
$\chi^2_{NORM}(2)$	**18.1**	**27.5**	0.47	3.87	**9.78**	
$F_{AR4}(100,136)$						1.16
$F_{HET}(330,221)$						0.62
$\chi^2_{NORM}(10)$						**42.2**
σ	0.05	0.04	0.04	0.02	0.02	
R^2	0.63	0.72	0.98	0.63	0.63	
Skewness	0.62	1.44	0.02	-0.22	0.46	
Excess Kurtosis	4.70	8.01	2.75	3.59	5.02	

Notes: Bold values of the computed test statistics indicate statistical significance at the 5 % level. F_{AR4} is an LM test for auto-correlated residuals (see Godfrey, 1988). F_{ARCH4} is an LM test for fourth-order ARCH effects (see Engle, 1982). F_{HET} is the White (1980) test for heteroscedasticity. χ^2_{NORM} is a test for normality (see Doornik and Hansen, 1994). σ and R^2 are the residual standard deviation and the 'goodness of fit' measure of each equation, respectively. 'Skewness' and 'excess kurtosis' are given to observe their relative effects on non-normality.

The diagnostic checking of the VAR(2) model performed by various misspecification tests indicates only a non-normality problem in the residuals of the system and some of the individual equations. However, as seen in Table 2, the non-normality seems to be due to excess kurtosis but not to skewness. In this regard, with reference to Gonzalo's (1994) findings that Johansen's maximum likelihood approach appears robust to excess kurtosis, the non-normality can be disregarded in the cointegration analysis.

2.3.2. Cointegration Analysis: The VAR(2) formulation, which is found to be a valid approximation of the data generation process, can now be exploited in testing the presence of cointegrating relationships among the variables **ΔM**, **ΔP**, **y**, **r** and **d**. Additional to the variables, the cointegration space includes a constant and three centred seasonal dummies entering unrestrictedly and a deterministic trend entering restrictedly. The number of the cointegrating vectors, that is, the rank of the cointegration, is determined through the *trace test statistic*[6].

Table 3. The Test for the Cointegration Rank

λ	0.586	0.406	0.276	0.218	0.078
H_0	r=0	r≤1	r≤2	r≤3	r≤4
λ_{trace}	**135.4**	**77.26**	42.61	21.61	5.37
(p)	(0.00)	(0.00)	(0.05)	(0.16)	(0.55)

Notes: λ denotes the eigenvalues while λ_{trace} is the trace test statistic with the probability values given below in parentheses to test the H_0 hypotheses of the cointegration rank. Bold values of the trace test statistics denote statistical significance at the 5 % level.

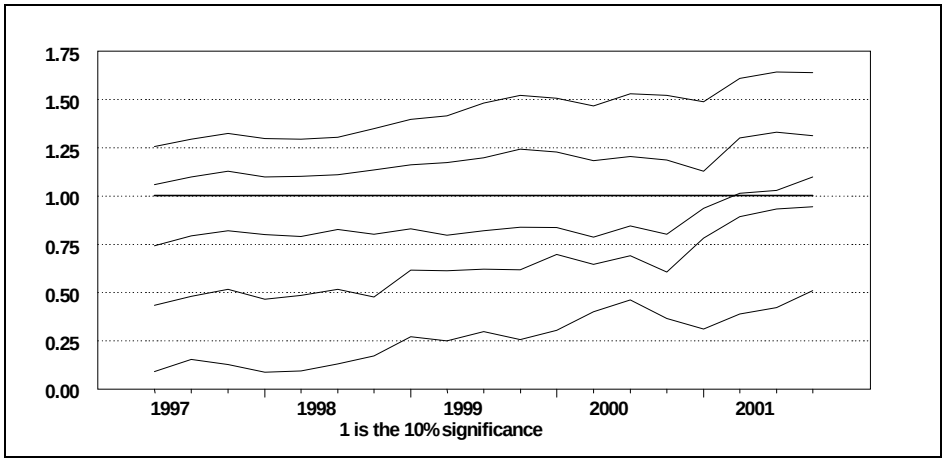

Figure 1. Plots of the Recursive Trace Tests

Table 3. exhibits the results of the Johansen estimation procedure, without any restrictions on the cointegration space. According to these results, the null hypothesis of r≤2 can hardly be rejected at the 5% significance level. For borderline cases, Hansen and Juselius (1995) suggest to look at the graphical behaviour of the estimated cointegration relations before deciding about the cointegration rank. In this respect, plots of the recursively estimated trace tests against time could be used as an auxiliary tool. Figure 1 shows these plots resulting from the recursive estimation over the sub-period 1997: 2-2001: 4. As stated by Hansen and Johansen (1993), the number of the upward sloping time paths over the critical value line indicates the cointegration rank. Therefore, the cointegration rank can be assumed as r=2.

Table 4. The Unrestricted Cointegration Analysis

Standardised Eigenvectors β′					
Variable	β_1	β_2	β_3	β_4	β_5
ΔM	**0.752**	***1.000***	0.444	**1.052**	-0.170
ΔP	-0.061	**-1.081**	*1.000*	**0.856**	-0.343
y	**0.271**	-0.132	-0.434	***1.000***	-0.396
r	**-0.683**	-0.299	-1.030	**-6.031**	*1.000*
d	***1.000***	**-1.199**	-0.156	**-2.487**	0.553
Trend	**-0.004**	0.003	0.005	**0.001**	0.009
Standardised Adjustment Coefficients α					
Equation	α_1	α_2	α_3	α_4	α_5
Δ^2M	**-0.916** (-4.554)	**-0.412** (-2.656)	**-0.508** (-2.712)	**0.117** (2.236)	-0.030 (-0.567)
Δ^2P	0.225 (1.620)	**0.382** (3.576)	**-0.504** (-3.900)	0.042 (1.163)	-0.008 (-0.234)
Δy	**-0.449** (-3.266)	0.116 (1.099)	**0.258** (2.013)	**-0.081** (-2.261)	-0.055 (-1.533)
Δr	-0.032 (-0.547)	0.012 (0.280)	0.055 (1.026)	**0.060** (3.975)	-0.011 (-0.729)
Δd	**-0.559** (-7.142)	**0.152** (2.524)	0.088 (1.208)	-0.011 (-0.558)	0.023 (1.154)
Long-run Exclusion Test					
rank	ΔM	ΔP	y	r	d	Trend
r=1	**11.67**	0.09	**8.20**	**6.12**	**12.46**	**11.25**
r=2	**22.31**	**6.08**	**8.83**	**6.44**	**24.69**	**14.16**
Long-run Weak Exogeneity Test						
Rank	ΔM	ΔP	y	r	d	
r=1	**9.32**	1.15	**6.73**	0.19	**18.49**	
r=2	**13.38**	5.43	**7.56**	0.23	**28.93**	
Multivariate Stationarity Test						
rank	ΔM	ΔP	y	r	d	
r=1	**40.10**	**45.13**	**51.83**	**52.41**	**43.00**	
r=2	**21.97**	**21.28**	**28.15**	**28.55**	**26.55**	

Notes: Bold values indicate statistical significance at the 5 % level. The figures in parenthesis under the α-coefficients are the t-statistics to test the speed of adjustment. The exclusion and the weak-exogeneity tests, and the stationarity test are LR tests with a χ^2distribution, with r degrees of freedom for the first two tests and p-r degrees of freedom for the last test. In testing the significance of the βs, the long-run exclusion test is used, by taking the differences of the LR statistics computed at consecutive cointegration ranks.

Under the assumption that there are two cointegrating vectors, all variables included in the cointegration space are found to have significant explanatory power in the long-run information set, with respect to the long-run exclusion tests in Table 4. The computed results of the multivariate stationarity test indicate the non-stationarity of the variables, confirming the unit root test results obtained before. The rejection of the long-run weak-exogeneity of **ΔM**, **y** and **d** for the parameters of interest reveals that the first two cointegrating vectors (β_1 and β_2) shown in Table 4. should reflect long-run economic relationships normalised for two of these three variables. In this regard, the first vector is normalised to reflect a long-run relationship explaining the deficit/GNP ratio (**d**) while the second is a relationship explaining the broad money growth (**ΔM**). The statistical significance of the adjustment coefficients attached to these variables in the first two vectors proves them to be error correcting. However, with respect to the statistical significance of the β-coefficients in these two vectors, some of the variables can be excluded from the cointegration vectors, as will be done below.

2.3.3. Economic Inferences from the Cointegration Analysis: The findings above require a closer look at the cointegrating vectors and the weak-exogeneity test results in order to reveal the characteristics of the monetary transmission process. In this respect, the weak-exogeneity of the consumer price inflation and the endogeneity evidence of the money supply growth invalidate the quantity-theoretic approach to inflation. It indicates that the broad monetary aggregate (M2Y) cannot be assumed to be under the control of the monetary authority. Moreover, the endogeneity of the budget deficits may be considered to be evidence of a limited scope for a fiscal policy control on budget deficits.

To investigate the long-run economic relations that the cointegrating vectors reflect, the normalised vectors can be analysed in terms of their estimated long-run response and short-run adjustment coefficients. However, more precise coefficient estimates can be obtained by excluding the variables that have insignificant long-run responses in the cointegrating vectors. This also provides a testable over-identified structure. The restricted model estimates are exhibited in Table 5. The over-identifying exclusion restrictions imposed on the model appear to be data-acceptable with respect to the LR test, given in Table 5. The negative and significant adjustment coefficients attached to the identified long-run relationships exhibit error-correcting behaviours, which justifies the vector normalisations in the analysis. Therefore, the cointegrating vectors are worth interpreting as plausible economic relationships.

Table 5. The Restricted Cointegration Vectors

Restricted Eigenvectors β'			**Adjustment Coefficients α**		
Variable	β_1	β_2	Equation	α_1	α_2
ΔM	**0.630** (5.625)	***1.000*** (0.000)	Δ^2M	**-0.735** (-2.964)	**-0.551** (-2.805)
ΔP	0.000 (0.000)	**-1.084** (-0.818)	Δ^2P	0.021 (0.117)	**0.472** (3.404)
y	**0.256** (3.821)	0.000 (0.000)	Δy	**-0.544** (-3.234)	0.136 (1.019)
r	**-0.590** (-3.471)	0.000 (0.000)	Δr	-0.024 (-0.321)	-0.041 (-0.706)
d	***1.000*** *(0.000)*	**-0.346** (-2.012)	Δd	**-0.669** (-7.352)	0.130 (1.808)
Trend	**-0.004** (-4.000)	0.000 (0.000)			
LR test for the over-identification restrictions$\Rightarrow\chi^2(2) = 1.55\ (0.51)$					

Notes: Bold values indicate statistical significance at the 5 % level. The figures in parentheses under the α- and β-coefficients are t-statistics.

The first cointegrating vector seems to be a long-run relationship for the trend-adjusted deficit/GNP ratio, explained by the short-term nominal interest rates, real output and broad money growth as in the following:

$$(\mathbf{d} - 0.004\ \mathbf{t}) = 0.590\ \mathbf{r} - 0.256\ \mathbf{y} - 0.630\ \mathbf{\Delta M}$$

This relationship confirms the fact that budget deficits in the Turkish economy are highly sensitive to changes in the short-term nominal interest rates. The relevance of such an inference is reflected by the significant long-run response coefficients of **r** and **ΔM** in the deficit equation. The positive coefficient on **r** reflects how higher interest rates increase the interest burden of domestic borrowing in the government budget through increased interest payments on debt. On the other hand, the negative coefficient on **ΔM** may be interpreted as the relieving impact of monetary expansion on budget deficits by decreasing the fiscal pressures on interest rates, which is consistent with the accommodative characteristic of the monetary policy in the Turkish economy in order to keep the costs of the debt rollover low. The negative long-run response of deficits to an increase in the real output (**y**) may indicate that high rates of real growth lighten the effects of factors that influence the deficit growth.

The estimated adjustment coefficients demonstrate that only money supply growth and real output reacts to short-run deviations from the error-correcting long-run path of budget deficits. Contrary to some views, neither the rate of inflation (CPI inflation) nor nominal interest rates (interest rates on deposits) are found to be adjusting to higher budget deficits in Turkey, reflected in the statistically insignificant adjustment coefficients in the short-run equations of $\Delta\mathbf{P}$ and **r**.

The second cointegrating vector is deemed to be a long-run relationship of broad money growth explained by the rate of inflation and budget deficits:

$$\Delta\mathbf{M} = 1.084\ \Delta\mathbf{P} + 0.346\ \mathbf{d}$$

This relationship implies that in the long-run money supply accommodates high inflation and budget deficits. The significant coefficient on $\Delta\mathbf{P}$ may be considered an indication of the demand determined characteristic of the nominal money supply growth in the Turkish economy. The significant coefficient on **d** may be interpreted as the credit-money expansion outcome of high public sector deficits, because the heavy reliance on commercial banks' resources in domestic debt finance results in parallel expansions in the assets and liabilities of commercial banks and hence, expansions in the interest-bearing money.

The quantity-theoretic approach to disinflation in the Turkish economy seems to be irrelevant considering the statistical findings of this analysis in which the weak-exogeneity status of the consumer price inflation is rejected while that of money stock is not. The weak-exogeneity of the inflation rate with respect to the parameters of interest may be considered an outcome of both the mark-up pricing behaviour of firms and the exchange rate pass-through prevalent in the Turkish economy. The oligopolistic structure of the private sector and the monopoly of the state enterprises in setting energy prices may constitute the basis for the weak-exogeneity of the inflation rate. Additionally, the high shares of imported capital and intermediate goods in the production strengthen the exchange rate pass-through argument in the determination of the inflation in Turkey.

The mark-up pricing argument can be valid also in explaining the weak-exogeneity evidence found for the nominal interest rates on bank deposits. Despite the impression that high nominal interest rates are the direct result of domestic borrowing in Turkey, for the deposit interest rates, the margin that exceeds the inflation rate seems to be a mark-up substantially determined by banks themselves depending on the risk structure of their portfolios and their

expectations. The indirect impact of deficits on interest rates may be explained by the risk premium set exogenously by the financial institutions, which is based on their expectations about the future course of the monetary and fiscal policies, expectations of failure in maintaining the debt rollover process and expectations of political instability. In this context, the estimated cointegration relationships show that the nominal interest rate influences budget deficits in the long-run whereas it adjusts to short-run deviations from the long-run paths of neither the budget deficits nor the money supply growth.

CONCLUSION

Empirical findings of this study verify the fiscal dominance over the monetary policy in the Turkish economy. Thus, the monetary transmission process in the economy is found to be extensively characterised by the accommodative behaviour of the monetary authority, which may be asserted by the rejection of the weak-exogeneity hypothesis of the monetary aggregate. Particularly, the monetary accommodation of budget deficits and inflation corresponds to the endogenous excessive expansion of the credit-money in the economy. The credit-money expansion occurs through the growth of commercial bank assets, which is led by debt sales to the banking system. Therefore, restrictive monetary policies may be ineffective and hazardous for the sustainability of budget deficits in Turkey because they increase the cost of domestic debt financing. This implies that high fiscal deficits, which are financed substantially by the banking system, reduce the likelihood of conducting any restrictive monetary policy in the Turkish economy. Thus, it may be argued that under such conditions the attempt to control monetary aggregates cannot be effective unless sound fiscal policies are conducted.

The findings of the study provide the evidence to the fact that the monetary authority can increase but not decrease the money supply exogenously in the long-run, consistent with the endogenous money hypothesis. Any attempt to reduce bank reserves may not slow down the money creation process because the supply of bank money is completely determined by the credit demand from both the private and public sectors. Note that the policy that encourages debt sales to the banking system in Turkey is nothing but the satisfaction of the public sector credit demand to finance budget deficits indirectly. Thus, the credit-money growth in Turkey is a by-product of efforts to keep the cost of domestic debt financing moderate and to maintain the stability of the financial system.

NOTES

[1] See Harris (1995: 39) for an overview of the "power and level of unit root tests".

[2] Juselius (1992) states "[S]ince a multivariate model is likely to give substantially lower residual variance than a univariate model, these tests are likely to be more efficient than the corresponding univariate tests."

[3] See Doornik and Hendry (2001) for the F-form of the likelihood ratio (LR) tests of system reduction.

[4] The computed SC(k) and HQ(k) statistics are as follows: SC(5)=-14.3, SC(4)=-15.5, SC(3)=-16.2, SC(2)=-16.8, SC(1)=-17.7, SC(0)=-16.3, HQ(5)=-17.4, HQ(4)=-18.1, HQ(3)=-18.3, HQ(2)=-18.4, HQ(1)=-18.7, HQ(0)=-6.8.

[5] The heteroscedasticity test $F_{HET}(180,297)=1.45$ and serial correlation test $F_{AR4}(100,160)=1.51$ are both significant at the 5 % significance level

[6] The trace statistic is argued to be more robust to skewness and kurtosis than the maximum eigenvalue statistic, as argued by Cheung and Lai (1993). Moreover, Doornik and Hendry (2001: 175) state that, "the sequence of trace tests leads to a consistent test procedure, but no such result is available for the maximum eigenvalue test". The critical values of the trace statistic are based on the approximations to the asymptotic distributions (see Doornik, 1998).

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