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Budget Deficits, Money Growth and Inflation: The Turkish Evidence

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Abstract: This paper investigates the long-run relationships between budget deficits, inflation and monetary growth in Turkey considering two alternative trivariate systems corresponding to the narrowest and the broadest monetary aggregates. While the joint endogeneity of money and inflation rejects the validity of the monetarist view, lack of a direct relationship between inflation and budget deficits makes the pure fiscal theory explanations illegitimate for the Turkish case. Consistent with the policy regime of financing domestic debt through commercial banking system, budget deficits lead to a growth not of currency seigniorage but of broad money in Turkey. This mode of deficit financing, leading to a creation of near money and restricting the scope for an effective monetary policy, may not be sustainable, as the government securities/broad money ratio cannot grow without limit.

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I. INTRODUCTION

The interrelationships between budget deficits, money growth and inflation have been at the heart of the monetary economics literature. It is often argued that monetization of budget deficits is the basic cause of inflation especially in developing countries. However, there appears to be no consensus, both theoretically and empirically, on the existence and the direction of the relationship between these three key macroeconomic variables.

Budget deficits are inflationary in the monetarist framework only to the extent that they are monetized. Sargent and Wallace (1981), on the other hand, argue that the monetarist arithmetic might be misleading as it ignores the fact that governments are constrained by their intertemporal budget. According to Sargent and Wallace (SW), tight money may lead to an unsustainable debt financing process and thus higher inflation in the long run. In this framework, inflation is a fiscal-driven monetary phenomenon, and nominal monetary growth is endogenously determined by the need to finance exogenously given deficit to satisfy the budget constraint. In the fiscal theory of the price level (FTPL), on the other hand, there is virtually no role for money in the determination of prices in a non-Ricardian world[1]. According to the FTPL, prices adjust to increases in nominal private sector wealth resulting from bond-financed deficits. In this non-Ricardian world, inflation is a symptom of too much nominal wealth chasing too few goods.

A trivariate system containing monetary growth (Δm), budget deficits (DEF) and inflation (Δp) can be postulated to test various alternative hypotheses. (i) An inflation equation with money and deficits being weakly exogenous can be interpreted as evidence supporting the monetarist approach. (ii) A money growth equation conditioned on weakly exogenous inflation supports the Keynesian endogenous money hypothesis. (iii) A DEF, Δm, Δp causal sequence, on the other hand, is consistent with the SW argument. (iv) The presence of deficits caused long-run inflation equation with money playing no role may be a strong support for the FTPL. The FTPL does not deny the possibility that fiscal disturbances may affect money supply. As noted by Woodford (2001), the monetary accommodation of inflation resulting from budget deficits is a story quite consistent with the FTPL[2]. That is, the causality is not from DEF to Δm and then from Δm to Δp as in SW, but instead from DEF to Δp and then from Δp to Δm in the FTPL.

Empirical studies investigating the links between budget deficits, money growth and inflation have yielded conflicting results. King and Plosser (1985) consider 12 countries and conclude that budget deficits do not contribute significantly to money growth or inflation. Similar results are reported by Joines (1985) for the USA, Karras (1994) for 32 countries and Sikken and Haan (1998) for 30 developing countries. On the contrary, Edwards and Tabellini (1991) find that inflation rates are determined by deficits in a sample of developing countries. The results of Favero and Spinelli (1999) support the hypothesis that the links between deficits, money, and inflation are not invariant to a policy regime change towards central bank independence in Italy. Metin (1998) considers Turkish annual data (1950-1987) and finds that budget deficits and debt monetization significantly affect inflation. The results by Özatay (2000) suggest that inflation adjusts to a monetary disequilibrium caused by budget deficits in Turkey.

In Turkey, fiscal deficits have often been interpreted as one of the major sources of severe inflation rates sustaining for more than two decades[3]. In comparison to OECD and many developing countries, the size of the public debt relative to GNP has been relatively modest. However, the basic problem has been its financing. After the financial liberalization program of 1980, the government's access to the Central Bank resources has been heavily restricted. The domestic debt financing has become the major source, especially after mid 1980s. The shift to the domestic debt financing from central bank monetization is often explained as the government avoiding inflation acceleration through the corresponding money supply growth. With the reserve accommodation, however, commercial banks have become the major source of the financing, and, the size of the banking system assets have increased correspondingly. The interrelationships between budget deficits, inflation and money growth under these conditions appear to be an important issue worthy of further investigation.

The plan of the rest of the paper is as follows. The following section investigates the long run relationships between the variables in two alternative systems containing money, inflation and budget deficits. Following the conventional literature, the narrowest monetary aggregate is considered in the first system in Section II.1. The fact that the bulk of government domestic debt instruments have been held by commercial banks leads us to consider also another system which contains the broadest monetary aggregate as the relevant money definition in Section II.2. Our concluding remarks are presented in Section III.

II. LONG-RUN RELATIONS IN THE SYSTEMS

We consider two alternative trivariate systems z_{1t} = [$DEFY_t$, Δp_t, Δcc_t] and z_{2t} = [$DEFY_t$, Δp_t, $\Delta m2y_t$] corresponding to the narrowest (currency in circulation, CC) and the broadest (M2Y) monetary aggregates. In the systems, DEFY is the ratio of consolidated budget deficit (deficit is a positive number) to GNP, p is the log of the wholesale price index.[4] All series are seasonally unadjusted. Throughout the paper lower-case letters signify natural logarithms of capital letters, and Δ denotes the first difference operator defined by $\Delta z_t = z_t - z_{t-1}$. The sample period is determined by the availability of quarterly data and is from 1983:1 to 1999:4.

The integration properties of the variables are investigated by conducting fourth-order augmented Dickey-Fuller (1981) (ADF(4)) tests. The results reported in Table 1 suggest that each of the variables in z_{it} is integrated of order 1 I(1).[5] This result is confirmed in the subsequent multivariate tests of stationarity.

[INSERT TABLE 1 ABOUT HERE]

For the 3x1 system z_{it}, we consider a VAR(k) process:

$$\Delta z_{it} = \Pi z_{it-1} + \Gamma_1 \Delta z_{it-1} + ... + \Gamma_{k-1} \Delta z_{it-k+1} + \mu + D_t + \varepsilon_t \qquad (1)$$

where μ is a vector of constants, D is a matrix of centered seasonal dummies and ε_t is a multivariate disturbance term. Given that z_{it} is an I(1) system, (1) represents a vector equilibrium correction mechanism (VECM) if the rank of Π, denoted by r, is such that $0 < r < 3$. For $0 < r < 3$, $\Pi = \alpha\beta'$, where α and β are 3 x r matrices of full column rank representing adjustment coefficients and cointegrating vectors, respectively. Under $0 < r < 3$, the VECM for long-run endogenous variable(s) can be written as:

$$\Delta z_{ijt} = \alpha\beta' z_{it-1} + \Gamma_1 \Delta z_{it-1} + \ldots + \Gamma_{k-1}\Delta z_{it-k+1} + \mu + \delta D_t + \varepsilon_t. \qquad (2)$$

A necessary condition for the weak exogeneity of a variable in the system is that no cointegration vector is significant ($\alpha = 0$) in the corresponding equation. Johansen (1988, 1995) and Johansen and Juselius (1990) provide a maximum likelihood procedure for the cointegration analysis of an I(1) system and it is this method we employ in this paper.

For each system we started with VAR(5) which is plausible for quarterly data and based our final choice of k on Likelihood Ratio (LR) tests of sequential system reduction. Furthermore, residual diagnostics of the unrestricted reduced form equations and system as a whole are also considered to evaluate the congruency of the maintained VAR system.

II.1. The Currency System

Table 3 reports the standard statistics and estimates for the Johansen procedure applied to the trivariate system $z_{1t} = [DEFY_t, \Delta p_t, \Delta cc_t]$. The VAR contains four lags, three centered seasonal dummies and an unrestricted constant[6]. Table 2 records various residual diagnostics to test the empirical adequacy of the system with $k = 4$. Each equation passes all of the diagnostics except the normality test. The residual non-normality may not be alarming as cointegration results appear robust to excess kurtosis (Gonzalo, 1994). These results suggest that the VAR(4) system can be interpreted as a valid approximation of the data generation process.

[INSERT TABLES 2 AND 3 ABOUT HERE]

The maximal eigenvalue (λ_{max}) and trace eigenvalue (λ_{trace}) statistics reject the null of no cointegration ($r = 0$), but not the null of at most one cointegrating vector ($r \leq 1$). As such, there appears to be a single cointegrating vector for the system. Table 3 also reports values of a multivariate statistic (Q_s) for testing the stationarity of the variables. This statistic tests the restriction that the coefficients of the designated variable and the rest of the variables in the maintained cointegration vector are unity and zero, respectively. The results of Q_s, under $r = 1$, strongly suggest the rejection of the stationarity null for each of the variables in z_{1t}.

The first vector β_1 normalised by Δp_t, can be interpreted as representing a long-run inflation equation:

$$\Delta p_t = 1.073\, \Delta cc_t + 0.002 DEFY_t$$

The coefficient estimates are consistent with expected sign priors. According to the LR statistics for testing long-run exclusion (Q_β), only DEFY is statistically insignificant in the cointegrating vector. Thus, the vector explains long-run currency growth-inflation relationship with budget deficits having no significant role. This preliminary result lends support to the argument that the fiscal theory of the price level (FTPL) is not supported by the Turkish data described by the currency variable space z_{1t}.

Table 3 reports t-statistics for standardized adjustment coefficients to test the weak exogeneity of the corresponding variable for the parameters of the long-run equation[7]. The test results strongly reject the quantity theory hypothesis, which states that money growth is exogenous for the long-run evolution of inflation. In contrast, the evidence reveals an economy in which money (currency) growth and inflation are jointly determined. As noted by Özmen (1998), currency seigniorage as a percentage of GNP has been relatively constant (around only 1%) in the face of fluctuating and accelerating inflation rates during the period. Thus, the evidence supports an argument that the seigniorage is the result of a passive monetary policy of nominal income accommodation, rather than the cause of inflation.

II.2. The Broad Money System

Table 4 reports the results of the cointegration anaysis for the system z_{2t} = [$DEFY_t$, Δp_t, $\Delta m2y_t$] with the VAR lag length[8] k = 4. The residual diagnostics recorded in Table 2 suggest the empirical validity of the VAR(4). According to the λ_{max} and λ_{trace} statistics there are two stationary relationships in the system. The results of Q_s, under r=2, strongly reject the stationarity null for each of the variables in z_{2t}.

[INSERT TABLE 4 ABOUT HERE]

The cointegration vectors are normalised to represent inflation and broad money growth equations, respectively:

$$\Delta p_t = 0.809\ \Delta m2y_t + 0.04DEFY_t$$

$$\Delta m2y_t = 0.12\ \Delta p_t + 0.434DEFY_t$$

According to the LR tests of exclusion (Q_β) reported in Table 4, all variables of the first vector except DEFY are significant. Thus the first vector can be interpreted to represent basically the long-run inflation-broad money growth relationship. Broad money and inflation appear to be endogenous for the parameters of the inflation equation as suggested by the results of the weak exogeneity tests. These results, rejecting the validity of both the quantity theory and the FTPL hypotheses for the Turkish data, are consistent with those obtained for the currency variable space.

The second vector indicates that broad money growth increases with budget deficits and inflation. All the variables are significant in the monetary growth equation, as implied by the significance of the exclusion tests (Q_β). Only DEFY is weakly exogenous, according to the tests for weak exogeneity. The adjustment coefficients indicate a rapid adjustment (less than two quarters) of inflation and money growth to disequilibrium.

The significant effect of budget deficits on money growth may not be an unexpected result as approximately 90% of government debt instruments have been held by the commercial banking system during the sample period. The numerical value of the DEFY coefficient is consistent with the fact that government securities has accounted for around 30% of the commercial banking system assets during most of the period. An active contractionary monetary policy may not be feasible in a high inflationary economy with substantial budget deficits financed via domestic commercial banking system. This is basically because an effective monetary control can increase the cost of domestic debt financing. Thus, a central bank concerned with the stability of a financial system and the cost of government debt financing may have no other operational choice than a reserve accommodation of credit money growth through domestic debt financing. This limits the possibility of financial crowding out and, hence, increases the expansionary impact of budget deficits on broad money supply.

III. CONCLUDING NOTES

In this paper we investigated the long-run relationships between budget deficits, inflation and monetary growth considering both the narrowest (CC) and the broadest (M2Y) monetary aggregates. The results of the Johansen cointegration analysis suggest that money and inflation are jointly determined. Consistent with the findings of Özmen (1998), the analysis provides no support for the quantity theoretical hypothesis that inflation is the result of an active monetary policy aiming to maximize seigniorage revenue. The joint endogeneity of currency growth and inflation lends support to the argument that the government has been attempting to keep the seigniorage revenue relatively constant by a passive monetary policy of nominal income accommodation.

The evidence does not support the fiscal theory of price level (FTPL), as there appears to be no direct link between budget deficits and inflation. Budget deficits exogenously determine money growth, which is consistent with the Sargent and Wallace (1981) argument. However, the growing monetary aggregate due to deficit financing is not outside money as postulated by SW but another aggregate, a bulk of which can be interpreted as inside or credit money (M2Y). This may not be an unexpected result, as throughout the sample period the government's access to the resources of Central Bank has been heavily restricted. There has been a shift to the domestic debt financing from central bank monetization. However, commercial banks have become the major source of financing, and with the reserve accommodation, the size of the banking system assets increased correspondingly. Consequently, budget deficits lead to a growth not of currency seigniorage but of interest-bearing broad money.

In Turkey, government domestic debt instruments are very liquid and their maturity is extremely short (see OECD 2001). Furthermore, commercial banks can hold government securities against their liquidity requirements and use them as collateral in the interbank money market. Consequently, government papers have become almost perfect substitutes for high-powered money. However, this mode of deficit financing leading to the creation of near money with very high real interest costs and heavily restricting the scope for an effective monetary policy may not be sustainable as the government securities-broad money ratio cannot grow without limit.

Table 1. ADF(4) Statistics for Testing for a Unit Root

	Levels		First differences
Series	λ_t	λ_m	λ_m
DEFY	-0.89	0.65	-4.08*
Δp	-3.37	-2.56	-4.85*
Δcc	-3.17	-2.50	-5.89*
Δm2y	-3.74*	-2.60	-5.24*
5 % CV	-3.57	-2.97	-2.97

Notes: All the test regressions contain three seasonal dummies and a constant term. The equations for λ_t include also a linear trend. An asterisk (*) indicates that the unit root null is rejected at the 5 % level, using MacKinnon (1991) critical values.

Table 2. Single equation residual diagnostics and system evaluation

System: z_1 = [DEFY, Δp, Δcc]

Statistic	ΔDEFY	$\Delta^2 p$	$\Delta^2 cc$	System
σ	0.03	0.05	0.07	
$F_{AR4}(4,42)$	1.40	0.26	0.52	
$F_{ARCH4}(4,38)$	0.22	0.05	0.30	
$\chi^2_{NORM}(2)$	2.97	48.9*	10.7*	
$F_{HET}(24, 21)$	0.48	0.56	0.44	
$F_{AR4}(36,95)$				0.83
$\chi^2_{NORM}(6)$				57.8*
$F_{HET}(144, 101)$				0.40

System: z_2 = [DEFY, Δp, Δm2y]

Statistic	ΔDEFY	$\Delta^2 p$	$\Delta^2 m2y$	System
σ	0.03	0.05	0.05	
$F_{AR4}(4,42)$	0.76	0.46	0.99	
$F_{ARCH4}(4,38)$	0.27	0.10	0.19	
$\chi^2_{NORM}(2)$	7.4*	33.1*	15.1*	
$F_{HET}(24, 21)$	0.48	0.56	0.44	
$F_{AR4}(36,95)$				1.34
$\chi^2_{NORM}(6)$				34.2*
$F_{HET}(144, 101)$				0.60

Notes: σ denotes equation residual standard deviations. The misspecification tests are against the alternative hypotheses: fourth-order autocorrelation (F_{AR4}), fourth-order ARCH (F_{ARCH4}), non-normality (χ^2_{NORM}) and heteroscedasticity (F_{HET}). See Doornik and Hendry (1997) for the tests. An asterisk (*) denotes rejection at the 5 % critical value.

Table 3. Cointegration analysis: Δcc, Δp, DEFY

Eigenvalues (λ)	0.343	0.190	0.0013
Hypotheses	r = 0	r ≤ 1	r ≤ 2
λ_{max}	26.42*	13.33	0.08
95% fractiles	21.0	14.1	3.8
λ_{trace}	39.83*	13.41	0.08
95% fractiles	29.7	15.4	3.8
Standardised eigenvectors β			
variable	β_1	β_2	β_3
Δp	1.000	0.338	-0.451
Δcc	-1.073	1.000	-0.048
DEFY	-0.002	-0.527	1.000
Standardised adjustment coefficients α			
equation	α_1	α_2	α_3
Δ^2p	-0.427	-0.565	0.009
	(-2.18)*	(-3.34)*	(0.091)
Δ^2cc	1.308	-0.562	-0.008
	(4.51)*	(-2.25)*	(-0.055)
ΔDEFY	-0.261	0.063	0.017
	(-1.94)*	(0.54)	(0.263)
LR test statistics for testing long-run exclusion: $Q_\beta \sim \chi^2(r)$			
r	Δp	Δcc	DEFY
1	12.13*	8.28*	0.02
Multivariate statistics for testing stationarity: $Q_s \sim \chi^2(p-r)$			
r	Δp	Δcc	DEFY
1	15.3*	17.5*	25.3*

Notes: The critical values for λ_{max} and λ_{trace} are from Table 2 of Osterwald-Lenum (1992). * indicates significance at the 5 % level.

Table 4. Cointegration analysis: Δm2y, Δp, DEFY

Eigenvalues (λ)	0.366	0.226	0.0013
Hypotheses	r = 0	r ≤ 1	r ≤ 2
λ_{max}	28.73*	16.13*	0.08
95% fractiles	21.0	14.1	3.8
λ_{trace}	44.95*	16.21*	0.08
95% fractiles	29.7	15.4	3.8
Standardised eigenvectors β			
variable	β_1	β_2	β_3
Δp	1.000	-0.120	-0.774
Δm2y	-0.809	1.000	0.515
DEFY	-0.040	-0.434	1.000
Standardised adjustment coefficients α			
equation	α_1	α_2	α_3
Δ^2p	-0.671	-0.753	0.011
	(-2.64)*	(-3.52)*	(0.102)
Δ^2m2y	0.768	-0.798	-0.002
	(3.01)*	(-3.70)*	(-0.022)
ΔDEFY	-0.048	-0.049	0.020
	(-0.28)	(-0.34)	(0.280)
LR test statistics for testing long-run exclusion: $Q_\beta \sim \chi^2(r)$			
r	Δp	Δm2y	DEFY
1	12.48*	6.31*	0.11
2	26.38*	21.80*	7.66*
Multivariate statistics for testing stationarity: $Q_s \sim \chi^2(p-r)$			
r	Δp	Δm2y	DEFY
1	16.3*	19.7*	26.4*
2	7.4*	7.3*	16.0*

Notes: The critical values for λ_{max} and λ_{trace} are from Table 2 of Osterwald-Lenum (1992). * indicates significance at the 5 % level.

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ENDNOTES

[1]See Woodford (1995, 2001) for the FTPL. For recent reviews of the FTPL, see Christiano and Fitzgerald (2000) and Buiter (2002).
[2]Woodford (2001, p. 674) explicitly states the causal sequence: "the government budget effects the general level of prices, and *only because prices change* does it also affect the money supply".
[3]Recent studies on Turkish fiscal policy include Özatay (1997, 2000), Özmen and Koğar (1998) and OECD (2001).
[4]All the data are from the Central Bank of the Republic of Turkey (CBRT) database and can be obtained by request from the authors at their internet addresses.
[5]The nonstationarity of DEFY is consistent with the weak sustainability of fiscal stance in Turkey (see, Özatay 1997 and Özmen and Koğar 1998).
[6]The sequential likelihood ratio (LR) test of system reduction from VAR(5) to VAR(4) and from VAR(4) to VAR(3) yielded F(9,99) = 1.11 (p = 0.36) and F(9, 107) = 1.81 (p = 0.07), respectively. As the reduction from VAR(4) to VAR(3) is not data-acceptable at the 10 % level, we preferred to commence from VAR (4).
[7]Note that these t-statistics has a meaning only if the corresponding cointegration vector defines a plausible economic relationship (see Johansen and Juselius 1990).
[8]The LR test of system reduction from VAR(5) to VAR(4) and from VAR(4) to VAR(3) yielded F(9,99) = 0.60 (p = 0.79) and F(9, 107) = 1.74 (p = 0.08), respectively. The system reduction from VAR(4) to VAR(3) is rejected at the 10% level, thus VAR(4) is chosen.