Budget Deficit, Money Supply and Inflation: Evidence from Low and High Frequency Data for Turkey

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1. INTRODUCTION

Many economists argue that inflation is strictly a monetary phenomenon and that inflation occurs when the rate of growth of the money supply is higher than the growth rate of the economy. This is the conventional monetarist linkage from the creation of base money to inflation when Central Banks issue money at a rate that exceeds the demand for cash balances at the existing price level and the increased demand in the goods market pushes up the price level as the public tries to get rid of its excess cash holdings. It is the contention of these economists that the Central Banks can eliminate the link between budget deficits and inflation by refusing to monetize the deficit, i.e., by not buying the bonds issued by governments.

Higher deficit policies may, however, lead to higher inflation even in the absence of monetization by Central banks. The governments borrowing requirement will increase the net credit demands in the economy, drive up the interest rates and crowd out private investment. The resulting reduction in the growth rate of the economy will lead to a decrease in the amount of goods available for a given level of cash balances and hence the increase in the price level. The other channel through which deficits can lead to higher inflation when Central Banks do not monetize the debt is the private monetization of deficits. This occurs when the high interest rates induce the financial sector to develop new interest bearing assets that are almost as liquid as money and are risk free. Thus, the government debt not monetized by the Central Bank is monetized by the private sector and the inflationary effects of higher deficit policies prevail.

The connection from deficits to inflation is nevertheless a difficult one to establish for a number of reasons. First, empirical studies trying to capture the link between budget deficits and inflation are bound to produce results that are quite sensitive to the choice of the model being used when one considers the number of possible versions that can be constructed intending to capture a given structure. [1] Secondly, it also needs to be noted that money-inflation relation is a highly dynamic one for the following reasons: (i) inflation will cause the velocity of circulation to rise and even an intact money supply will generate more inflation; (ii) the rise in inflation will reduce the available inflation tax base for the government and the attempt on the part of the government to collect a given inflation tax revenue will bring forth an increase in the tax (inflation) rate; (iii) inflation might cause budget deficits to rise (revenues to fall) due to the Tanzi effect, and the pursuing monetization could lead to even higher rates of inflation.

Turkish economic policies were overwhelmingly conducted according to the five year development plans until the early 1980s. Budget deficits were highly exogeneous, deficit monetization by the Central Bank was routine and monetary policy was subordinate to fiscal policy. A stabilization and liberalization package was introduced in January 1980 and gradual steps were taken towards capital account liberalization which resulted in the full convertibility of the Turkish Lira in August 1989. Domestic borrowing by the Treasury was initiated in 1986 which introduced an alternative source of financing for the government.

This study aims at gaining an insight into the channels through which the budget deficit has been operating in Turkey both in the post-World War II era and in the post-liberalization era where bond financing has been an additional source of deficit financing. The existence (and the nature) of a stable long run relationship between budget deficits, money and inflation, and the short-run dynamics of the inflationary process as of 1986 are the two major issues to be analyzed in the study.

[1] Paldam (1994) states that robustness of results is limited with the finding that the effect of the deficit is larger when the deficit is monetized. He presents this as an explanation for the more common presence of inflation in middle income countries with thin capital markets.

II. DATA AND METHODOLOGY

A. Methodology

In order to analyze short-run dynamics and long-run relationships among budget deficits, money and inflation, we make use of Vector Autoregression (VAR) and Vector Error Correction (VEC) specifications in this study. VARs are appropriate for cases where the aim is to examine the channels through which a number of variables interact as potentially spurious *a priori* constraints are not imposed in these models. As unrestricted VARs do not impose co-integration on its variables, a VEC model needs to be set up if the variables are known to be non-stationary and co-integrated.

In the VARs, two useful tools for short-run dynamics are the impulse response functions (IRFs) and the variance decompositions (VDCs). Estimates of the effects of changes in one variable on all the variables of the model are investigated by means of VDCs and IRFs. The VDCs indicate the percentage of the expected k-step ahead squared prediction of a variable produced by an innovation in another variable. The impulse response functions (IRFs) show the expected response of each variable in the system to a one standard deviation shock in one of the systems variables.

B. Data

We use annual data for the period 1948-1994 and quarterly data for 1987:01-1995:4 in the analysis. The annual variables are consolidated budget deficit over Gross National Product (GNP), *defognp*, percentage change of currency in circulation, *percur*, and GNP deflator based inflation, *inf.* The quarterly variables are consolidated budget deficit over Gross Domestic Product (GNP is not available quarterly), *defogdp*, percentage change in the Central Bank money, *percbmq*, and the Wholesale Price Index based inflation, *infwpi*. The quarterly data were obtained from the Central Bank of the Republic of Turkey and the annual data from the State Institute of Statistics and the Ministry of Finance.

III. EMPIRICAL RESULTS

Before we estimate the system that governs the relationship among *defognp*, *percur*, and *infwpi*, we check for the order of integration of these variables. Table 1 reports results of various unit root tests developed by Dickey and Fuller (1979) (ADF), by Phillips and Perron (1988) (PP), and by Kwiatkowski, Phillips, Schmidt and Shin (1992) (KPSS). The null hypothesis for the ADF and PP tests is that the series in question has a unit root whereas the KPSS test has the null hypothesis of level or trend stationarity.

Table 1

Unit Root Tests

Series	ADF[†]	PP[††]	KPSS[†††]	
defognp	2.611 [1, 2]	-4.328*	0.873*	0.211*
percur	-2.359 [1, c, t]	-4.522*	0.901*	0.182*
inf	-3.189 [c, t]	-2.447	0.887*	0.156*

Notes: * denotes significance at 5%.

[†] Augmented Dickey Fuller Unit root test. Numbers in square brackets denote the lags, constant (c) and trend (t) used in the augmentation of the test regression.

[††] Phillips-Perron Unit root test

[†††] Kwiatkowski, Phillips, Schmidt and Shin Unit root test for 4 lags. First column corresponds to the test of stationarity of the series in levels, second column gives the test of stationarity around a trend.

For *defognp* and *percur* ADF and PP test results are in conflict, and we base our inference of stationarity on the KPSS test and conclude that they are nonstationary. Regarding the order of integration, we conduct the same tests on the differences of the series and find them stationary.

We thus conclude that all variables are level and trend nonstationary.

Utilizing a likelihood ratio test, we select the lag length of two for the VEC mechanism that we estimate. On the basis of the plots of the series (Figure 1A), we assumed a model with a linear trend in the non-stationary part of the process. After we estimate the VEC following Johansens (1990) full-systems maximum likelihood estimation procedure, the trace test for the null hypothesis of no greater than *r* cointegrating vectors versus the alternative of more than *r* cointegrating vectors as well as the maximum eigenvalue test for the null hypothesis of *r* versus the alternative of *r+1* cointegrating vectors indicate with 95% confidence that there are two cointegrating vectors. The cointegrating vector coefficients and their weights in the error correction mechanism are in Table 2. The reported numbers are normalized with respect to inflation. Since there are two cointegrating vectors, the first two columns of the cointegration matrix are relevant.

Table 2

Cointegration results: $\hat{\alpha}$ and $\hat{\beta}$ matrices

f(*defogdp*, *percur*, *inf*) = 0

Variable	$\hat{\alpha}$ (weighting matrix)		
defognp	-0.026	0.006	0.001
percur	0.180	-0.253	0.013
inf	-0.788	-0.139	0.001
Variable	$\hat{\beta}$ (cointegration matrix)		
defognp	0.424	-46.947	38.265
percur	-1.139	2.134	0.098
inf	1.00	1.00	1.00

When we consider the two individual cointegrating relations, we see that in the first cointegration vector, *defognp* enters with an incorrect sign since an increase in the deficit GNP ratio seems to imply a lower inflation rate at the steady state, and in the second cointegrating equation *percur* enters with an incorrect sign since an increase in the growth rate of money seems to be associated with a lower long-run inflation rate. However, any linear combination of these two vectors is also a legitimate candidate for characterizing the long-run relationship among the three variables. We choose one such combination that is economically interesting through multiplying the first cointegration vector by 22.5468 and adding it to the second one and normalizing with respect to inflation. The following cointegrating relation is accordingly obtained:

$$inf = percur + 1.59\ defognp.$$

In this cointegration vector, long-run money neutrality is assumed. The relation reveals that a one unit increase in the deficit GNP ratio under money neutrality will increase the long-run inflation by

1.59 units. When we consider the weighting matrix a_i can be interpreted as the weight with which excess inflation enters equation *i*. The coefficients -0.79 and -0.14 indicate that excess inflation enters the error correction mechanism for inflation with a negative coefficient. This is expected since, if there is excess inflation, this excess is corrected towards the long-run equilibrium next period by an amount that is given by the weighting coefficient.

The assumed existence of a linear trend in the non stationary part of the process is tested for by means of a likelihood ratio test. The test statistic of 22.53 which is distributed as chi-squared with one degree of freedom leads to the rejection of the null hypothesis that linear trend is absent in the nonstationary process at $\alpha = 0.05$, validating the maintained assumption. We then check whether any of the variables can be excluded from the cointegration vectors. The likelihood ratio statistics of 20.52, 28.21, and 29.26 which are distributed as chi-squared with two degrees of freedom rejects the exclusion of *defognp, percur*, and *inf* from the cointegration vectors, respectively, at $\alpha = 0.05$. We finally test for weak exogeneity of the variables in the cointegration vectors. The likelihood ratio test statistics of 18.72 and 32.83 lead to the rejection of the null hypothesis that *percur* and *inf* are weakly exogenous at 5% level of significance. For *defognp*, the rejection of weak exogeneity depends whether one uses 95% or 97.5% level of confidence, thus one may treat deficit over GNP ratio as weakly exogenous to the system.

Since the sample period for the above analysis is 1948-1994 while bond financing has become available as of 1986, we rerun the VEC mechanism estimation for the period 1948-1985 to investigate if deficits had a greater impact on inflation during the pre-bond financing period. The cointegrating relation under the money neutrality assumption using the two cointegrating vectors obtained from this particular sample period is as follows:

$$inf = percur + 5.67\ defognp.$$

In comparison to the previous relation, we see that a one unit increase in the deficit GNP ratio under money neutrality will increase the long-run inflation by 5.67 which is much higher than 1.59 for the whole sample indicating greater impact of deficit on inflation during pre-bond financing period.

analyzed the post-World War II data in two sample periods: 1948 to 1986 and 1948 to 1994. Unrestricted VAR estimates are given in Table ?. Figures 2A and 2B report the IRFs of the estimated VAR whereas Figures 3A and 3B report the percentage of the k-step- ahead prediction error in *defognp*, *percur,* and *inf* due to innovations in the same variables for selected values of k .for sample periods 1948-1994 and 1948-1986, respectively. For both sample periods, *defognp* seems to be exogeneous; forecast error variance of *defognp* is overwhelmingly due to innovations in *defognp* itself. Squared prediction error of *infwpi* is explained less by *defognp* and *percur* and more by *inf* itself when the full sample period is used as opposed to the period up to 1986. This might be interpreted as being suggestive of increased inertia in the inflationary process which was one source of motivation for analyzing the post -1986 period using higher frequency data.

Analyzing the plots of the quarterly series given in Figure 4, we observed the following: (i) unlike the annual data, there is no linear trend in any of the series; (ii) there exists non-stochastic seasonality in all of the series; (iii) there is a spike in the first quarter of 1994 in the inflation series due to the 1994 financial crisis.

We start out by checking the level of integration of the variables. Using the aforementioned stationarity tests we find the *defogdp*, *percbmq*, and *infwpiq* series to be level and trend stationary.

Table 3

Unit Root Tests

Series	ADF†	PP††	KPSS†††	
defogdp	-6.170* [1, c, t]	-5.364*	0.329	0.107
percbmq	-3.798* [1, ... , 5, c]	-4.575*	0.218	0.147
infwpiq	-5.443* [c]	-5.470*	0.282	0.110

Notes: * denotes significance at 5%.

† Augmented Dickey Fuller Unit root test. Numbers in square brackets denote the lags, constant (c) and trend (t) used in the augmentation of the test regression.

†† Phillips-Perron Unit root test

††† Kwiatkowski, Phillips, Schmidt and Shin Unit root test for 9 lags. First column corresponds to the test of stationarity of the series in levels, second column gives the test of stationarity around a trend.

The likelihood ratio test was utilized to select the appropriate order of the VAR estimation and a lag order of four is chosen. VAR estimation results are presented in Table 3 and IRFs and VDCs are given in Figures 5 and 6, respectively. IRFs indicate decaying (not persistent) responses as expected since the series are stationary. A one standard deviation shock to *defogdp* causes an increase in *defogdp* overall, but shocks to *percbmq* and *infwpiq* do not seem to affect *defogdp* which is consistent with *defogdp* being weakly exogeneous. The effect of shocks to *defogdp* on *percbmq* and *infwpiq* are insignificant. This may be due to the introduction of a new source of deficit financing, namely domestic borrowing, in the post-1986 period. VDCs in Figure 6 indicate that the forecast error variances of all variables for all given values of k are accounted for mostly by the variables themselves, indicating an evolution of each series to a univariate process. Since this is suggestive of increased inertia in the inflationary process, we finally applied recursive Box-Jenkins modelling to the inflation series and estimated it as an AR(1) process. The autoregression coefficient was observed to approach 1.0 through time which lent support to Paldams (1994) assertion that inflation in high inflation countries is a process with much inertia.[2]

[2] Paldam (1994) states that for the countries in his sample, the autocorrelation is approximately 0.60. See Figure 1B for the graphical representation of the recursive AR(1) estimation for the Turkish case.

IV. CONCLUSION

Using annual Turkish data covering the post World War II period, the existence of a stable long-run relationship between budget deficits, money growth and inflation is tested in this study and the results have been affirmative. Using the cointegrating vectors found in the study, a significant impact of budget deficits on inflation can not be refuted under the assumption long-run monetary neutrality.

However, when we utilized an unrestricted VAR model using quarterly data corresponding to the post-bond financing period, the results are suggestive of a weakened link from the other variables to inflation. A further check using an ARIMA approach validated the same result for the post- World War II period and it is shown that the inertia in the inflation process was increasing over time. The availability of bond financing after 1986 might account for the weakening in the link from the budget deficits to inflation to a certain extent. The increased inertia in the inflation could be due to the accumulation of inflationary expectations. It can also be explained by the general equilibrium nature of the price system where the increase in one price will drive up the other ones and the continuity of the process will be guaranteed whenever overshooting of some prices occurs.

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