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The Relationship Between Inflation and the Budget Deficit in Turkey

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This article analyzes the empirical relationship between inflation and the budget deficit for the Turkish economy by a multivariate cointegration analysis. A single-equation model shows that the scaled budget deficit (as well as income growth and debt monetization) significantly affects inflation in Turkey. The conditional model of inflation is constant, and it encompasses a previously estimated model.

KEY WORDS: Cointegration; Encompassing; Exogeneity; Turkish inflation.

An extensive literature has examined the relationship between the budget deficit and inflation. At a theoretical level, Sargent and Wallace (1981) showed that under certain conditions, if the time paths of government spending and taxes are exogenous, bond-financed deficits are nonsustainable, and the central bank should eventually monetize the deficit. This will increase the money supply and inflation in the long run. These findings have subsequently been generalized for the open economy case and for alternative forms of financing (see Scarth 1987; Langdana 1990).

The empirical relationship between the deficit and inflation in developed countries has been studied in detail (see Hamburger and Zwick 1981; Dwyer 1982; Hein 1983; Ahking and Miller 1985; King and Plosser 1985; Protopapadakis and Siegel 1987; Burdekin and Wohar 1990; Ho 1990). Empirical studies of developing countries include those of Dornbush and Fisher (1981), Bhalla (1981), Siddiqui (1989), Choudhary and Parai (1991), Buiter and Patel (1992), Dogas (1992), Sowa (1994), Hondroyiannis and Papapetrou (1994), and Metin (1995). These studies did not yield conclusive results on the relationship between the budget deficit and inflation, either in the short run or in the long run. Specifically, Hamburger and Zwick (1981) found that growth in Federal Reserve debt holdings exerted a significant inflationary impact on the U.S. economy over 1961–1982, yet a growth in nonmonetized debt had a negative short-run effect on inflation. Ahking and Miller (1985) modeled deficits, money growth, and inflation over 1950–1980 as a trivariate autoregressive process. They found government deficits to be inflationary in the 1950s and 1970s but not in the 1960s. Using a rational-expectations macro model of Peruvian inflation, Choudhary and Parai (1991) found that budget deficits, as well as the growth rate of money supply, have significant impacts on inflation. Similarly, Dogas (1992) found that the public deficit affects inflation in Greece. Hondroyiannis and Papapetrou (1994) also found a relationship between the Greek government budget and price level. Using an error-correction model, Sowa (1994) found that inflation in Ghana is influenced more by output volatility than by monetary factors, both in the long run and in the short run.

For Turkey, Metin (1995) analyzed inflation using a general framework of sectoral relationships and found that fiscal expansion was a determining factor for inflation. The excess demand for money affected inflation positively, but only in the short run. On the other hand, imported inflation, the excess demand for goods, and the excess demand for assets in the capital markets had little or no effect on inflation. A key policy implication of Metin (1995) is that Turkish inflation could be reduced rapidly by eliminating the budget deficit.

The aforementioned general literature influences the current study, which builds directly on Metin (1995). The large public-sector budget deficits and the relatively high inflation in Turkey during the last four decades have sparked debate on their consequences for the Turkish economy. The main question is whether bond-financed deficits are inflationary or whether only monetized deficits are inflationary. To answer this question, this article investigates the relationship between Turkish inflation and budget deficits over 1950–1987. Although the government shifted from monetizing the deficit to bond financing in the mid-1980s, the short annual sample on Treasury bonds precluded sorting out the effects of this alternative means of deficit financing. Therefore, I have used Metin's (1995) dataset for analyzing the relationship between inflation and the public-sector budget deficit, considering a closed-economy public-finance approach. The closed-economy assumption may appear restrictive, but Metin (1995) showed the lack of external effects in the determination of Turkish inflation. The empirical analysis herein is of general interest because many other developing countries have experienced budget and inflation difficulties similar to those in Turkey.

Section 1 presents a historical background to the Turkish economy for 1950–1987, and Section 2 develops a theoretical framework based on the public-finance approach. Section 3 tests for budget deficits and inflation being cointegrated (and finds that they are). Although weak exogeneity does not appear valid, a parsimonious conditional model is still developed (Sec. 4). This model is empirically constant, whereas the corresponding marginal model is not, thus showing super exogeneity for dynamics parameters.

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Additionally, the new conditional model encompasses the model of Metin (1995).

1. HISTORICAL BACKGROUND

This section presents a brief economic history of Turkey, focusing on inflation and budget financing.

From the 1950s until 1980, the Turkish government consistently followed a policy of import substitution, with prohibitions on imports of commodities. State economic enterprises (SEE's) were established to produce agricultural commodities, several manufactured goods, and minerals. In the late 1950s, the Turkish economy experienced severe balance-of-payment difficulties and rising inflation. Efforts to control inflation consisted largely of price controls. Private-sector firms responded either by shutting down or by selling on the black market. SEE's, however, sold at official prices and experienced losses. As inflation increased, these losses reached enormous amounts. The losses were automatically financed by the credits extended by the Central Bank to the SEE's, resulting in high money growth (see Aktan 1964; Okyar 1965; Fry 1972, 1980; Krueger 1974, 1995; Onis and Riedel 1993).

In 1958, Turkey implemented a fairly typical International Monetary Fund (IMF)-supported stabilization program, which improved the foreign-exchange situation and drastically reduced inflation. The most important component of the program was an increase in the prices of SEE goods, a component that was featured prominently in the 1970 and 1980 reforms as well. Raising those prices in 1958 resulted in an immediate and once-and-for-all increase in the price level, after which the reduced rate of expansion of Central Bank credits reduced inflation. Although inflation dropped from 25% in 1958 to less than 5% in 1959, real gross domestic product (which had been declining) started growing immediately due to the greater availability of imports.

Turkey was among the more rapidly growing developing countries during most of the 1960s, with an annual inflation rate of 5%–10%. The nominal exchange rate was kept constant after the 1958 devaluation. Investment spending increased and was financed mainly by foreign aid. In the late 1960s, foreign aid did not increase, but the rate of investment spending was maintained. In addition, some difficulties appeared in obtaining imports, creating visible restraints on economic activity and growth.

Although inflation was rising at the time, the main reason for the 1970 devaluation was foreign-exchange difficulties. After the devaluation, export earnings increased sharply, and Turkish workers in Germany and other western European countries started remitting a significant amount of foreign exchange. Because there was no mechanism readily at hand for the Central Bank to sterilize these inflows, the money supply expanded rapidly and inflation increased, reaching an annual rate of 25% by 1973. In the early and the mid-1970s, the problem of the growing public-sector deficit also arose from the expenditure side. In particular, large salary increases were granted to civil servants, and substantial increases in transfer payments were made to SEE's, which had financial deficits due to both increased wage costs and a rise in the rate of investment by the SEE's (see Onis and Riedel 1993). The growth of government spending during a boom in the mid-1970s led to rising budget deficits, for which the Central Bank provided a major part of the financing. The public sector borrowing requirement (PSBR) was 4.3% of gross national product (GNP) in 1973, more than doubling to 10.7% in 1979.

Inflation reached about 100% in 1980, apparently fed by monetization of the public-sector deficit. Policy changes in the early 1980s were designed to shift Turkey's growth strategy away from import substitution and toward greater integration with the international market. The 1980 stabilization program attempted to deal with inflation by creating greater efficiency in operating the SEE's, restraining the growth of public expenditure, reducing subsidies, and attempting to improve revenue collection. Under the government's liberalization program, the financial performance of SEE's improved substantially. Unlike their performance during the previous decades, SEE's appeared to have contributed positively to the financial position of the central government in the 1980s. The government's restrictive stance could not be fully maintained, however. The PSBR remained at about 6% of GNP during the first half of the 1980s and rose to 8.3% in 1987, the highest since 1980. Contributing factors included slow growth of revenues, a strong increase in budget transfers to loss-making SEE's, higher than planned wage and salary raises in the public sector, and an election. After 1980, policy reforms continued. Although inflation fell to approximately 35% in 1982, it started rising again and continued to be a problem throughout the 1980s.

2. THE ECONOMIC FRAMEWORK

This section summarizes the theoretical model underlying the empirical analysis. In a closed economy, it is assumed that all public debt takes the form of noninterest-bearing money. The public sector budget identity is then

$$G - T = \Delta H \tag{1}$$

or

$$\frac{G-T}{PY} = \frac{\Delta H}{PY}, \tag{2}$$

where G is public-sector expenditures, T is public-sector revenues, Y is real income, P is the price level, and H is base money. In a steady-state growing economy, it follows that

$$\begin{aligned} \Delta(H^*) &= (H^*)\left(\frac{\Delta H}{H} - \frac{\Delta P}{P} - \frac{\Delta Y}{Y}\right) \\ &= \frac{\Delta H}{PY} - H^*(\Delta p + \Delta y), \end{aligned} \tag{3}$$

where Δ is the difference operator; $H^*, \Delta p$, and Δy are scaled base money (H/PY), inflation, and the growth rate of real income, respectively; and variables in lower case are in logarithms. It is assumed that the long-run income elasticity of the demand for money is unity. Then the simplified

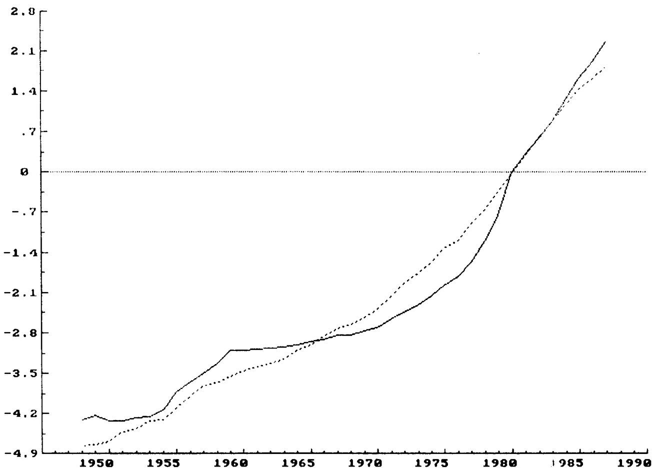


Figure 1. Consumer Price Index and Base Money: p = ——, h = ---.

budget constraint is

$$\Delta(H^*) = \frac{G-T}{PY} - H^*(\Delta p + \Delta y). \tag{4}$$

Solving (4) for Δp, I obtain the following relation:

$$\Delta p = c + \psi_1 B - \psi_2 \Delta y, \tag{5}$$

where B is the scaled budget deficit $(G-T)/H, c$ is the constant term (interpretable as the inertial inflation rate), and ψ_1 and ψ_2 are slope coefficients associated with the scaled deficit and income growth. Here, ψ_1 and ψ_2 are equal coefficients with an opposite sign [see Phelps (1973), Anand and van Wijnbergen (1989), and Rodrik (1990) for theory and empirical analysis]. The remainder of this article empirically analyzes the relationship between the budget deficit, inflation, base money, and real income growth.

3. THE DATA, UNIT-ROOT TESTS, AND COINTEGRATION ANALYSIS

This section tests for unit roots in the series of interest (Sec. 3.1) and for cointegration between the series (Sec. 3.2).

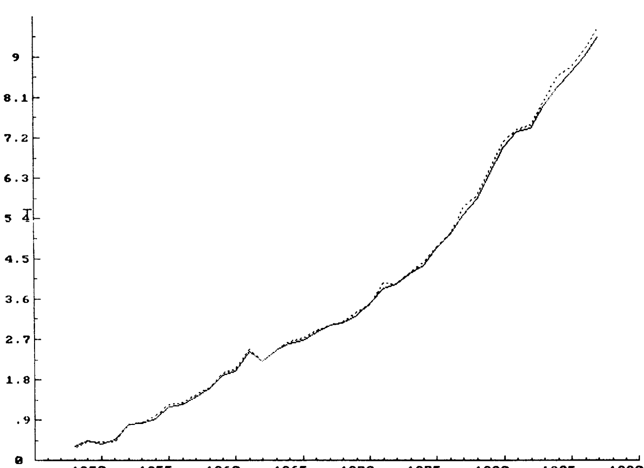


Figure 2. Revenues and Taxes: g = ——, t = ---.

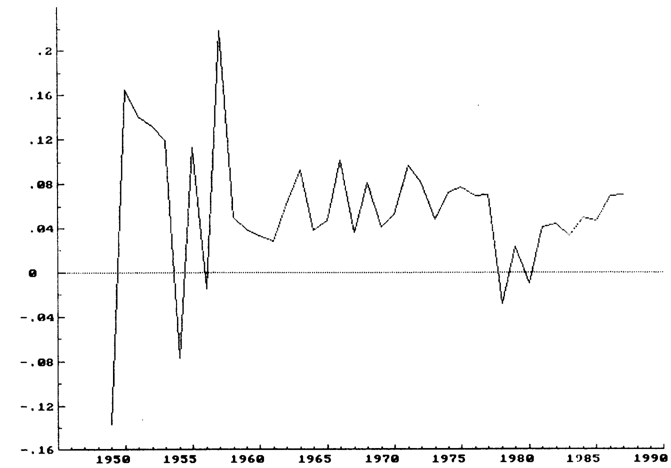


Figure 3. The Growth Rate of Real Income: Δy = ——.

3.1 The Data and Unit-Root Tests

The data used are annual over 1950–1987. Budget expenditures (G) and budget revenues (T) are from the budget and final accounts, respectively [Turkish lira (TL) Billion]. The general budget deficit $(G-T)$ is the primary deficit, which excludes interest payments (TL Billion). The budget deficit does not include the SEE's deficit. Because reliable statistics about SEE's deficits are available only after the second half of the 1970s, the general budget deficit is therefore used as a proxy for the total deficit. The price level (P) is the consumer price index with base year 1980, Y is real GNP (TL 1980 Billion), and H is base money. The components of base money are currency in circulation, vault cash, legal reserves, and Central Bank sight deposits (TL Billion). The Appendix describes the data in greater detail.

Figures 1–4 show $(h, p), (t, g), \Delta y$, and $(\Delta p, B)$, respectively. Visually, all series appear at least I(1); the augmented Dickey–Fuller (1981) (ADF) test statistics in Table 1 support the graphical explanation. p and h appear I(2) (Fig. 1), and h^* is I(1). Government expenditures (g) and revenues (t) also seem to be I(2) (Fig. 2), but the scaled deficit B is clearly I(1). Δy is I(0) and, from its plot, looks like a

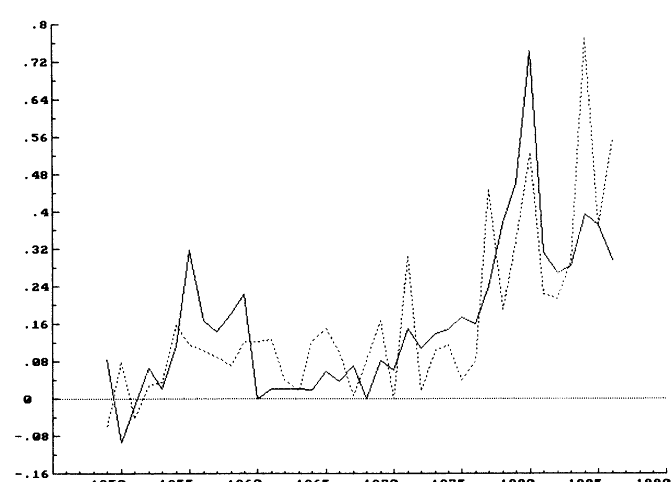


Figure 4. Inflation and the Rescaled Budget Deficit: Δp = ——, B = ---.

Table 1. Augmented Dickey–Fuller Test Statistics

Null order	Variable g	t	B	p	h	y	h^*
I(1)	−1.15	−.89	−1.03	−.28	−.44	−1.39	−1.11
	(10)	(1)	(2)	(1)	(1)	(5)	(3)
I(2)	−2.14	−2.20	−8.07**	−2.78	−3.68	−8.51**	−5.04**
	(3)	(3)	(1)	(0)	(0)	(0)	(2)
I(3)	−5.95**	−6.68**		−7.26**	−9.02**		
	(2)	(2)		(0)	(0)		

NOTE: For a given variable and null order, two values are reported. The first row is the t value, which is the ADF statistic, and the second row is the longest significant lag with significant t value. Five lags are allowed in each variable's ADF regression, but twelve lags are allowed for g and t. All regressions include a constant term and a trend. The sample is 1954–1987 (T = 34) if the variables are in their log levels (except B), 1955–1987 (T = 33) if they are in first differences, and 1956–1987 (T = 32) if variables are in second differences. The critical values are from MacKinnon (1991, table 1). Here and elsewhere in this article, ** and * denote rejection at the 1% and 5% critical values.

stationary heteroscedastic series (Fig. 3). Figure 4 captures the essence of the cointegration analysis: Both Δp and the scaled budget deficit B share the same upward trend over time.

3.2 System Cointegration Analysis

This subsection tests for cointegration among the series $(\Delta p, h^*, B, \Delta y)$. I test for cointegration in a first-order vector autoregression (VAR), using the multivariate cointegration procedure of Johansen (1988) and Johansen and Juselius (1990). The VAR includes a constant term, a trend, and an impulse dummy ($i1980$). The impulse dummy represents the structural change in the Turkish economy that took place in 1980. The constant and $i1980$ enter the system unrestrictedly. The trend is restricted to lie in the cointegration space because a quadratic deterministic trend in levels of economic variables is not usually a sensible long-run outcome (see Doornik and Hendry 1994). The cointegration results are quite sensitive to the lag length of the VAR. Our choice of one lag is based on the Schwarz and Hannan–Quinn criteria, both of which pointed to a single lag. The estimation period is 1952–1987.

Table 2 summarizes the cointegration results. It includes the eigenvalues, the max and trace statistics, the standardized estimated feedback coefficients α and cointegrating vector β', and statistics for testing restrictions on α. The cointegration test statistics are corrected for sample size (see Reimers 1992), and they suggest three cointegrating vectors. The residual misspecification tests appear satisfactory. None of the equations exhibits autocorrelation, and the equations for B and Δy have nonnormal residuals.

Because I find three stationary relations, I need to identify the estimated cointegrating vectors before I interpret them. Assuming that Δy is trend stationary, the second row of the β is an inflation relation, and the third cointegrating vector is including just Δp and B, I test the identification of all cointegrating vectors. The expected β' matrix will be

$$\beta' = \begin{bmatrix} 1 & 0 & 0 & 0 & * \\ 0 & 1 & * & * & 0 \\ 0 & * & 1 & 0 & * \end{bmatrix},$$

and implementing those identification restrictions leads to the restricted form β' and α matrix reported in Table 3. The likelihood ratio test statistic suggests that all three cointegrating vectors are identified $\chi^2(2) = 1.1559\,[.5611]$ (see Johansen 1991, theorem 5.1).

From the standardized β' eigenvectors, the first cointegrating vector is the growth rate of real income. The second one is an inflation relation:

$$\Delta p = .58B + .35h^*. \tag{6}$$

The public sector deficit B enters with a positive coefficient (.58), and scaled base money h^* also has a positive coef-

Table 2. A Cointegration Analysis of $\{\Delta y, \Delta p, B, h^*\}$

Eigenvalues	.739	.662	.445	.085
Hypotheses	$r = 0$	$r \le 1$	$r \le 2$	$r \le 3$
Max statistic	40.4	32.5	17.7	2.7
95% critical value	27.1	21.0	14.1	3.8
Trace statistic	93.2	52.9	20.4	2.7
95% critical value	47.2	29.7	15.4	3.8

Standardized eigenvectors β'

Variable	Δy	Δp	B	h^*	Trend
	1	.188	−.124	−.088	.0006
	−1.222	1	−2.515	−.769	.0128
	−1.042	1.745	1	.009	−.0191
	.125	.611	−.443	1	.0052

Standardized adjustment coefficients α

Δy	−1.200	.097	.054	.006
Δp	−.692	−.129	−.201	.042
B	.079	.337	−.185	.071
h^*	−.016	.076	−.037	−.115

Weak exogeneity test statistics

Variable	Δy	Δp	B	h^*
$\chi^2(5)$	2.41	12.963	12.506	38.848
p value	[.4911]	[.0047]**	[.0058]**	[.000]**

Diagnostic statistics

Variable	Δy	Δp	B	h^*
Normality $\chi^2(2)$	11.35**	.61	7.93*	.24
ARCH 1 $F(1, 25)$	1.14	.58	.25	1.61
AR 1-2 $F(2, 25)$	.86	1.29	1.27	1.44

NOTE: r is the hypothesized number of cointegrating vectors. The critical values for the cointegration tests are from Osterwald-Lenum (1992). The Jarque-Bera (1980) normality test statistic has a χ^2 distribution with 2 df under the null of normal errors. ARCH F(df1, df2) refers to the test for ARCH errors, introduced by Engle (1982). The AR1 F(df1, df2) is the test for residual autocorrelation.

Table 3. A Restricted-Form Cointegration Analysis

	Standardized eigenvectors β'				
Variable	Δy	Δp	B	h^*	Trend
	1.000	0.000	0.000	0.000	0.000
	0.000	1.000	−0.585	−0.349	0.000
	0.000	1.148	1.000	0.000	−0.012
	Standardized adjustment coefficients α				
Δy	−1.348	−.024	−.038		
Δp	−.348	−.487	−.085		
B	−.157	.747	−.611		
h^*	−.067	.162	−.134		
	Weak exogeneity test statistics				
Variable	Δy	Δp	B		
$\chi^2(6)$	74.151	16.136	23.989		
p value	[.000]	[.000]	[.000]		

ficient (.35). The third stationary relationship is between inflation and the scaled budget deficit.

The standardized α coefficients show that the main effect of the first cointegrating vector is on Δy. From the second column of α, feedback of the second cointegrating vector on both B and Δp is .75 and $-.49$, respectively. The third cointegrating vector primarily affects the scaled deficit B. Weak exogeneity for β can be tested using the Johansen (1992a,b) procedure. The results suggest that $\Delta p, B$, and h^* cannot be assumed weakly exogenous for β, but Δy can be (see Table 2). Weak exogeneity of the variables is also tested jointly with identification restriction and rejected for $\Delta p, B$, and Δy (see Table 3).

For inference, conditional models should have regressors that are weakly exogenous; see Engle, Hendry, and Richard (1983). In the context of cointegration, weak exogeneity means that inference about the cointegrating vector can be performed on the conditional model without loss of information relative to a system analysis. Even lacking weak exogeneity, single-equation modeling can proceed, treating the system-based estimated cointegration coefficients as given; see Juselius (1992). Section 4 develops such a conditional model and examines its properties.

4. SINGLE-EQUATION MODELING

This section develops a parsimonious, conditional, single-equation model for inflation, in which inflation depends on the scaled budget deficit, the real growth rate of income, and scaled base money. Section 4.1 develops a parsimonious conditional model from a general autoregressive distributed lag and shows the constancy of this conditional model. Section 4.2 estimates some marginal equations and tests their constancy. Finally, Section 4.3 compares the model estimated by Metin (1995) with the conditional model developed in this article, using the standard encompassing framework.

4.1 Single-Equation Analysis and the Constancy of a Conditional Model

Because weak exogeneity does not appear valid (except for Δy), Juselius's (1992) approach is used for single-equation modeling. Recalling the cointegration analysis in the previous Section 3.2, a single inflation equation is constructed. The inflation model includes the error-correction terms (ECM's) obtained from the earlier cointegration analysis. The first ECM (CI2) is constructed using Equation (6), and the second ECM (CI3) is obtained from the third row of the β' matrix given in Table 3. Then the general ECM model involves $\Delta^2 p, \Delta B, \Delta y$ (because it is stationary), Δh^*, their lags, and the lagged ECM's. Here, single-equation modeling starts with an unrestricted fourth-order autoregressive distributed lag (ADL) in the (log) levels of the variables, written as an error-correction model:

$$\begin{aligned}\Delta^2 p_t &= \sum_{i=0}^{k-2} \beta_{1i} \Delta B_{t-i} + \sum_{i=0}^{k-2} \beta_{2i} \Delta y_{t-i} + \sum_{i=0}^{k-2} \beta_{3i} \Delta h^*_{t-i} \\ &\quad + \sum_{i=0}^{k-2} \beta_{4i} \Delta^2 p_{t-i} + \beta_5 \mathrm{CI2}_{t-1} \\ &\quad + \beta_6 \mathrm{CI3}_{t-1} + c + u_t, \end{aligned} \tag{7}$$

where $k = 4$ and c represents the constant term, trend, and impulse dummies $i1980$ and $d55$. The model suffered from a major outlier in 1955 that was not explained by the variables in the information set and did not correspond to any previous historical events. Thus, I created a dummy ($d55$) to pick this up. This equation is a reparameterization of the ADL model and is in I(0) space. Furthermore, this equation obviates the need for weak exogeneity with respect to the cointegrating estimates from the Johansen-system procedure.

Equation (7) is fitted over 1954–1986. Estimation results and diagnostic statistics are reported in Table 4, column 2. The diagnostic statistics test against several alternative hypotheses—residual autocorrelation (DW and AR), skewness and excess kurtosis (normality), autoregressive conditional heteroscedasticity (ARCH), and heteroscedasticity (RESET). The estimated ECM model embodies the sensible long-run solution in (6) and has good diagnostic statistics. The RESET test suggested a possible nonlinearity in the model, however, perhaps because many of the disequilibria are likely to interact.

The general ECM can be simplified. Modeling general to specific, a parsimonious model of inflation is obtained (Table 4, col. 3):

$$\begin{aligned}\Delta^2 p_t &= \underset{[.1912]}{+\,.2487} + \underset{[.00142]}{.002153\,\mathrm{trend}} + \underset{[.0479]}{.3762 i1980} \\ &\quad \underset{[.0583]}{+\,.3357 d55} - \underset{[.1864]}{.3729 \Delta^2 p_{t-1}} + \underset{[.1451]}{.2031 \Delta B_t} \\ &\quad \underset{[.3809]}{-\,.704 \Delta y_t} + \underset{[.3128]}{.5179 \Delta y_{t-2}} + \underset{[.2884]}{.5045 \Delta h^*_{t-2}} \\ &\quad \underset{[.1110]}{-\,.1772\,\mathrm{CI2}_{t-1}} - \underset{[.0561],}{.1062\,\mathrm{CI3}_{t-1}} \end{aligned} \tag{8}$$

where $R^2 = .89, \hat{\sigma} = .0476$, DW $= 1.60$, AR$(2, 20) = 1.77$, ARCH: $F(1, 20) = .13$, Normality: $\chi^2(2) = 1.85$, and RESET: $F(1, 21) = 4.57$.

Table 4. The Conditional and Marginal Models

	Dependent variable				
	$\Delta^2 p$	$\Delta^2 p$	ΔB	Δy	Δh^*
	Estimation method				
	OLS	OLS	RLS	RLS	RLS
Sample	1954–1986	1954–1986	1954–1986	1954–1987	1954–1986
Constant	.097(0.659)	.249(0.191)	.038(0.026)	.057(0.013)	
Trend	.0019(0.0019)	.0021(0.0014)			
i1980	.340(0.159)	.376(0.048)	.351(73.44)	−.067(157.94)	−.324(170.72)
d55	.350(0.078)	.336(0.058)			
$\Delta^2 p_{t-1}$	−.359(0.282)	−.373(0.186)			
$\Delta^2 p_{t-2}$	.0532(0.201)				
ΔB_t	.165(0.278)	.203(0.145)			
ΔB_{t-1}	.140(0.591)		−1.04(0.185)		
ΔB_{t-2}	.079(0.355)		−.862(0.326)		
ΔB_{t-3}			−1.100(0.603)		
ΔB_{t-4}			−.680(0.316)		
ΔB_{t-5}			−.437(0.314)		
Δy_t	−.739(0.567)	−.704(0.381)			
Δy_{t-1}	−.010(0.580)				
Δy_{t-2}	.435(0.494)	.518(0.313)			
Δy_{t-3}				−.392(0.223)	
Δy_{t-4}					
Δy_{t-5}				.358(0.154)	
Δh^*_t	−.183(0.230)				
Δh^*_{t-1}	.139(0.271)				
Δh^*_{t-2}	.495(0.450)	.505(0.288)			
Δh^*_{t-3}					
Δh^*_{t-4}					
Δh^*_{t-5}					.400(0.115)
$CI2_{t-1}$	−.086(0.104)	−.177(0.111)			
$CI3_{t-1}$	−.138(0.361)	−.106(0.056)			
R^2	.9032	.8937	.5896	.4347	.4680
$\hat{\sigma}$	.0533	.0476	.0967	.0385	.0674
F, df	9.3352(16, 16)	18.494(10, 22)	6.226(6, 26)	7.433(3, 29)	
DW	1.62	1.60	2.01	2.55	1.68
Normality χ^2	.2379	1.848	6.130*	1.009	.145
AR1-2 F, df	1.9397(2, 14)	1.77(2, 20)	.168(2, 24)	1.580(2, 27)	.925(2, 29)
ARCH 1 F, df	.205(1, 14)	.1306(1, 20)	.561(1, 24)	.869(1, 27)	.634(1, 29)
RESET F, df	4.747(1, 15)*	4.571(1, 21)*	.027(1, 25)	.000(1, 28)	1.194(1, 30)

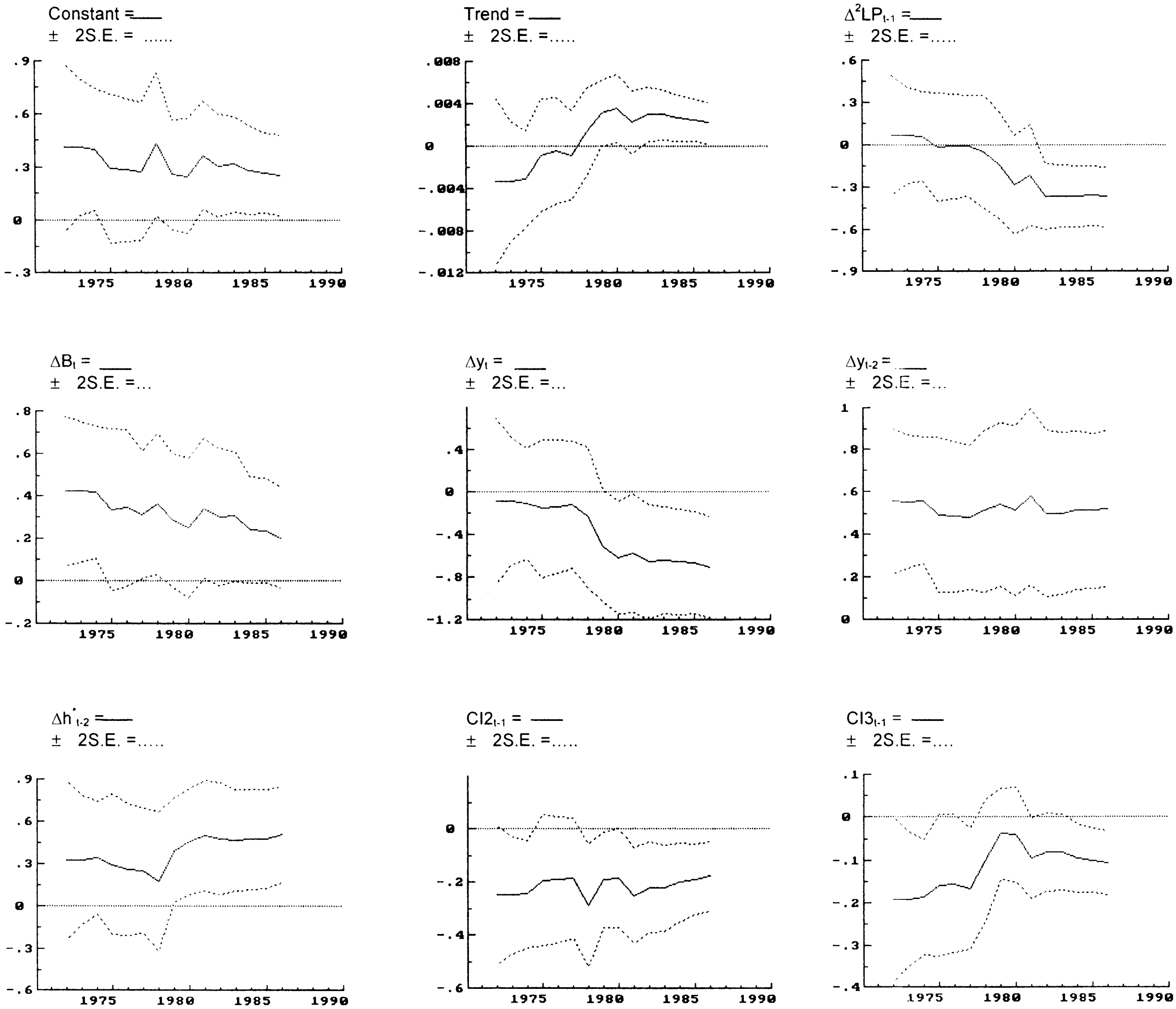
NOTE: The diagnostic checks for residual autocorrelation (AR 1-2F test with the degrees of freedom shown) confirm the choice of relevant lag, residual heteroscedaticity of the ARCH form (ARCH 1 *F* test) suggested by Engle (1982). RESET-*F* is a regression specification test. It tests the null of correct specification of the original model against the alternative that powers of the dependent variable are present.

White (1980) estimated standard errors are in parentheses. $\Delta^2 p$ depends on its own first lag and the current scaled public-sector deficit. It is also influenced by real income growth, its second lag, and the lagged monetization of the economy. The time trend and dummies have an impact on inflation. Equation (8) suggests a positive relationship between inflation and an appropriately scaled deficit. The ECM's explain the behavior of inflation by revealing relatively rapid reactions. This model closely matches the theory model and appears statistically satisfactory from the diagnostic tests except for the RESET F.

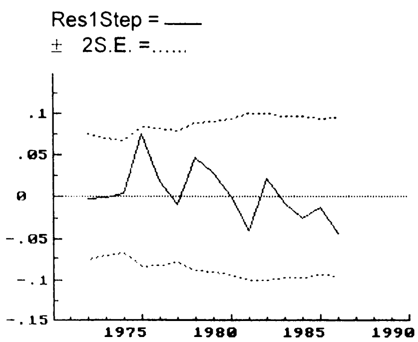
Parameter constancy is also an important statistical property. To examine constancy, recursive least squares is used because sequences of constancy tests yield tools for investigating constancy from the corresponding one-step innovations. From the sequence of innovations, Chow tests can be constructed for parameter constancy [distributed as $F(1, t-k-1)$ on the null]. Graphs provide a convenient way of portraying evidence about constancy. Figure 5 shows the recursively estimated coefficients of variables in (8) and plus or minus twice their recursively estimated standard errors. Coefficients vary only slightly relative to their ex ante standard errors. Figure 5 also records one-step residuals and corresponding calculated equation standard errors for conditional inflation equation with 0 ± 2 estimated standard errors. The equation standard error varies little. Figure 5 finally plots the breakpoint Chow (1960) statistic for the inflation equation, which remains constant over the sample period considered.

4.2 Nonconstancy of Marginal Models

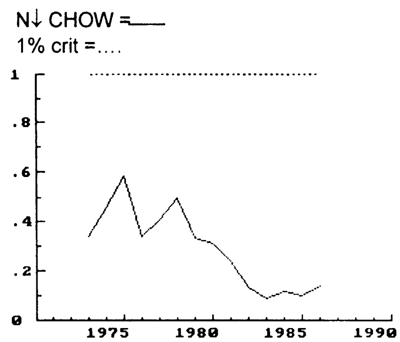
Nonconstancy of the marginal models is related to the concept of super exogeneity, which implies that the parameters of the conditional model remain constant, even while those of the marginal model change (i.e., the Lucas critique does not hold). This subsection estimates marginal models for Δy, ΔB, and Δh^*. Because of the results in Section 3.2, the parameters of interest here include just the parameters



(a)



(b)



(c)

Figure 5. (a) The Recursive Estimates for Nine Coefficients; (b) One-Step Residuals From a Conditional Model for Δp With 0 ± 2 Estimated Standard Errors; (c) Breakpoint Chow Statistics for a Conditional Model of Δp Normalized by Their One-off 1% Critical Values.

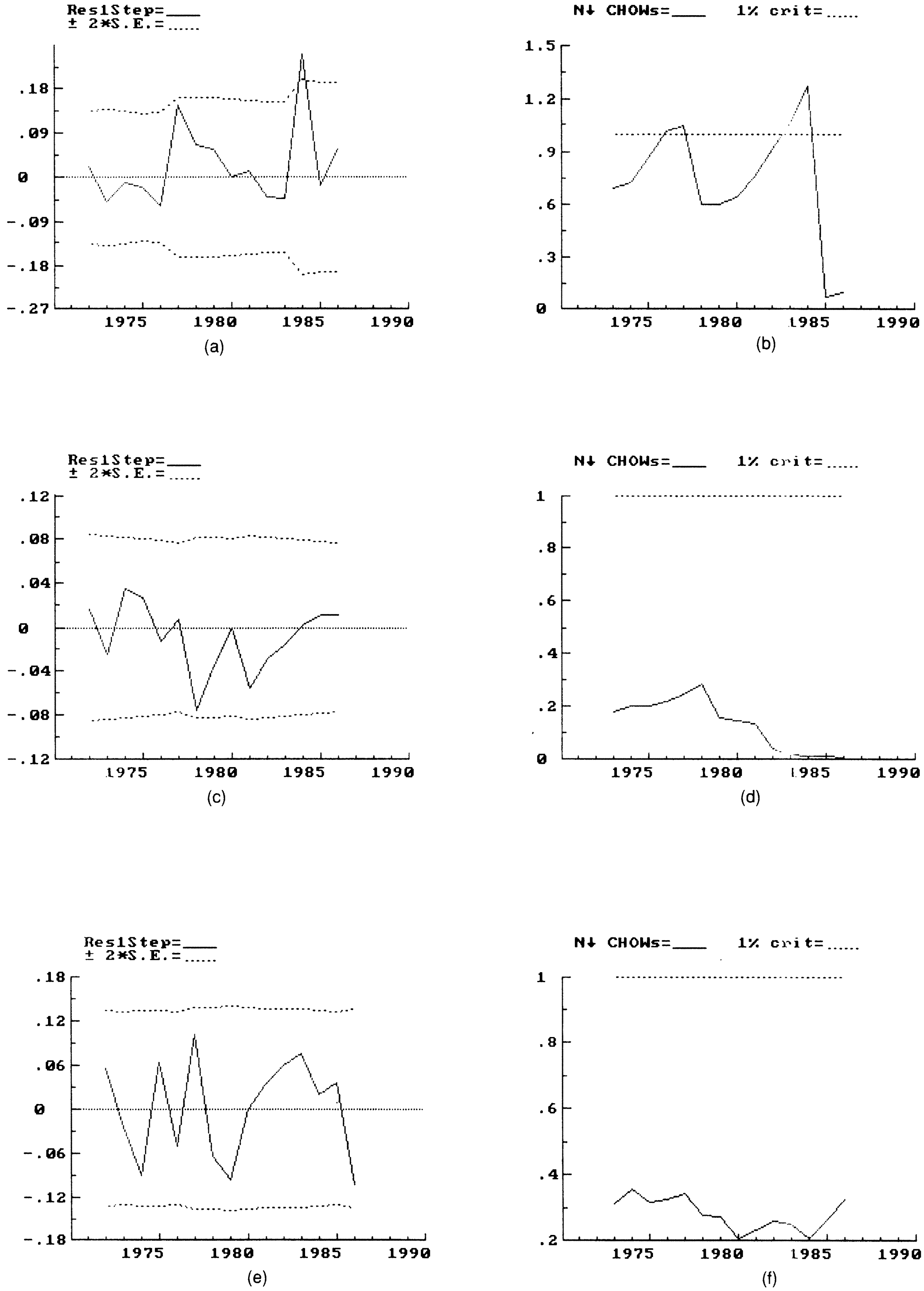


Figure 6. (a) One-Step Residuals From a Marginal Model for ΔB With 0 ± 2 Estimated Standard Errors; (b) Breakpoint Chow Statistics for a Marginal Model of ΔB Normalized by Their One-off 1% Critical Values; (c) One-Step Residuals From a Marginal Model for Δy With 0 ± 2 Estimated Standard Errors; (d) Breakpoint Chow Statistics for a Marginal Model of Δy, Normalized by Their One-off 1% Critical Values; (e) One-Step Residuals From a Marginal Model for Δh^* With 0 ± 2 Estimated Standard Errors; (f) Breakpoint Chow Statistics for a Marginal Model of Δh^*, Normalized by Their One-off 1% Critical Values.

for dynamics in the conditional model. For each marginal variable, we began with fifth-order autoregression (including a constant, trend, and $i1980$) and applied a sequential reduction procedure. The results are reported in Table 4, columns 4–6. For ΔB all lags matter. The residuals are nonnormal. Figure 6, (a) and (b), graphs the one-step resid-

Table 5. Encompassing Test Statistics for Equation (8) and Metin's (1995) Equation (9)

	Null hypothesis			
	Equation (8)		Metin (1995)	
Statistic	Distribution		Distribution	
Cox	$N(0, 1)$	−2.75	$N(0, 1)$	−7.39
Ericsson	$N(0, 1)$	1.86	$N(0, 1)$	3.95
Sargan	$\chi^2(6)$	6.79	$\chi^2(9)$	14.71
F	$F(6, 15)$	1.19	$F(9, 15)$	2.63
$\hat{\sigma}$		.0487%		.0601%

NOTE: T = 1954–1986.

uals and the sequence of breakpoint Chow statistics, which show considerable nonconstancy, with possible breaks in 1977 and 1984.

For Δy, the third and fifth lags matter. Statistically, the model appears well specified with no rejections from the diagnostic tests available. Figure 6, (c) and (d), plots the recursively estimated equation standard errors and the breakpoint Chow statistics. The marginal model of Δy appears constant.

For Δh^*, only the fifth lag matters. The equation is statistically satisfactory, and it appears constant [Fig. 6, (e) and (f)]. Because the conditional model for $\Delta^2 p$ is constant and the marginal model of ΔB is nonconstant, ΔB (at least) appears super exogenous for the dynamic parameters in the inflation equation.

4.3 Encompassing Implications of the Conditional Model

A congruent model should encompass previous empirical findings explaining the same dependent variable (see Hendry and Richard 1982, 1989; Mizon and Richard 1986). Consider two rival explanations, denoted M1 and M2. The question was whether M2 can explain features of the data that M1 cannot. This can be a test of M1, with M2 providing an alternative to see whether M2 captures any specific information not embodied in M1 (see Doornik and Hendry 1994, p. 237). Several variants of encompassing have been proposed—variance (Cox 1961), parameter (Hendry 1983), reduced-form (Ericsson 1983), exogeneity (Hendry 1988), and forecast (Chong and Hendry 1986). In this subsection we compare Equation (8) with an inflation equation estimated by Metin (1995), using such encompassing tests. The model from Metin (1995) is

$$\begin{aligned}\Delta p_t = & -\underset{[.039]}{.064} + \underset{[.135]}{1.111}B_t - \underset{[.670]}{3.901}\Delta((G-T)/Y)_t \\ & + \underset{[.362]}{1.663}\Delta p_{w_t} + \underset{[.099]}{.229}\Delta\text{ECM-M}_t \\ & - \underset{[.093]}{.272}(\text{ECM-UIP})_t/2 + \underset{[.044]}{.074}\text{ECM-PPP}_{t-1} \\ & + \underset{[.020]}{.257}d55_t - \underset{[.166]}{.234}\Delta y_t, \end{aligned} \tag{9}$$

where $R^2 = .8973, \hat{\sigma} = .0601$, DW $= 2.072$, AR(2,26) $= .55$, ARCH: $F(1,26) = 2.77$, normality: $\chi^2(2) = 1.33$, and RESET: $F(1,27) = 3.74$. In the work of Metin (1995), ECM represents sectoral excess demands, where ECM-M, ECM-PPP, and ECM-UIP were derived from the monetary sector, from purchasing power parity, and from uncovered interest-rate parity, and $d55$ is a dummy variable, which picks up a major outlier in 1955. Finally Δp_w is consumer price index (CPI) inflation for industrial countries. Table 5 reports the encompassing test results. As shown in Table 5, Equation (8) variance dominates Equation (9) (.00487 vs. .0601). None of the encompassing tests reject (8), and all reject (9); the new model encompasses the old one. (Note that Δp_{t-1} was added to (8) to calculate the encompassing tests.)

5. CONCLUDING REMARKS

This article examines the relationship between the public-sector deficit and inflation. System cointegration analysis suggests three stationary relationships. Although weak exogeneity does not hold for variables concerned (except Δy), one is still able to develop a conditional model for inflation. In that model, an increase in the scaled budget deficit immediately increases inflation. Real income growth has a negative immediate effect and positive second-lag effect on inflation. Monetization of the deficit also affects inflation at a second lag. These dynamics are consistent with institutional and general knowledge of the economy. The conditional model of inflation is constant over the sample period, even though several significant structural breaks occurred during the period. Breaks included three devaluations, structural stabilization, and economic liberalization programs. As further evidence of its specification, the new conditional model of inflation encompasses the inflation equation of Metin (1995). The major finding from the new equation is that budget deficits (as well as real income growth and debt monetization) significantly affect inflation in Turkey.

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APPENDIX: DATA

This appendix describes the data, lists the definitions used, and gives their units and sources. The sample period is 1950–1987.

G, T: The budget expenditure (G) and the revenue (T) are the general budget expenditures and revenues from the budget and final accounts, respectively (TL Billion). *Ministry of Finance and Custom General Directorate of Accounting, Statistical Year Book of Turkey 1990,* State Institute of Statistics Prime Ministry Republic of Turkey, Table No. 367, page 471.

$G - T$: The general budget deficit is the general budget expenditure minus the general budget revenue—that is, the primary deficit, which excludes interest payments (TL Billion). The budget deficit does not include the SEE's deficit. Because reliable statistics about SEE's deficits are available only after the second half of the 1970s, the general budget deficit is therefore used as a proxy for the total deficit.

P: Price level is the CPI. The base year is 1980 (*IMF International Financial Statistics,* several issues).
Y: *Y* is nominal GNP, divided by the GNP deflator (TL Billion). Nominal GNP is obtained from *IMF International Financial Statistics,* several issues.
H: *H* is base money. The components of base money are currency in circulation, vault cash, legal reserves, and Central Bank sight deposits (TL Billion). Reserve money is obtained from the database of the Central Bank of Turkey.

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