Journal of Monetary Economics 23 (1989) 219–238. North-Holland

U.S. FEDERAL INDEBTEDNESS AND THE CONDUCT OF FISCAL POLICY*

Jeroen J.M. KREMERS

International Monetary Fund, Washington, DC 20431, USA

Received January 1988, final version received December 1988

This paper examines whether the conduct of U.S. fiscal policy since the 1920s has been influenced by constraints on the accumulated stock of outstanding Federal debt. The evidence is affirmative: during most of the inter- and post-war period, policy in effect stabilized the Federal debt relative to GNP. But after 1981, large debt and debt service costs failed to produce deficit reductions consistent with the pattern of previous decades.

1. Introduction and synopsis

The large U.S. Federal budget deficits of recent years have featured prominently in the outlook for U.S. and world economic growth. Two key empirical questions are: (1) does the dynamic pattern of deficits have economic effects? and (2) are there, in a longer perspective, limits on the degree to which the government can rely on debt accumulation to finance its expenditure? These short- and long-run aspects of public deficits are clearly not unrelated, but empirical research has tended to address them separately. For example, Hamilton and Flavin (1986) introduced the use of stationarity tests to analyze long-run limitations on debt financing, but Barro's (1979) dynamic model of the determination of Federal deficits, on which several studies of their economic effects have relied,[1] does not satisfy the requirement that the Federal

*This paper is based on Chapter 4 of my D. Phil. thesis at Nuffield College, Oxford and on IMF working paper 87/27. It benefited from comments by E. Borensztein, L. Bovenberg, W.M. Corden, N. Ericsson, C. Granger, J. Hamilton, D. Hendry, P. Masson, J. Muellbauer, D. Wilcox, and seminar participants at the Federal Reserve Board in Washington, DC and the European Meeting of the Econometric Society in Copenhagen. The editors and the referee made valuable suggestions for improvement. The views expressed are mine alone and not necessarily those of the IMF.

[1] For instance, Barro (1980), Canarella and Garston (1983), Poterba and Summers (1987).

debt remain within the limits of the government's collateral – even though Barro has recognized the potential relevance of this requirement.[2]

Theoretical considerations bearing on long-run stability requirements and the constraints they impose on short-run budgetary management are well known (some are summarized below). Moreover, from a more practical perspective both former IMF Managing Director de Larosière and former Federal Reserve Chairman Volcker placed particular emphasis on linkages between short- and long-run aspects of budgetary policy:

> '[...] in spite of a growing awareness of the problem of debt accumulation, there are still too many observers who judge fiscal deficits exclusively in terms of their immediate impact on the economy, ignoring their long-run consequences' [de Larosière (1984)].

> 'Ultimately, debt can only be serviced from income. If that relationship is strained, financial pressures will jeopardize further growth in income itself, aggravating the difficulties. The time to act is before the strains become oppressive, not after' [Volcker (1986)].

This paper examines whether long-run constraints on the accumulated stock of outstanding Federal debt have influenced the annual conduct of budgetary policy since the 1920s. The evidence is affirmative: during most of the inter- and post-war period policy in effect stabilized the Federal debt relative to GNP. However, budgetary policy appears to have changed after 1981, as large debt and debt service costs failed to produce deficit reductions consistent with the pattern of previous decades.[3]

Section 2 of the paper discusses the government's intertemporal budget constraint and summarizes theoretical arguments relating to the optimal pattern of debt/GNP stabilization. It also sets out the econometric background of this paper, which was motivated in part by concern with conventional tests for stationarity of the Federal debt. These may have low power on the relatively short post-war sample analyzed by Hamilton and Flavin and others. Therefore, in addition to extending the sample to include the inter-war period, the paper follows an alternative, dynamic modeling strategy with more promising power characteristics. The next two sections present the empirical results, concentrating on long-run aspects in section 3 followed by the dynamic analysis in section 4.

[2]See Barro (1976, pp. 343–345; 1979, fn. 3 and 4; and 1986a, b, fn. 9).

[3]Barro (1986a, b) argued that these deficits were not exceptionally large, but noted subsequently (1987a, b) that his model produced large residuals after 1983. The view that the Reagan deficits constituted a change of policy was shared by von Furstenberg (1983) and Poterba and Summers (1987).

2. Analytical framework

This section first briefly reviews the dynamics of the intertemporal budget constraint.[4] For the purpose of transparency, the introduction of random shocks to the budget is postponed until the second part of the section.

The government's budgetary policy for period t must satisfy

$$\Delta B_t = G_t - T_t + r_{t-1}B_{t-1}. \tag{1}$$

B_t is the stock of interest-bearing public debt outstanding at the end of period t; G_t is government expenditure net of interest payments; T_t is tax revenue; r_t is the effective rate of interest paid on the public debt; and Δ is the first-difference operator. All variables are in real terms. Money financing is not considered explicitly, although it can be thought of as being part of tax revenue T_t.

Recursive forward substitution of (1) from period t to period $n > t$ yields the following expression for public debt in terms of future debt and primary surpluses:

$$B_t = R_{t,n}B_n + \sum_{j=t+1}^{n} R_{t,j}(T_j - G_j). \tag{2}$$

Discount factors are given by $R_{t,j} = \prod_{h=t+1}^{j} 1/(1 + r_{h-1})$. If budgetary policy follows a rule which ensures that the real debt will grow asymptotically at an average rate smaller than the asymptotic rate of interest (denoted by r), then the first term on the right-hand side of (2) will converge to zero as n becomes large. In that case budgetary policy respects the present-value budget constraint (3) below, which maintains that, viewed from any period t, future primary surpluses will be needed in order to service a currently positive stock of debt:

$$B_t = \sum_{j=t+1}^{\infty} R_{t,j}(T_j - G_j). \tag{3}$$

Hamilton and Flavin (1986) failed to reject requirement (3) using annual U.S. data from 1960 to 1984, and Hakkio and Rush (1986) supported that result with tests on quarterly data from 1950 to 1986.

With a currently finite stock of outstanding debt, the sequence of primary surpluses consistent with (3) must grow at an asymptotic rate below r as well.

[4]The exposition of the first part of the section benefited considerably from suggestions made by the referee. More extensive discussions can be found in Sargent and Wallace (1981), McCallum (1984), and Blanchard et al. (1985).

However, the path of the debt and the primary surplus associated with this policy may involve taxes that grow faster than taxing capacity, thus violating the government's collateral. It seems reasonable to assume that taxes (corrected for essential non-interest spending) are limited to a fraction, τ, of GNP (denoted by Y_t).[5] Then the government's collateral at any t amounts to

$$C_t = \tau Y_t \sum_{j=t+1}^{\infty} Q_{t,j}. \tag{4}$$

Discount factors are given by $Q_{t,j} = \prod_{h=t+1}^{j}(1+\rho_{h-1})/(1+r_{h-1})$, where $\rho_t = Y_{t+1}/Y_t - 1$. The collateral remains finite relative to GNP if the latter grows asymptotically at a rate ρ smaller than r.[6] This implies that, in order for taxes to remain feasible and the debt to remain bounded by the government's collateral, both must grow asymptotically at rates at most equal to ρ.[7] If ρ is indeed smaller than r, then (3) is necessary but not sufficient for a policy that respects both the present-value constraint and the government's taxing capacity.

The present paper deals with a condition that is both necessary and sufficient, namely the boundedness of the ratio of debt or debt service to GNP.[8] Within the limits set by this condition, the government decides on its plan for the intertemporal pattern of primary surpluses. But before proceeding to a closer look at that pattern, it is important to consider how the available finite sample can best be exploited for inference regarding this long-run aspect of policy. The theory of cointegrated variables deals with this issue in general terms.

The idea underlying cointegration is to study the behavior of variables moving apart in the short run, but brought together again by government policy or market forces, or both, if they continue to be too far apart in the longer run. A variable is defined to be integrated of order one [denoted by

[5]Similar assumptions were made by Barro (1976), Sargent and Wallace (1981), Blanchard (1984), Blanchard et al. (1985), and Masson (1985).

[6]For empirical evidence, see Abel et al. (1986). If the rate of GNP growth is larger than the rate of interest, the government's collateral is infinite. Most studies, including those referred to in footnote 5, have been based on the premise that the collateral is finite. And even if during a prolonged period of time ρ exceeded r, it would, in the presence of uncertainty regarding future growth, still seem a matter of prudence to keep the debt/GNP ratio within limits, validating the main argument of this paper.

[7]McCallum (1984) presented a general equilibrium model in which the tax/GNP and debt/GNP ratios grow without bound, but not the ratios of tax, interest, and debt to households' disposable income (defined to include the unbounded interest on public debt). Nevertheless, McCallum noted that the latter is of little economic significance in these circumstances; default and tax-evasion motives would presumably further restrict taxation and debt relative to GNP, as presumed in this paper.

[8]As suggested in footnote 6, this condition probably remains relevant even if ρ were to exceed r during a prolonged period of time.

$I(1)$] if it must be differenced once to induce stationarity [denoted by $I(0)$].[9] A set of variables, each $I(1)$, is said to be cointegrated if a linear combination of them is $I(0)$ (i.e., if they 'move together in the long run'). That linear combination is characterized by a cointegrating vector.

Natural logarithms of the debt/GNP and interest/GNP ratios are related by the identity[10]

$$(b-y)_t = (i-y)_t - \ln(r_{t-1}) + \Delta b_t. \tag{5}$$

Lower-case letters denote natural logarithms of the corresponding upper-case variables, and $I_t = r_{t-1}B_{t-1}$ denotes interest payments. Provided b_t is $I(1)$ so that Δb_t is $I(0)$, it is clear from (5) that stationarity of the debt/GNP ratio is equivalent to stationarity of the interest/GNP ratio if the rate of interest is stationary as well.

However, if the rate of interest (or Δb_t) is stationary with a component moving at a frequency that is low relative to the finite-sample period, then it is possible that nonstationarity of one of the two ratios cannot be rejected, even if in a longer perspective both are stationary. Rejection of nonstationarity of at least one of the two ratios can then still be interpreted as finite-sample evidence for stationarity of both in a longer perspective, given the assumption that in the long run the rate of interest and Δb_t remain $I(0)$.[11]

In any case, it is evidently important to employ tests with high power to reject nonstationarity. This justifies the analysis of long-run aspects of budgetary policy embedded in their dynamic context, for the following argument due to Banerjee et al. (1986) and Engle and Granger (1987). Two complementary approaches to testing for cointegration have been proposed in the literature. One consists of applying the usual univariate tests for (unit root) nonstationarity to deviations of the data from their long-run cointegrating combination, and the other involves testing for the presence of equilibrating feedback in the context of a multivariate dynamic specification. In the present study the scalar that cointegrates debt, or interest payments, with GNP is assumed to be given at unity. Thus, following the first approach, the variable

[9]Provided the variable has no deterministic component – see Granger (1986), Hendry (1986), and Engle and Granger (1987) for introductions to cointegration. Note that for such a variable to be bounded it is necessary but not sufficient that it is stationary. Thus, the logic of tests conducted in the literature on this subject, including this paper, is that boundedness cannot be accepted if nonstationarity cannot be rejected.

[10]Stationarity of a variable is equivalent to stationarity of its natural logarithm, provided the latter exists [Granger (1986)].

[11]Bhargava and Dickey–Fuller tests (introduced below) do not lead to rejection of the maintained hypothesis that the real effective rate of interest on the Federal debt (defined in correspondence with real interest payments as in appendix 1) and the growth rate of the real Federal debt (the nominal debt deflated by the CPI) are stationary variables, on annual data covering the combined inter- and post-war period. Test results are available from the author.

to be tested for nonstationarity will for instance be $\hat{u}_t$, representing deviations of the debt/GNP ratio from its mean $\hat{c}_0$:

$$\hat{u}_t = b_t - y_t - \hat{c}_0. \tag{6}$$

As a basis for the second approach to testing for cointegration, Engle and Granger (1987) established that cointegration of, say, debt and GNP is necessary and sufficient for the existence of dynamic error-correction models for the short-run pattern of debt or GNP, or both:

$$\Delta b_t = \alpha_1 \hat{u}_{t-1} + \text{short-run dynamics } [I(0)], \tag{7a}$$

$$\Delta y_t = \beta_1 \hat{u}_{t-1} + \text{short-run dynamics } [I(0)], \tag{7b}$$

with $|\alpha_1| + |\beta_1| \neq 0$. This approach tests whether the coefficient of at least one of the two feedbacks from past disequilibria is significantly different from zero.[12] Banerjee et al. (1986) and Engle and Granger (1987) showed that, in finite samples, the second approach may have larger power to reject nonstationarity than the first. In this paper both will be applied in a complementary fashion.

The section concludes by considering how long-run limitations on budgetary policy may be reflected in its short-run pattern. Error-correction mechanisms of the type introduced above will provide the linkage between both of these aspects of budgetary policy, by injecting an element of contingency to ensure consistency with the government's collateral in the presence of random shocks.[13] As indicated by (7), feedback could operate directly through the determination of deficit policy ($\alpha_1 < 0$), indirectly through the determination of GNP ($\beta_1 > 0$), or through both these channels. This paper focuses on the former.

At a theoretical level, the operating characteristics of various budgetary policy rules with a constant rate of debt-related error-correcting feedback[14] have been analyzed by Buiter (1984), Sachs and Wyplosz (1984), and Masson (1987). However, Blanchard et al. (1985) argued that the optimal rate of feedback depends on the state of the economy, with slower fiscal adjustment taking place during periods of recession. Indeed, as shown in section 4 below,

[12] Monte Carlo evidence, for instance in Engle and Granger (1987), suggests that under the null of no cointegration the usual t-statistic on the error-correction variable may not have the usual asymptotic distribution. However, Banerjee et al. (1986) obtained Monte Carlo results indicating that, remarkably, at the 5 percent level the usual t-test may have about the correct size.

[13] See Salmon (1982) and Nickell (1985) for the theoretical desirability of error correction under forward-looking behavior in a random environment.

[14] I.e., with no influences on the growth of debt other than from disequilibrium feedback with constant coefficient [no short-run dynamics in (7a) with a constant $\alpha_1 < 0$].

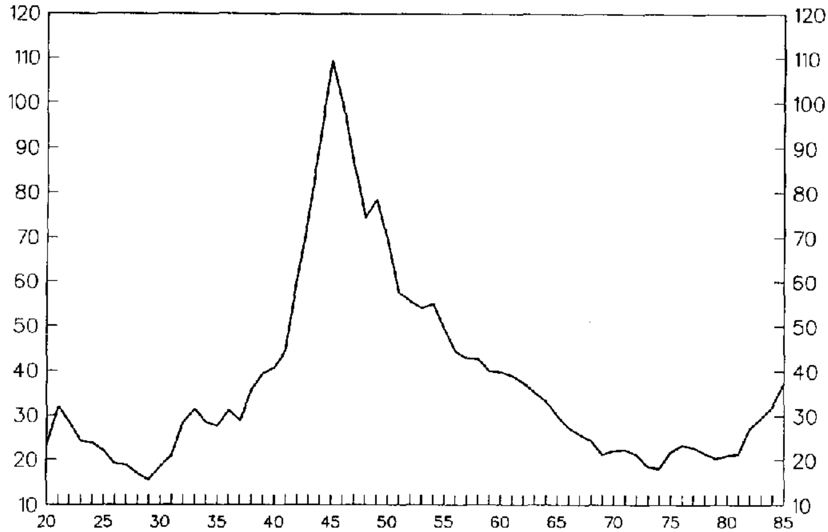

Fig. 1. The ratio of the Federal debt to GNP (B/Y). Data are annual, in percent, from 1920 to 1985. *Source*: appendix 1.

cyclical variations in the rate of Federal debt stabilization occurred during most of the inter- and post-war period. Similarly, Masson (1985) suggested that the rate of feedback may be related to the level of real interest rates, particularly if the latter are perceived to be affected by excessive public indebtedness.

3. Empirical results: The long-run

This section provides a first look at the long-run tendencies of the debt/GNP and interest/GNP ratios, which are depicted in figs. 1 and 2, respectively.[15] Interest payments were corrected for inflation (and denoted IC_t) in order to remove the fraction corresponding to the nominal component of interest rates [see, e.g., Jump (1980) and Eisner and Pieper (1984)].

The graphical evidence suggests a sustained tendency to return to a stationary long-run equilibrium on the part of the real interest/GNP ratio, which has fluctuated relatively tightly around an average of about $\frac{3}{4}$ percent since the 1920s. This stands in apparent contrast with the debt/GNP ratio, which has

[15]See appendix 1 for precise descriptions of the data. The debt/GNP ratio based on the measure for Federal indebtedness of Hamilton and Flavin (1986) has fluctuated similarly to the measure employed here, albeit at a lower level. The main difference between the two ratios consists of the government's gold holdings and the Treasury's operating cash balance, both netted out by Hamilton and Flavin.

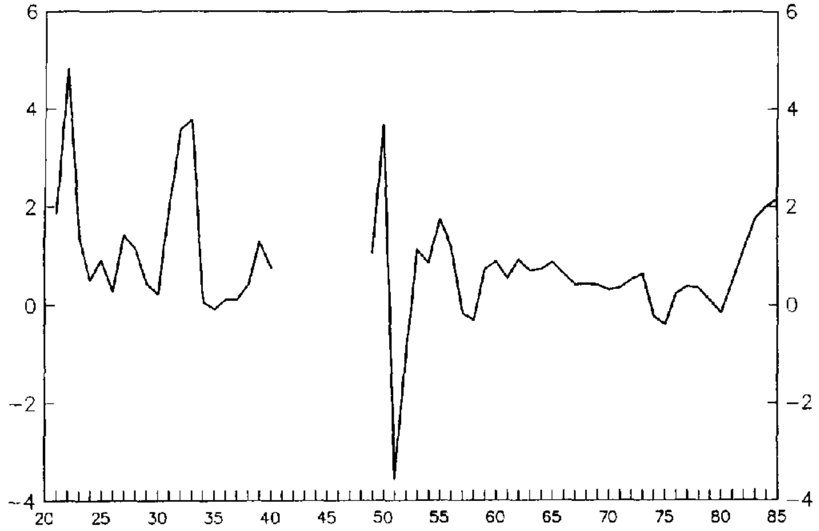

Fig. 2. The ratio of real interest payments on the Federal debt to GNP (IC/Y). Data are annual, in percent, from 1921 to 1985. *Source*: appendix 1.

crossed any possible equilibrium level only infrequently during this sample period.

The graphical impression is confirmed by formal tests: Table 1 reports Augmented Dickey–Fuller (ADF) (1981) and Bhargava (BHA) (1986) tests for nonstationarity of these ratios.[16] It appears that neither during the combined inter- and post-war period nor during the post-war period alone can nonstationarity of the debt/GNP ratio be firmly rejected. ADF test statistics for nonstationarity of the debt/GNP ratio over the entire sample [(1.5) and (1.6)] are close to the 10 percent critical level of about -2.60, suggesting that in a longer perspective the ratio may be stationary.

But there is firmer evidence against nonstationarity of the real interest/GNP ratio, as witnessed by BHA tests for subperiods and for the entire sample and by the ADF test for the post-war period (table 2). As regards the exact time period for which nonstationarity can be rejected, two results are worth noting. First, in part because of the small size of the inter-war sample, given the high order of autoregression, it proved difficult to estimate a satisfactory ADF

[16] These and other tests were discussed by Engle and Granger (1987), whose findings suggest that the ADF test is recommended if it is unknown whether the true model is of first or of higher order. Still, BHA has power in both cases as well. The finite-sample performance of these tests is quite sensitive with respect to the exact characteristics of the underlying true model – reason to consider both tests and see whether they point to the same conclusion. Throughout the paper, for reasons of data availability World War II data were excluded from the empirical analysis. All computations used the PCGIVE program by D.F. Hendry and the Oxford Institute of Economics and Statistics.

Table 1

Testing for nonstationarity: the debt/GNP ratio.[a]

Bhargava tests

(1.1) $\widehat{(B/Y)}_t = 0.32$, $BHA = 0.30$, $T = 1920\text{–}40, 1949\text{–}82$

(1.2) $\widehat{(B/Y)}_t = 0.32$, $BHA = 0.30$, $T = 1920\text{–}40, 1949\text{–}85$

(1.3) $\widehat{(B/Y)}_t = 0.31$, $BHA = 0.05$, $T = 1953\text{–}82$

(1.4) $\widehat{(B/Y)}_t = 0.31$, $BHA = 0.06$, $T = 1953\text{–}85$

Augmented Dickey–Fuller tests

(1.5) $$\Delta\widehat{(B/Y)}_t = \underset{(0.01)}{0.02} - \underset{(0.03)}{0.09}(B/Y)_{t-1} + \underset{(0.12)}{0.33}\Delta(B/Y)_{t-1}$$

$ADF = -2.70,^{*}$ $T = 1923\text{–}40, 1951\text{–}82$

(1.6) $$\Delta\widehat{(B/Y)}_t = \underset{(0.01)}{0.02} - \underset{(0.03)}{0.08}(B/Y)_{t-1} + \underset{(0.12)}{0.37}\Delta(B/Y)_{t-1}$$

$ADF = -2.55$ $T = 1923\text{–}40, 1951\text{–}85$

(1.7) $$\Delta\widehat{(B/Y)}_t = \underset{(0.01)}{0.02} - \underset{(0.03)}{0.07}(B/Y)_{t-1} + \underset{(0.21)}{0.27}\Delta(B/Y)_{t-1}$$

$ADF = -2.22$ $T = 1953\text{–}82$

(1.8) $$\Delta\widehat{(B/Y)}_t = \underset{(0.01)}{0.02} - \underset{(0.03)}{0.06}(B/Y)_{t-1} + \underset{(0.17)}{0.46}\Delta(B/Y)_{t-1}$$

$ADF = -1.86$ $T = 1953\text{–}85$

[a] B_t is the nominal market value of interest-bearing, privately-held Federal debt and Y_t is nominal GNP. Data sources are given in appendix 1. BHA is the Bhargava test statistic, with critical levels in Bhargava (1986, table 1). ADF is the Augmented Dickey–Fuller test statistic, with critical levels in Fuller (1976, table 8.5.2). Statistics rejecting nonstationarity at the 10 percent critical level are marked by one asterisk (*) and those rejecting at 5 percent by two (**). Conventionally computed standard errors are in parentheses. Regressions 5 to 8 have been checked for autoregressive errors up to fourth order.

equation for the interest/GNP ratio covering the combined inter- and post-war period. Second, all post-war tests appear to be sensitive with respect to the inclusion of 1983–1985. However, in view of the long-run character of these stationarity tests, they do not warrant precise inference regarding the time coverage of their results. Issues of timing are more appropriately addressed within the dynamic approach that follows.

4. Empirical results: The dynamics

The model of Federal deficits developed by Barro (1979) will be used as a starting point for the dynamic analysis. The background of this model is a neoclassical theory of tax smoothing, whose empirical specification and further elaborations by Barro (1986a, b) allow for countercyclical fluctuations of the

Table 2

Testing for nonstationarity: the interest/GNP ratio.[a]

Bhargava tests

(2.1) $\widehat{(IC/Y)}_t = 0.008$, $BHA = 1.53^{**}$, $T = 1921\text{–}40,\ 1949\text{–}82$

(2.2) $\widehat{(IC/Y)}_t = 0.008$, $BHA = 1.46^{**}$, $T = 1921\text{–}40,\ 1949\text{–}85$

(2.3) $\widehat{(IC/Y)}_t = 0.005$, $BHA = 1.00^{**}$, $T = 1953\text{–}82$

(2.4) $\widehat{(IC/Y)}_t = 0.007$, $BHA = 0.58$, $T = 1953\text{–}85$

Augmented Dickey–Fuller tests

$$\text{(2.5)}\quad \Delta\widehat{(IC/Y)}_t = \underset{(0.001)}{0.003} - \underset{(0.16)}{0.56}(IC/Y)_{t-1} + \underset{(0.12)}{0.46}\Delta(IC/Y)_{t-1} - \underset{(0.10)}{0.23}\Delta(IC/Y)_{t-2} - \underset{(0.04)}{0.20}\Delta(IC/Y)_{t-4}$$

$ADF = -3.48,^{**}\quad T = 1954\text{–}82$

$$\text{(2.6)}\quad \Delta\widehat{(IC/Y)}_t = \underset{(0.001)}{0.002} - \underset{(0.15)}{0.22}(IC/Y)_{t-1} + \underset{(0.14)}{0.41}\Delta(IC/Y)_{t-1} - \underset{(0.12)}{0.29}\Delta(IC/Y)_{t-2} - \underset{(0.05)}{0.21}\Delta(IC/Y)_{t-4}$$

$ADF = -1.42,\quad T = 1954\text{–}85$

[a] IC_t is nominal interest payments on the Federal debt, corrected for inflation, and Y_t is nominal GNP. Data sources are given in appendix 1. For further explanations, see table 1. Regressions 5 and 6 have been checked for autoregressive errors up to fourth order.

Federal budget as well. Using annual data from 1920 to 1982 without war years, the estimate of this model by Barro (1986a, table 6.1, set 2) can be closely approximated as follows:[17]

$$\widehat{\Delta pvb_t} = \underset{(0.006)}{0.010} + \underset{(0.09)}{0.98}\hat{\pi}^{e}_{t,t+1} - \underset{(0.29)}{3.96}\,YVAR_t + \underset{(0.10)}{0.29}\,GVAR_t, \tag{8}$$

$T = 1920\text{–}40,\ 1948\text{–}82,\quad R^2 = 0.97,\quad \hat{\sigma} = 2.38\%,\quad DW = 2.25,$

$AC(4)_1^4 = 8.88,\quad AC(1)_1^1 = 1.15,\quad AC(1)_2^2 = 3.02,$

$ARCH(3,48) = 1.07,\quad NORM(2) = 0.25,\quad CHOW(17,35) = 1.08.$

PVB_t is the nominal par value of interest-bearing, privately-held Federal debt; $\hat{\pi}^{e}_{t,t+1}$ is inflation anticipated from January of year t to January of year $t+1$; $YVAR_t$ is a cyclical variable (negative during recession); and $GVAR_t$ is

[17] See appendix 2 for brief explanations of the diagnostic statistics.

transitory Federal defense spending. Heteroscedasticity-consistent standard errors are in parentheses [White (1980)].

The parameter estimates do not differ significantly from those of Barro,[18] nor do the diagnostic tests (most of which were not reported by Barro) point to any misspecification. Actual and fitted values are depicted in the upper panel of fig. 3.

The model as it stands in (8) contains no mechanism to prevent the stock of outstanding debt from exceeding the government's collateral. One possible test for the presence of such a mechanism was conducted by Barro (1979, p. 966), who included the lagged debt/GNP ratio and found that it was insignificant. This can be confirmed in (8),[19] and, in light of section 3, is perhaps not surprising: over this sample period the debt/GNP ratio does not exhibit a typical stationary pattern.

However, if added to (8), feedback from the lagged real interest/GNP ratio is significant at a critical level of 5 percent:

$$\widehat{\Delta pvb_t} = \underset{(0.006)}{0.015} + \underset{(0.12)}{0.92}\,\hat{\pi}^{e}_{t,t+1} - \underset{(0.33)}{4.02}\,YVAR_t + \underset{(0.10)}{0.29}\,GVAR_t - \underset{(0.26)}{0.44}\,(IC/Y)_{t-1}, \tag{9}$$

$$T = 1922\text{–}40, 1950\text{–}82, \quad R^2 = 0.97, \quad \hat{\sigma} = 2.37\%, \quad DW = 2.07,$$

$$AC(4)^4_1 = 4.37, \quad AC(1)^1_1 = 0.12, \quad AC(1)^2_2 = 0.94,$$

$$ARCH(3,43) = 0.44, \quad NORM(2) = 0.10, \quad CHOW(17,30) = 0.96.$$

This suggests that, in accordance with the previous section, during most of the inter- and post-war period budgetary policy included a mechanism to avoid excessive accumulation of Federal indebtedness, by stabilizing the cost of debt service relative to GNP and thus indirectly stabilizing the debt/GNP ratio. However, historical evidence in Ahearn (1963) and Vatter (1963) would seem to indicate that the abolition in 1952 of the Federal Reserve's policy of pegging interest rates on Federal debt may have had repercussions for the possible dependence of debt/GNP stabilization on interest rates [cf. Masson

[18]Because the equation for inflation was not jointly estimated here, the estimated standard error of the coefficient on this variable is slightly smaller [Pagan (1984)] and the equation standard error is slightly larger.

[19]Included in (8), $(B/Y)_{t-1}$ generates an estimated coefficient of -0.01 with standard error 0.03.

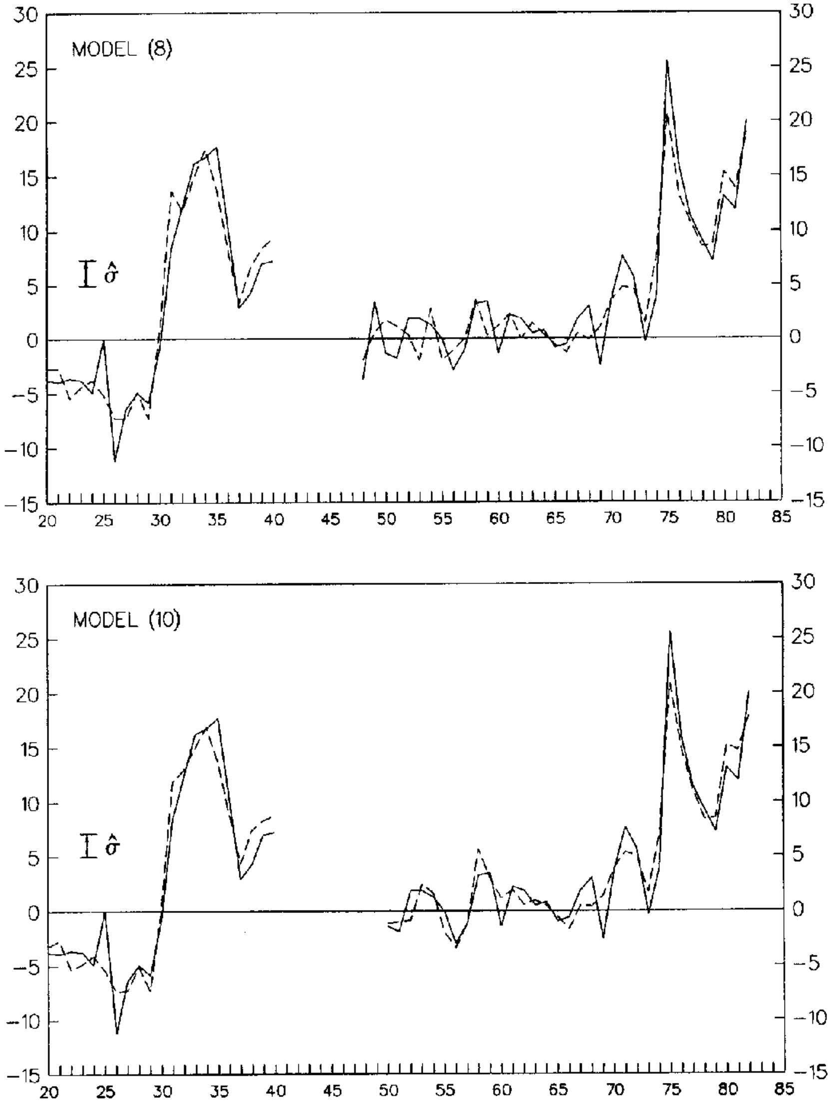

Fig. 3. Actual and fitted values for the nominal growth rate of the Federal debt. Data are annual, in percent, from 1920 to 1982. The solid line represents the actual values and the arrow measures the estimated standard error of the equation ($\hat{\sigma}$). *Source*: models (8) and (10).

(1985)]. To test for a change of feedback regime around that event, regression (8) was run including both the lagged debt/GNP and the lagged real interest/GNP ratios while allowing different coefficient estimates before and after 1952. That regression (not reported) indicated that, indeed, debt/GNP stabilization became dependent on the level of interest rates only after 1952. Imposing appropriate restrictions produces firm evidence of a change in feedback regime:[20]

$$\begin{aligned}\widehat{\Delta pvb_t} = & -\underset{(0.0141)}{0.0002} + \underset{(0.15)}{0.82}\,\hat{\pi}^e_{t,t+1} - \underset{(0.41)}{3.66}\,YVAR_t + \underset{(0.12)}{0.26}\,GVAR_t \\ & -\underset{(0.04)}{0.08}\,(B/Y)_{t-1}D3252 + \underset{(0.024)}{0.044}\,D3252 \\ & -\underset{(0.54)}{2.44}\,(IC/Y)_{t-1}D5382 + \underset{(0.016)}{0.029}\,D5382, \end{aligned} \tag{10}$$

$T = 1920\text{–}40, 1950\text{–}82,\quad R^2 = 0.98,\quad \hat{\sigma} = 2.16\%,\quad DW = 2.20,$

$AC(4)^4_1 = 8.38,\quad AC(1)^1_1 = 0.87,\quad AC(1)^2_2 = 1.36,$

$ARCH(3,42) = 0.48,\quad NORM(2) = 1.63,\quad CHOW(17,29) = 1.44.$

D3252 is unity during 1932–52 and zero elsewhere; *D5382* is defined in a corresponding manner. Actual and fitted values of (10) are in the lower panel of fig. 3.

The inclusion of debt-related feedback affects the Barro model in two dimensions. First, it improves the model in a statistical sense: the estimated error variance of (8) is reduced by 17 percent. Although the dummies for changing constants (*D3252* and *D5382*) form an integral part of the error-correction feedback mechanism (they allow for a difference in mean of the changing long-run targets), the statistical improvement clearly does not come from the dummies but from the two error-correction variables: the *F*-statistic

[20] Estimating from 1922, an *F*-test cannot reject (10) against a more general specification allowing different coefficients on both forms of feedback during 1922–31, 1932–52, and 1953–82 (the test statistic is 0.6 against a 5 percent critical value of 2.6). The results of direct (nonnested) Cox tests for variance encompassing [see Cox (1961) and Ericsson (1983)] are unambiguous: they reject (9) against (10), but fail to reject (10) against (9) (the test statistics are 13.1 and 0.7, respectively, against a 5 percent critical value of 2.0). No significant feedback was found during the 1920s, which may reflect the less than fully satisfactory quality of some of the data for that period.

for adding the two dummies to (8) is 0.93 (5 percent critical value is 3.19), but the F-statistic for including in addition the two error-correction variables is 6.05 (significant even below 1 percent). The F-statistic for adding all four variables together is 3.59 (significant at about 2 percent).

Second, the debt-related feedback affects the model in an economic sense. The presence in the dynamic model of this feedback, in conformity with the static analysis of section 3, suggests that the Federal debt and GNP are cointegrated.[21] Since cointegration is by its very nature a long-run rather than a short-run dynamic phenomenon, this is likely to affect the model's long-run properties more markedly than its short-run dynamics (including the error variance). In particular, given the findings of Engle and Yoo (1987), one would expect that incorporation of the apparently valid cointegration constraint, which was justified from a theoretical perspective in section 2, would improve the model's multi-step forecast performance. This points to possible policy implications and an avenue for further research. In a world of rational expectations, tests of the effects of deficit policy on key economic variables such as interest rates and economic growth must rely on models for expected policy. It might be true that the robustness of those tests can be enhanced by imposing the cointegration constraint, the more so to the extent that economic behavior is framed in a long-term perspective.

Turning, finally, to the large deficits of the 1980s, in accordance with Barro (1987a, b), Poterba and Summers (1987), and von Furstenberg (1983), the model including the cointegration constraint indicates that deficit policy changed under the Reagan Administration. The upper panel of fig. 4 shows that (10), when estimated up to 1965, predicts debt growth from 1966 to 1981 quite well [note the Chow test of (10)], but significantly underpredicts it subsequently (cf. $\hat{\sigma}$).

Sargent (1986) discussed two alternative ways of interpreting these large deficits. The first introduces an exogenous change of the long-run target for public spending into the Barro model:

> 'Assume that the election of Reagan signalled a downward revision in the size of the U.S. government, as measured by the expected present value of federal expenditures. Assume further that the path of reductions, compared to the path that could have been expected prior to Reagan, was skewed toward the future or 'back-loaded'. [...] Barro's tax smoothing model predicts that the (optimal) response of the government would be an immediate permanent reduction of tax collections, relative to the pre-Reagan path' [Sargent (1986, p. 7)].

[21]As discussed in footnote 12, under the null of no cointegration larger critical values than those for the usual t-test may be required at levels other than 5 percent. Note, however, that the t-value on the post-war error-correction variable is as large as 4.5.

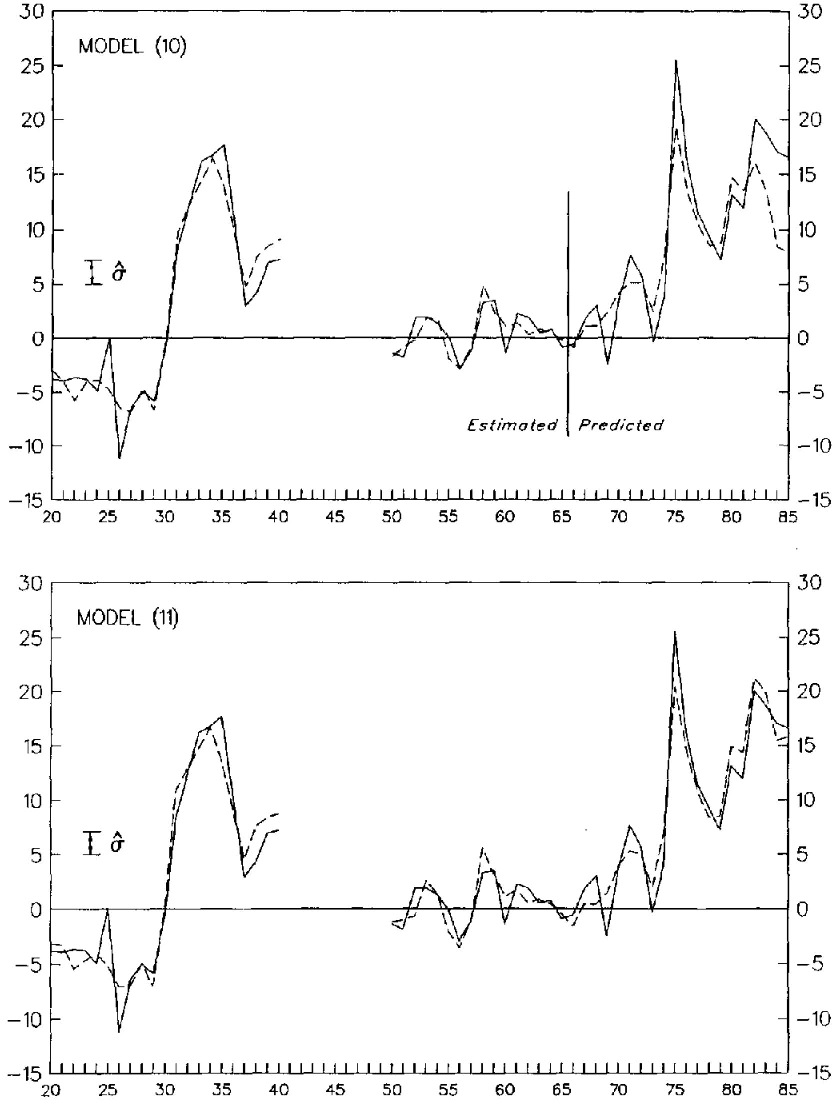

Fig. 4. Actual and fitted values for the nominal growth rate of the Federal debt. Data are annual, in percent, from 1920 to 1985. The solid line represents the actual values and the arrow measures the estimated standard error of the equation ($\hat{\sigma}$). *Source*: models (10) and (11).

The second interpretation, due to Neil Wallace, essentially views fiscal policy during the Reagan period as a game between the President (the taxing authority), whose objective it is to reduce public spending as in the first interpretation above, and the Congress (the spending authority), which desires higher spending than the President does.[22] Both are bound, in the longer term, by the intertemporal budget constraint. The President has opened the game by lowering taxes upfront so as to force the Congress into cutting expenditure. Whether Congress will feel impelled to do so depends on the credibility of the President's commitment to keeping taxes at the new low level.

In the first scenario, the exogenous drop in spending would enter the model through an initial jump in $GVAR$ (measuring the deviation of current from permanent spending), exerting upward pressure on the deficit until $GVAR$ would have returned to zero. The counterbalancing role of debt feedback, ensuring long-run policy feasibility, would not be affected. In the second interpretation, by contrast, a temporarily lower priority for debt-related feedback (combined with a temporary, exogenous jump in the deficit) could be a characteristic of the early stage of the game as credibility is being established. Heeding a caveat due to the limited degrees of freedom, regression (11) provides suggestive evidence for the latter interpretation (actual and fitted values are in the lower panel of fig. 4):

$$\begin{aligned}\widehat{\Delta pvb_t} = & -\underset{(0.013)}{0.002} + \underset{(0.15)}{0.83}\,\hat{\pi}^{e}_{t,t+1} - \underset{(0.42)}{3.42}\,YVAR_t + \underset{(0.12)}{0.22}\,GVAR_t \\ & - \underset{(0.04)}{0.09}\,(B/Y)_{t-1}D3252 + \underset{(0.024)}{0.053}\,D3252 \\ & - \underset{(0.53)}{2.63}\,(IC/Y)_{t-1}D5381 + \underset{(0.015)}{0.029}\,D5381 \\ & + \underset{(0.020)}{0.057}\,D8285, \end{aligned} \tag{11}$$

$$T = 1920\text{–}40, 1950\text{–}85, \quad R^2 = 0.98, \quad \hat{\sigma} = 2.12\%, \quad DW = 2.13,$$

$$AC(4)^4_1 = 6.55, \quad AC(1)^1_1 = 0.32, \quad AC(1)^2_2 = 0.97,$$

$$ARCH(3,41) = 0.65, \quad NORM(2) = 2.94.$$

F-tests for the inclusion of feedback from the debt/GNP or interest/GNP

[22] I abstract from the role of the Federal Reserve.

ratios in 1982–85 (not reported) indicate the insignificance of both. Should this interpretation of recent deficits be correct, then debt feedback could return to normal as soon as the game has a winner.

Appendix 1: Data sources

Flow variables correspond to calendar years, and stock variables are end-of-calendar-year.

IC_t $= I_t - \hat{\pi}^e_{t-1,t} B_{t-1}$.

I_t = Nominal net Federal interest payments [1921–28: calculations based on Dewey (1924), Newcomer (1937) and Ratchford (1947) (values: 0.997, 1.026, 1.056, 0.969, 0.882, 0.993, 0.924, 0.855); 1929–85: U.S. Department of Commerce, Survey of Current Business, table 3.2, line 22].

$\hat{\pi}^e_{t,t+1}$ = Inflation anticipated in period t for period $t+1$ [January values, 1920–40 and 1948–82: predictions from model in Barro (1986a, table 6.4); 1983–85: based on Barro (1986b, table 4) (values: 0.048, 0.066, 0.050)].

B_t = Nominal end-of-year market value of interest-bearing, privately-held Federal debt [1919–41: PVB_t multiplied by market-to-par ratio from Seater (1981, table 4); 1942–84: end-of-December market value from Cox (1985, table 2); 1985: obtained from W.M. Cox, Federal Reserve Bank of Dallas (value: 1483698)].

PVB_t = Nominal end-of-year par value of interest-bearing, privately-held Federal debt [1919–76: Barro (1979, table 3); 1977–85: IMF International Financial Statistics, 88-88aa, or Barro (1986a, table 6.4) for rates of growth].

Y_t = Nominal GNP [1919–85: U.S. Council of Economic Advisers, Economic Report of the President, table B.1].

$GVAR_t$ = Transitory Federal defense spending, scaled by moving average of PVB_t [1920–82: Barro (1986a, table 6.5); 1983–85: based on Barro (1986b, table 3) (values: -0.4; -0.49; -0.49)].

$YVAR_t$ = Business cycle variable based on the rate of unemployment [1920–82: Barro (1986a, table 6.5); 1983–85: based on Barro (1986b, table 3) (values: 0.033, 0.016, 0.021)].

Appendix 2: Test statistics

This appendix contains brief descriptions of the reported test statistics, with degrees of freedom in brackets. The methodological background of these and other diagnostic tests can be found in Hendry (1987) and Spanos (1986).

Autocorrelation

$AC(n)_i^j$ = Lagrange multiplier test for autocorrelation from lags i to j $(j-i+1=n)$, χ^2-form of Godfrey (1978) [see also Spanos (1986, p. 521)]. Computed by regressing the residuals on all the regressors of the original model and the lagged residuals for lags i to j, and testing the joint significance of the latter.

Heteroscedasticity

$ARCH(n,\cdot)$ = Lagrange multiplier test for nth-order autoregressive conditional heteroscedasticity, F-form of Engle (1982) [see also Spanos (1986, p. 548)]. Computed by regressing the squared residuals on the lagged squared residuals up to lag n and testing their joint significance.

Normality Residuals

$NORM(2)$ = χ^2-test of Jarque and Bera (1980) [see also Spanos (1986, p. 454)]. Based on the estimated skewness and kurtosis of the residuals compared to their counterparts of the normal distribution.

Parameter Constancy

$CHOW(n,\cdot)$ = F-test for parameter constancy between first $T-n$ and last n observations, based on residual sums of squares [Chow (1960), Spanos (1986, p. 486)].

References

Abel, A.B., N.G. Mankiw, L.H. Summers, and R.J. Zeckhauser, 1986, Assessing dynamic efficiency: Theory and evidence, NBER working paper no. 2097.

Ahearn, D.S., 1963, Federal Reserve policy reappraised, 1951–1959 (Columbia University Press, New York, NY).

Banerjee, A., J.J. Dolado, D.F. Hendry, and G.W. Smith, 1986, Exploring equilibrium relationships in econometrics through static models: Some Monte Carlo evidence, Oxford Bulletin of Economics and Statistics 48, 253–277.

Barro, R.J., 1976, Reply to Feldstein and Buchanan, Journal of Political Economy 84, 343–349.

Barro, R.J., 1979, On the determination of the public debt, Journal of Political Economy 87, 940–971.

Barro, R.J., 1980, Federal deficit policy and the effects of public debt shocks, Journal of Money, Credit, and Banking 12, 747–761.

Barro, R.J., 1986a, The behavior of United States deficits, in: R.J. Gordon, ed., The American business cycle: Continuity and change (University of Chicago Press, Chicago, IL).

Barro, R.J., 1986b, U.S. deficits since World War I, Scandinavian Journal of Economics 88, 195–222.

Barro, R.J., 1987a, Macroeconomics, 2nd ed. (Wiley, New York, NY).

Barro, R.J., 1987b, Budget deficits: Only a minor crisis, The Wall Street Journal, Jan. 16th.

Bhargava, A., 1986, On the theory of testing for unit roots in observed time series, Review of Economic Studies 53, 369–384.

Blanchard, O.J., 1984, Current and anticipated deficits, interest rates and economic activity, European Economic Review 25, 7–27.

Blanchard, O.J., R. Dornbusch, and W. Buiter, 1985, Public debt and fiscal responsibility, Paper no. 22 (Centre for European Policy Studies, Brussels).
Buiter, W.H., 1984, Fiscal policy in open, interdependent economies, NBER working paper no. 1429.
Canarella, G. and N. Garston, 1983, Monetary and public debt shocks, Journal of Money, Credit, and Banking 15, 199–211.
Cox, D.R., 1961, Tests of separate families of hypotheses, in: J. Neyman, ed., Proceedings of the fourth Berkeley symposium on mathematical statistics and probability, Vol. I (University of California Press, Berkeley, CA).
Cox, W.M., 1985, The behavior of Treasury securities, monthly, 1942–1984, Journal of Monetary Economics 16, 227–240.
de Larosière, J., 1984, The growth of public debt and the need for fiscal discipline, 40th Congress of the International Institute of Public Finance, Innsbruck.
Dewey, D.R., 1924, Financial history of the United States, 9th ed. (Longmans, London).
Dickey, D.A. and W.A. Fuller, 1981, Likelihood ratio statistics for autoregressive time series with a unit root, Econometrica 49, 1057–1072.
Eisner, R. and P.J. Pieper, 1984, A new view of the Federal debt and budget deficits, American Economic Review 74, 11–30.
Engle, R.F., 1982, Autoregressive conditional heteroscedasticity with estimates of the variance of United Kingdom inflation, Econometrica 50, 987–1008.
Engle, R.F. and C.W.J. Granger, 1987, Co-integration and error-correction: Representation, estimation, and testing, Econometrica 55, 251–276.
Engle, R.F. and B.S. Yoo, 1987, Forecasting and testing in co-integrated systems, Journal of Econometrics 35, 143–159.
Ericsson, N.R., 1983, Asymptotic properties of instrumental variables statistics for testing non-nested hypotheses, Review of Economic Studies 50, 287–304.
Fuller, W.A., 1976, Introduction to statistical time series (Wiley, New York, NY).
Godfrey, L.G., 1978, Testing against general autoregressive moving average error models when the regressors include lagged dependent variables, Econometrica 46, 1293–1301.
Granger, C.W.J., 1986, Developments in the study of cointegrated economic variables, Oxford Bulletin of Economics and Statistics 48, 213–228.
Hakkio, C.S. and M. Rush, 1986, Co-integration and the government's budget deficit, Research working paper no. 86-12 (Federal Reserve Bank of Kansas City, KS).
Hamilton, J.D. and M.A. Flavin, 1986, On the limitations of government borrowing: A framework for empirical testing, American Economic Review 76, 808–819.
Hendry, D.F., 1986, Econometric modelling with cointegrated variables: An overview, Oxford Bulletin of Economics and Statistics 48, 201–212.
Hendry, D.F., 1987, Econometric methodology: A personal perspective, in: T.F. Bewley, ed., Advances in econometrics, Vol. 2 (Cambridge University Press, Cambridge).
Jump, G.V., 1980, Interest rates, inflation expectations and spurious elements in measured real income and saving, American Economic Review 70, 990–1004.
Masson, P.R., 1985, The sustainability of fiscal deficits, IMF Staff Papers 32, 577–605.
Masson, P.R., 1987, The dynamics of a two-country minimodel under rational expectations, Annales d'Economie et de Statistique, no. 6/7, 37–69.
McCallum, B.T., 1984, Are bond-financed deficits inflationary? A Ricardian analysis, Journal of Political Economy 92, 123–135.
Newcomer, M., 1937, Analysis of the nature of American public debts, American Economic Review 27, suppl., 49–57.
Nickell, S.J., 1985, Error correction, partial adjustment and all that: An expository note, Oxford Bulletin of Economics and Statistics 47, 119–130.
Pagan, A., 1984, Econometric issues in the analysis of regressions with generated variables, International Economic Review 25, 221–247.
Poterba, J.M. and L.H. Summers, 1987, Finite lifetimes and the effects of budget deficits on national saving, Journal of Monetary Economics 20, 369–391.
Ratchford, B.U., 1947, History of the Federal debt in the United States, American Economic Review 37, suppl., 131–141.
Sachs, J. and C. Wyplosz, 1984, Real exchange rate effects of fiscal policy, NBER working paper no. 1255.

Salmon, M., 1982, Error correction mechanisms, The Economic Journal 92, 615–629.
Sargent, T.J., 1986, Interpreting the Reagan deficits, Economic Review of the Federal Reserve Bank of San Francisco, Fall, 5–12.
Sargent, T.J. and N. Wallace, 1981, Some unpleasant monetarist arithmetic, Quarterly Review of the Federal Reserve Bank of Minneapolis, Fall, 1–17.
Seater, J.J., 1981, The market value of outstanding government debt, Journal of Monetary Economics 8, 85–101.
Spanos, A., 1986, Statistical foundations of econometric modelling (Cambridge University Press, Cambridge).
Vatter, H.G., 1963, The U.S. economy in the 1950s (Greenwood Press, Westport, CT).
Volcker, P., 1986, Statement before the Subcommittee on Telecommunications, Consumer Protection and Finance of the Committee on Energy and Commerce of the U.S. House of Representatives, April 23rd.
von Furstenberg, G.M., 1983, Fiscal deficits: From business as usual to a breach of the policy rules?, The National Tax Journal 36, 443–457.
White, H., 1980, A heteroscedasticity-consistent covariance matrix estimator and a direct test for heteroscedasticity, Econometrica 48, 817–838.