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ON THE PITFALLS OF UNTESTED COMMON-FACTOR RESTRICTIONS: THE CASE OF THE INVERTED FISHER HYPOTHESIS

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A common approach to single-equation time-series modelling is to start with a static regression suggested by some economic theory; to estimate it using ordinary least squares; and, if the Durbin-Watson statistic is sufficiently below two, suggesting serial correlation in the residuals, to re-estimate it using a Cochrane-Orcutt transformation. Then, if the Durbin-Watson statistic looks respectable, one proceeds to hypothesis testing on the basis of the estimated standard errors of the coefficients of the (transformed) static regression. The Cochrane-Orcutt transformation imposes a common-factor restriction on the estimated regression. Although it is well-known that such restrictions may not be justified, the ease with which they are imposed in common econometric computer packages and the fact that many econometrics texts present the Cochrane-Orcutt transformation and its generalizations as a panacea for serial correlation means that unwarranted common-factor restrictions are routinely and wrongly imposed. It is too rarely appreciated that unwarranted common-factor restrictions may lead an investigator to draw false statistical inferences — *sometimes the opposite of what the data in fact support.*

In this paper, we shall review the theory of common factors. We shall then illustrate a case of a wrongly-imposed common-factor restriction using Carmichael and Stebbing's (1983) paper in support of the inverted Fisher hypothesis. Finally, we shall use the same data to illustrate a constructive alternative to blithely imposing common-factor restrictions.

I. THE COCHRANE-ORCUTT TRANSFORMATION AND COMMON-FACTOR RESTRICTIONS

A simple two-variable static regression model may take the form

$$y_t = \alpha_0 + \alpha_1 x_t + \varepsilon_t, \tag{1}$$

On estimating such a regression, an investigator may discover that the estimates of ε_t are serially-correlated. Econometrics textbooks tell him that serial

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correlation in the residuals produces inefficient estimates of the parameters α_0 and α_1 and biased estimates of their standard errors, seriously interfering with hypothesis testing (e.g., Johnston, 1972, pp. 246–49).

The simplest model of serial correlation is that it is first-order autoregressive — i.e.,

$$\varepsilon_t = \rho\varepsilon_{t-1} + v_t, \tag{2}$$

where v_t is assumed to be a serially uncorrelated random error. Adopting this model as his working hypothesis, the investigator can then apply a Cochrane-Orcutt transformation (usually by altering a single parameter in a regression package). The Cochrane-Orcutt transformation lags equation (1) by one period, multiplies through by ρ and subtracts the new equation from (1) to yield

$$y_t = \alpha_0(1-\rho) + \rho y_{t-1} + \alpha_1 x_t - \rho\alpha_1 x_{t-1} + v_t. \tag{3}$$

The error term of equation (3) is serially uncorrelated *ex hypothesi*; and consistent and efficient estimates of the parameters, α_0, α_1 and ρ, can be recovered from estimates of equation (3). Of course all of the desirable properties of the transformed estimates of (1) rest on the truth of the untested assumption that its error follows the stochastic process described in (2).

The practice of ridding regressions of serial correlation through Cochrane-Orcutt transformations rests on an odd notion of the nature of random error. It assumes that, when a regression such as (1) is found to have serially-correlated residuals, it should still be taken to be correctly specified as far as it goes. The only problem is to guess the nature of the stochastic process generating the error term in order to supplement the original equation — as if serial correlation were a natural process independent of the variables of the regression. The usual justification for equations such as (1) is that they are derived from theory. It is odd that few attempts are made to derive supplemental processes such as (2) from any theory.

A more reasonable view of random error would be that the estimated residuals are measures of our ignorance of the correct specification of the underlying process that generates the data. Then the discovery of serial correlation indicates that our regression is misspecified, that there are systematic improvements to be made in our specification. An equation such as (1) would need to be respecified so that its residuals were not serially correlated. The respecified equation would not necessarily be correct; although the original equation was certainly wrong.

This alternative view would work backwards from the usual practice. Suppose that a careful specification search yields

$$y_t = \beta_0 + \beta_1 y_{t-1} + \beta_2 x_t + \beta_3 x_{t-1} + \omega_t \tag{4}$$

as a form with no serial correlation (among other desirable properties). Comparing (4) term by term with (3) shows that, if $\beta_1 = \rho$, $\beta_2 = \alpha_1$ and $\beta_3 = -\rho\alpha_1$, then (4) could be simplified to (1) supplemented by (2) (with a gain of efficiency, since there is one fewer parameter to estimate). Another way of saying this is that equation (3) is equation (4) subject to the *common-factor restriction*, $\beta_1\beta_2 = -\beta_3$. Such a restriction is imposed every time the Cochrane-Orcutt transformation is performed.

The Cochrane-Orcutt transformation is, therefore, sometimes a 'convenient simplification' of a more general equation (see Hendry and Mizon, 1978). This view of common-factor restrictions is already evident in Sargan's (1964) seminal paper, and is typical of the 'British' or 'LSE approach' to econometrics (see Gilbert 1986a, b). The critical feature of this approach, however, is that a common-factor restriction should be imposed *only* when valid. And validity can be tested using standard χ^2 or F-tests against the alternative hypothesis that an unrestricted equation is correct.

The Cochrane-Orcutt transformation imposes the simplest common-factor restriction. To generalize, consider the model

$$A(L)y_t = B(L)X_t + v_t, \tag{5}$$

where $A(L)$ is a 1×1 polynomial in the lag operator, $L(Lz_t = z_{t-1})$; $B(L)$ is a $1\times k$ polynomial in the lag operator; X_t is a $k\times 1$ vector of independent variables; and v_t is a residual, serially uncorrelated by construction (i.e., through the choice of parameters for $A(L)$ and $B(L)$). Common factors occur whenever $A(L)=0$ and $B(L)=0$ have roots in common: there are as many common factors as there are common roots. By dividing through by the common factors, the number of parameters to be estimated is reduced with a gain to efficiency.

To return to our earlier example, equation (4) may be written as (5) with $A(L)=(1-\beta_1 L)$, $B(L)=[\beta_0, (\beta_0, (\beta_2+\beta_3 L)]$ and $X=[1, x_t]'$. If the common-factor restriction, $\beta_1\beta_2=-\beta_3$, is correct, then $B(L)= [\beta_0, (\beta_2-\beta_1\beta_2 L)] = (1-\beta_1 L)[\beta_0/(1-\beta_1 L), \beta_2]$. Dividing both sides of (5) through by the common factor, $(1-\beta_1 L)$, yields

$$y_t = [\beta_0/(1-\beta_1 L), \beta_2][1, x_2]' + v_t/(1-\beta_1 L) \tag{6}$$
$$= \kappa + \beta_2 x_t + \varepsilon_t,$$

where $\kappa = \beta_0/(1-\beta_1 L) = \text{constant}$ and $\varepsilon_t = v_t/(1-\beta_1 L)$.

Obviously $v_t = (1-\beta_1 L)\,\varepsilon_t = \varepsilon_t - \beta_1\varepsilon_{t-1}$ or

$$\varepsilon_t = \beta_1\varepsilon_{t-1} + v_t. \tag{7}$$

Equations (6) and (7) have precisely the same form as equations (1) and (2):

$$\kappa = \alpha_0,\ \beta_2 = \alpha_1 \text{ and } \beta_1 = \rho.$$

II. THE INVERTED FISHER HYPOTHESIS

Common-factor restrictions, *if they are valid*, provide convenient simplifications of general forms such as (4) and (5). If they are not valid, imposing them through Cochrane-Orcutt transformations may seriously mislead the investigator. To illustrate this, we examine Carmichael and Stebbing's (1983) attempt to test the inverted Fisher hypothesis. Carmichael and Stebbing are not singled out as having committed an error any more egregious than other practitioners; but rather as providing a clear illustration, with a well-defined hypothesis, of how an invalid common-factor restriction can reverse the apparent thrust of empirical evidence.

The Fisher theorem states that, in long-run equilibrium, the market rate of interest should move directly with changes in the rate of inflation. Carmichael and Stebbing (1981, p. 7) argue that, with no uncertainty, optimization requires the real after-tax rate of interest to equal the real rate of return on non-interest-bearing money — that is, minus the rate of inflation. This implies the 'inverted Fisher hypothesis': *the nominal rate of interest is independent of the rate of inflation and the after-tax real rate of interest inversely reflects the rate of inflation.*

Using data for the US three-month Treasury bill and the consumer price index for the period 1953:I-1978:IV, Carmichael and Stebbing specify a static equation of the form[1]

$$(\text{YT}-\text{INF}) = \alpha_0 + \alpha_1\text{INF} + \varepsilon, \qquad (8)$$

where YT is the after-tax rate of return (see Darby, 1975), and INF is the rate of inflation. The inverted Fisher hypothesis implies $\alpha_1 = -1$; while Fisher's own hypothesis implies $\alpha_1 = 0$. Carmichael and Stebbing's estimates of ε in equation (7) are serially correlated. They therefore perform a Cochrane-Orcutt transformation and re-estimate (8) (Carmichael and Stebbing, 1983, Table 1). Their estimate is reproduced with complete summary statistics as regression A.1 of our Table 1:

$$(\text{YT}-\text{INF}) = 1.19 - 1.02\text{INF}, \rho = 0.997. \qquad (\text{A.1})$$

The coefficient on inflation is nearly negative one, which seems to support the inverted Fisher hypothesis. And the autocorrelation coefficient is nearly one, which suggests to Carmichael and Stebbing that A.1 could be re-estimated in first differences (reproduced as A.2, Table 1):

$$\Delta(\text{YT}-\text{INF}) = -1.02\Delta\text{INF}. \qquad (\text{A.2})$$

again apparently confirming the inverted Fisher hypothesis.

Equation A.1 is nested in

$$(\text{YT}-\text{INF}) = \gamma_0 + \gamma_1\text{INF} + \gamma_2\text{INF}_{-1} + \gamma_3(\text{YT}-\text{INF})_{-1} + \omega. \qquad (9)$$

Equation (8) is (9) subject to the common-factor restriction that $\gamma_1\gamma_3 = -\gamma_2$. The column labelled 'COMFAC' in Table 1 tests this restriction. It is decisively rejected at every conventional level of significance.

[1] Carmichael and Stebbing's (1983) original data (1953:I-1978:IV) is used throughout this paper. Consult that paper for sources and definitions.

Because $\rho \simeq 1$, A.1 is essentially a random walk, and the variance of (YT−INF) is not well-defined. The significance levels of tests based on A.1 are thus doubtful. Nevertheless, conditional on A.1 the COMFAC test is as valid as any that Carmichael and Stebbing perform.

To avoid this problem, we may concentrate our attention of Δ(YT−INF) as in A.2. This implies (as of course does A.1 with $\rho = 1$) that there is no long-run relationship between the levels of YT and INF. Even so A.2 is also nested in (8): $\gamma_0 = 0$, $\gamma_1 = \gamma_2$ and $\gamma_3 = 1$. The COMFAC test decisively rejects these restrictions as well. Carmichael and Stebbing's equations are clearly misspecified.

III. RESPECIFICATION OF INTEREST RATE DYNAMICS

How important is this misspecification? We answer this question by trying to specify a better equation using a general-to-specific modelling strategy (see Hendry and Richard 1982; Hendry 1983 and 1986; and Gilbert 1986a). We restrict ourselves to the time period and the variables used in Carmichael and Stebbing's paper.

The specification search begins with a vector autoregression [VAR(4)] of YT on a constant, four lags of itself and the current value and four lags of INF and TAX (the marginal tax rate used to construct YT), estimated for the period 1954:I-1975:IV (twelve quarters of data are reserved for a Chow test of out-of-sample parameter constancy — a test both for stability and to prevent illicit data-mining (Harvey, 1981, pp. 181–182). Owing to multicollinearity, VAR (4)'s coefficients are virtually uninterpretable but its $R^2 = 0.91$, SER = 0.44 and Sum of Squared Residuals = 13.36. A Chow test for out-of-sample parameter constancy $F(12, 70) = 0.80$ shows that the hypothesis of structural stability cannot be rejected at any conventional significance level.[2]

The most parsimonious representation of VAR(4) uncovered through the specification search is reproduced from Table 1:

$$\underset{\substack{(0.66)\\ [0.59]}}{\Delta YT = 2.46} \; \underset{\substack{(0.08)\\ [0.10]}}{-\,0.37\, YT_{-1}} \; \underset{\substack{(0.03)\\ [0.04]}}{+\,0.13\, INF_{-1}} \; \underset{\substack{(0.03)\\ [0.03]}}{+\,0.05\Delta INF} \qquad \text{(B.1)}$$

$$\underset{\substack{(1.73)\\ [1.42]}}{-\,5.14\, TAX_{-1}} + \text{two seasonal dummies}$$

$R^2 = 0.29$; SER = 0.42; AR(4): $F(4, 77) = 0.69$;
Normality: $\chi^2 = 5.11$; ARCH: $F(4, 77) = 1.93$
CHOW: $F(12, 81) = 1.09$

This is a 'levels-and-differences' form, which allows the relationship to have a well-defined long-run solution, while at the same time adequately

[2] For the whole sample: $R^2 = 0.91$, SER = 0.43, RSS = 15.18.

TABLE 1

Estimates of the Effect of CPI Inflation on the US 3-month T-bill Rate.

Equation	*Dependent variable (mean) [standard deviation]*	*Independent variables (standard errors) [Heteroskedasticity-corrected standard errors]* Constant	YT_{-1}	$(YT\text{-}INF)_{-1}$	INF	INF_{-1}	ΔINF	TAX_{-1}	S_1§	S_2§
A.1 1953:II–1978:IV	(YT-INF) (−0.73) [2.29]	1.19 (*)			−1.02 (0.02)					
A.2 1953:II–1978:IV	Δ(YT-INF) (−0.52) [2.27]						−1.02 [0.02]			
B.1 1954:I–1975:IV	ΔYT (0.041) [0.48]	2.46 (0.66) [0.59]	−0.37 (0.08) [0.10]			0.13 (0.03) [0.04]	0.05 (0.03) [0.03]	−5.14 (1.73) [1.42]		−0.38 (0.11) [0.12]
B.2 1954:I–1978:IV	ΔYT (0.046) [0.48]	2.35 (0.63) [0.57]	−0.37 (0.07) [0.09]			0.13 (0.03) [0.03]	0.05 (0.03) [0.03]	−4.88 (1.66) [1.37]		−0.37 (0.11) [0.11]
B.3 1954:I–1978:IV	ΔYT (0.046) [0.48]	2.33 (0.63) [0.57]	−0.35 (0.07) [0.08]			0.12 (0.03) [0.03]	0.05 (0.03) [0.03]	−4.64 (1.66) [1.40]	−0.15 (0.12) [0.12]	−0.45 (0.12) [0.14]
C.1 1954:I–1978:IV	Δ(YT-INF) (−0.04) [2.22]						−1.03 (0.02) [0.02]			
C.2 1954:I-1978:IV	ΔYT (0.046) [0.48]	0.30 (0.48) [0.33]		−0.03 (0.02) [0.02]				−0.21 (1.45) [0.99]	−0.22 (0.13) [0.14]	−0.43 (0.13) [0.14]

*Summary statistics***

Equation	S_3§	ρ	R^2	SER	RSS	AR(4) $F(,)$	*Normal* $x^2(2)$	ARCH $F(,)$	Chow‡ $F(,)$	COMFAC $F(,)$	*Nested in Eq. 3* $F(,)$	*Nested in* VAR(4)† $F(,)$
A.1 1953:II–1978:IV		0.997 (0.004)	0.96	0.47	23.06	1.67 (4,98)	9.74	1.12 (4,98)	1.03 (12,90)	11.76 (1,99)		
A.2 1953:II–1978:IV			0.96	0.48	23.12	1.66 (4,98)	10.22	1.12 (4,98)	1.03 (12,90)	3.93 (3,99)		
B.1 1954:I–1975:IV	−0.23 (0.11) [0.11]		0.29	0.42	14.42	0.69 (4,77)	5.11	1.93 (4,77)	1.09 (12,81)			0.51 (11,70)
B.2 1954:I–1978:IV	−0.21 (0.11) [0.10]		0.27	0.42	16.75	0.88 (4,89)	5.00	1.29 (4,89)			1.53 (1,92)	0.77 (11,82)
B.3 1954:I–1978:IV	−0.27 (0.12) [0.12]		0.27	0.42	16.47	0.84 (4,88)	4.84	1.30 (4,88)	0.96 (12,80)			0.70 (10,82)
C.1 1954:I–1978:IV			0.95	0.48	22.98	1.64 (4,95)	10.17	1.08 (4,95)	1.00 (12,87)		5.19 (5,92)	2.48 (17,82)
C.2 1954:I–1978:IV	−0.19 (0.13) [0.13]		0.13	0.46	20.07	0.81 (4,90)	9.77	1.22 (4,90)	0.89 (12,82)		10.04 (2,92)	2.20 (12,82)

* Constant constrained to take Carmichael and Stebbing's estimated value, since their results could not be reproduced on *Micro TSP*'s Cochrane-Orcutt routine when freely estimated.

** Descriptions and references in text.

§ Seasonal dummies: $S_1 = 1$ in first quarter; $S_2 = 1$ in second quarter; $S_3 = 1$ in third quarter; each $= 0$ otherwise.

‡ Test of 12 quarter out-of-sample estimates against a baseline regression estimated from beginning of indicated sample to 1975.IV.

† VAR(4) is fitted over the same period as the equation reported.

Equations 1A, estimated by ordinary least squares with AR1 (Cochrane-Orcutt) correction using *Micro TSP, Version 4.0*; the remainder estimated by ordinary least squares using PC-GIVE, Version 4.2.

capturing short-run behaviour (see Harvey, 1981, pp. 290–92). Neither the null hypothesis of no serial correlation of up to fourth order (AR(4)) (Harvey, 1981, pp. 276–77) nor the null hypothesis of no autoregressive conditional heteroskedasticity (ARCH) (Engle, 1982) of up to fourth order can be rejected at the 95 per cent level. The hypothesis of normal residuals (Jarque and Berra, 1980) cannot be rejected at the 95 per cent level, which suggests that the F-test are appropriate. The close correspondence of the usual form of the standard errors (parentheses beneath the estimated coefficients) with the heteroskedasticity-corrected forms (square brackets) shows that heteroskedasticity is not a problem (White, 1980). The Chow test of out-of-sample parameter constancy cannot reject the null hypothesis of stable parameters and lends some support to belief that desirable properties of this equation are not artefacts of its construction but correspond to the underlying reality.[3]

Equation B.2 reports the same specification estimated over the entire set of available data. Its properties are similar to B.1 in every respect:

$$\begin{aligned}\Delta YT = \underset{\substack{(0.63)\\ [0.57]}}{2.35} - \underset{\substack{(0.07)\\ [0.09]}}{0.37}\, YT_{-1} + \underset{\substack{(0.03)\\ [0.03]}}{0.13}\, INF_{-1} + \underset{\substack{(0.03)\\ [0.03]}}{0.05}\Delta INF \\ - \underset{\substack{(1.66)\\ [1.37]}}{4.88}\, TAX_{-1} + \text{two seasonal dummies}\end{aligned} \qquad (B.2)$$

$R^2 = 0.27$; SER $= 0.42$; AR(4): $F(4, 89) = 0.88$;
Normality: $\chi^2 = 5.00$; ARCH: $F(4, 89) = 1.29$

Eleven restrictions must be placed on VAR(4) in order to derive B.1 or B.2. The test of the null hypothesis that these restrictions are valid (reported in Table 1 as 'Nexted in VAR(4)') cannot be rejected at any conventional level of significance.

While it is possible that a more parsimonious model than B.2 is a valid restriction of VAR(4), it is clear that we cannot reject the hypothesis that B.2 encompasses its obvious rivals (see Hendry and Richard, 1982, pp. 16–20). Equation C.1 has the same form as A.2, but is estimated over the same period as VAR(4) and B.2 (1954:IV-1978:IV). Equation C.2 is the best equation found with an error-correction term for inflation, which ensures that the Fisher hypothesis holds in the long run (Salmon, 1982 and Harvey, 1981, pp. 290–92). Both equations are rejected against VAR(4) at the 99 per cent confidence level. In contrast, equation B.2 cannot be rejected at even the 95 per cent level against VAR(4). Furthermore, B.2, C.1 and C.2 are all nested in B.3, which is B.2 with the addition of the third, insignificant seasonal dummy.

[3] While its R^2 is very low compared to VAR(4) or A.1 and A.2, this is simply an artefact of specifying the dependent variable as a first difference of the nominal rate. Judged more accurately by the standard error of regression, equations B.1-3 are the best fitting regressions reported in Table 1. R^2 for B.1, recomputed in terms of levels, is 0.91.

Equation B.2 cannot be rejected as a valid restriction of B.3; while C.1 and C.2 can be rejected at the 99 per cent level (see the column 'Nested in Equation B.3' in Table 1). We are justified, then, in tentatively accepting B.2 as the best parsimonious representation of VAR(4).

What difference does this result make to the inverted Fisher hypothesis? Obviously, it does not hold in the short run as Carmichael and Stebbing take their results to imply. Fisher took his own theorem as a theoretical conclusion about long-run equilibrium. Let us then calculate the static long-run solution to B.2:

$$\overline{YT} = 6.35 + 0.35\,\overline{INF} - 13.19\,\overline{TAX} \tag{10}$$

Equation (10) indicates that while the inverted Fisher hypothesis as stated by Carmichael, and Stebbing is not supported by the data, neither is Fisher's theoretical prediction that nominal interest rates reflect inflation one for one in the long run. Only 35 per cent of an increase in inflation seems to be reflected in the long run in rate of interest. With a mean lag of 2.32 quarters in B.2, the effect of inflation on interest rates is also not immediate. While these results may bring little comfort to supporters of extreme forms of either hypothesis, they are qualitatively the same *empirical* results reported by Fisher (1930, p. 451) himself.

IV. SHOULD WE STOP HERE?

Carmichael and Stebbing's claim that the evidence supports a radically different understanding of the relation of interest rates to inflation than has been widely accepted in the profession is unsupported. They are misled, specifically, by the habit of applying the Cochrane-Orcutt transformation without testing the implicit common-factor restriction and, generally, by inattention to dynamic specification. We are entitled to ask, however, if we can rest easy with the specification of equation B.2. Unfortunately, the answer is, no.

One symptom of a misspecified regression is the instability of the coefficients out-of-sample. Carmichael and Stebbing tested the within-sample stability of A.1 and A.2 with a Chow test. Table 1 shows that another Chow test cannot reject the hypothesis of stability of these regressions over twelve quarters beyond a baseline regression running from 1953.I to 1975.IV. But these are minimal tests at best. Figure 1 presents a more searching test of C.1, which has the form of A.2 but which, like B.1-3, is estimated over 1954.I-1978.IV. Figure 1 graphs the sequence of Chow test statistics scaled by their 5 per cent significance levels, so that unity represents rejection at the 95 per cent level. Successive points on the graph are the Chow tests for parameter constancy of a baseline regression (1954.I-1958.II) against longer and longer samples (1954.I to t, where t = 1958.III, 1958.IV, ..., 1978.IV).[4]

[4] The statistic used is Chow's (1960) first test, see pp. 594, 595 and 598.

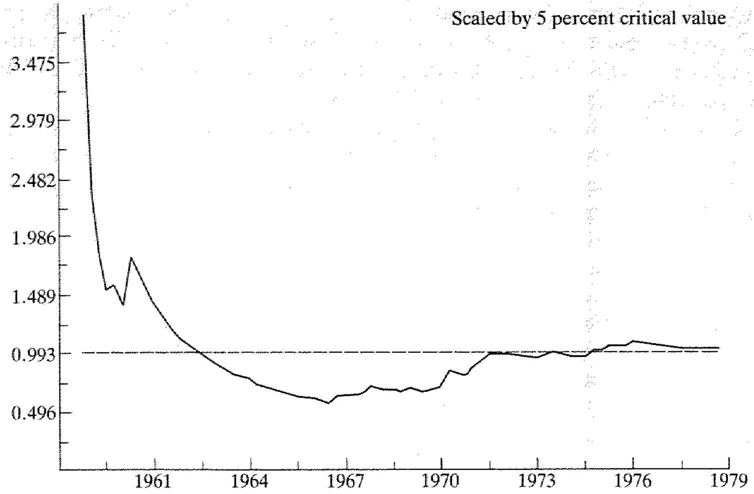

Fig. 1.

The graph clearly shows rejection of stability in the early and later years of the sample period. Further evidence against Carmichael and Stebbing's specification.

Figure 2 presents the same sequential Chow tests for our preferred equation B.2. Stability cannot be rejected at any point in the sample period, lending further support to B.2.

The sequential Chow test with a constant base is not the only possible one. One alternative is a one-step-ahead Chow test in which the base advances one period for each test in the sequence (i.e., the sequence of regressions 1954.I to t, where t=1958.II, 1958.III, ..., 1978.III, is tested against the regression 1954.I to $t+1$). The results of this test for C.1 are reported in Figure 3. Stability is clearly rejected in the earliest observations and again for much of the later period. Figure 4 reports the same test for B.2. Although stability is not rejected in the early observations, it is frequently rejected for much of the later period. The rejection frequency of this test should be around 4 (5 per cent of 80).

A third alternative Chow test, tests a sequence of base periods from 1954.I-1958.II to the end of the sample against the entire sample period (i.e., the sequence of regressions 1954.I to t, where t=1958.II, 1958.III, ..., 1978.III, is tested against the regression 1954.I to 1978.IV). Figure 5 reports this test for C.1. Once again stability is rejected in the earliest observations. It is also rejected in the middle of the sample, although not at the end. Figure 6 reports the same test for B.2. Again, stability cannot be rejected in the early observations; but it is clearly rejected in the middle of the sample.

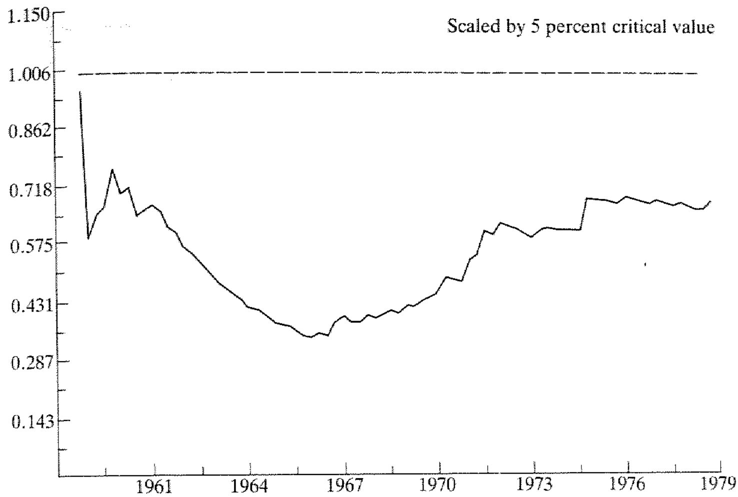
1.150
1.006
0.862
0.718
0.575
0.431
0.287
0.143
Scaled by 5 percent critical value
1961
1964
1967
1970
1973
1976
1979

Fig. 2

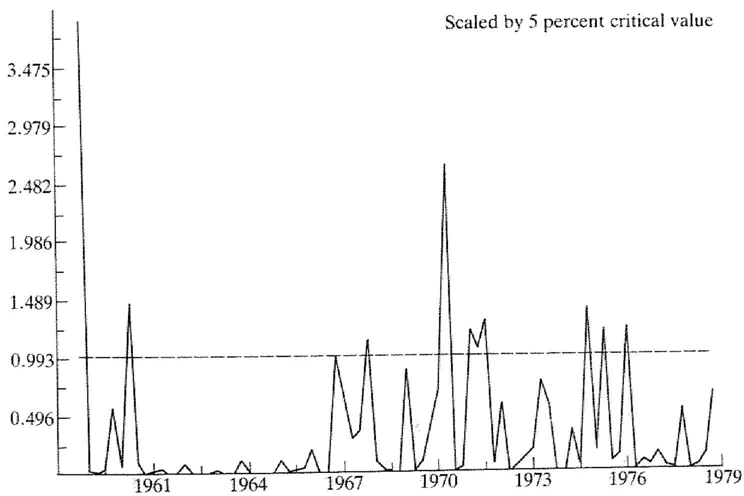
3.475
2.979
2.482
1.986
1.489
0.993
0.496
Scaled by 5 percent critical value
1961
1964
1967
1970
1973
1976
1979

Fig. 3

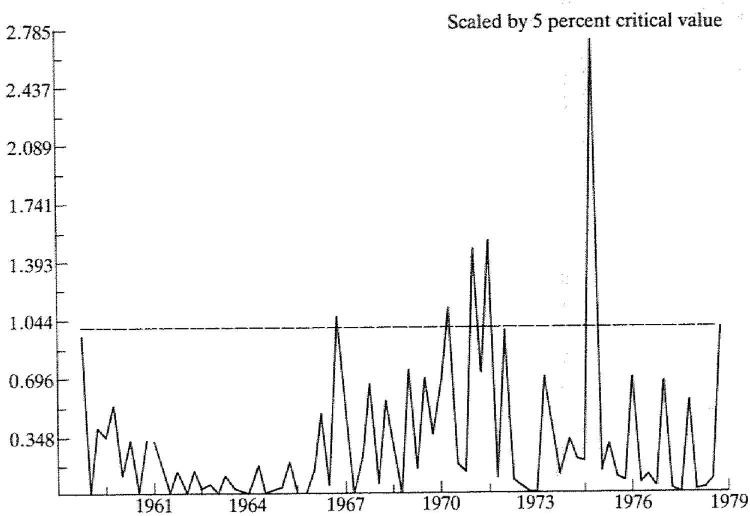
Scaled by 5 percent critical value
2.785
2.437
2.089
1.741
1.393
1.044
0.696
0.348
1961
1964
1967
1970
1973
1976
1979

Fig. 4

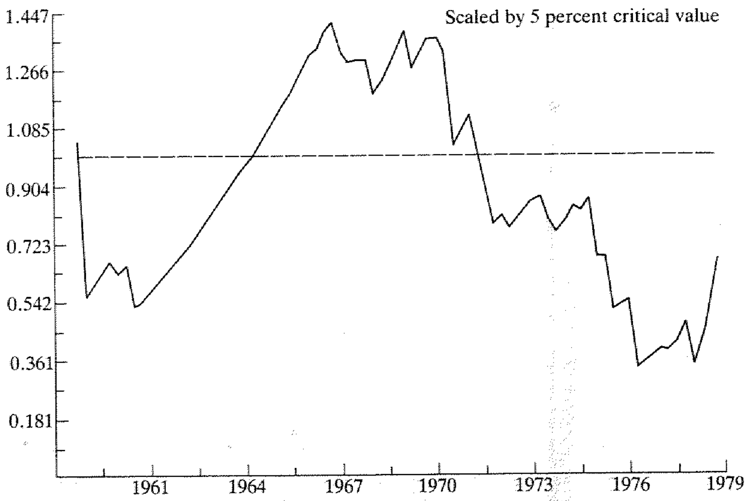
Scaled by 5 percent critical value
1.447
1.266
1.085
0.904
0.723
0.542
0.361
0.181
1961
1964
1967
1970
1973
1976
1979

Fig. 5

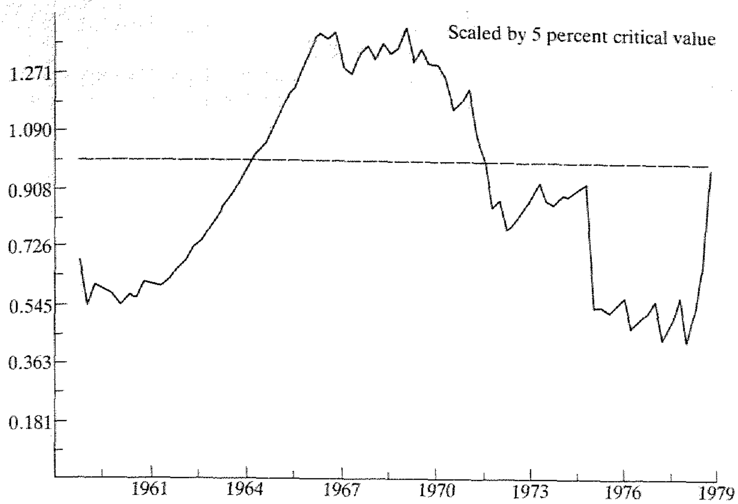

Fig. 6

These more rigorous tests of stability show that B.2 performs better, although not decisively better, than C.1. This suggests that it too is misspecified. Given the nature of the specification search, it seems unlikely that a greatly superior specification is to be found using only Carmichael and Stebbing's data. The evidence of misspecification suggests that proper specification will probably be had only if our nets are cast wider and other factors besides inflation and marginal tax rates are considered as determinants of interest rates.

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