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Co-integration, error correction and the Fisher effect

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I. INTRODUCTION

A question of central concern in macroeconomics is that of the role played by interest rates. If some representative nominal interest rate behaves in a classical manner, a change in the nominal interest rate simply accounts for a change in inflationary expectations in equilibrium. If this is true, then the real interest rate is determined by real factors and cannot be influenced by monetary policy. This is one of the reasons why the Fisher effect has attracted so much empirical attention. In this paper we present evidence from both the US and Australian economies that post-tax nominal interest rates and inflation are co-integrated in the sense of Granger and Engle (1987), and that these variables have a joint error correction representation. This allows the interpretation that movements in nominal interest rates and inflation have long-run components which are consistent with the quantity theory of money, and short-run components which are consistent with adjustment to equilibrium.

II. NOMINAL AND REAL INTEREST RATES AND INFLATION

The Fisher effect is a comparative statics result which states that in long-run equilibrium a fully perceived change in the rate of growth of the nominal money stock leads to a fully perceived change in inflation and a concomitant adjustment of nominal interest rates. Equilibrium here presumably means that all agents possess all information that can profitably be learned from market interaction. Therefore, as pointed out by Sargent (1976), the manner in which inflationary expectations are formed has important implications for the Fisher effect. In long-run equilibrium inflationary expectations equal actual inflation and thus rational expectations and the Fisher effect form what Sargent calls 'a natural package'. In this situation it would be proper, from an econometric perspective, to regress nominal interest rates on actual inflation as the least squares orthogonality requirement would be met. However, there is a simultaneity problem here that is more than a simple violation of this technical requirement. Unless we possess data which pertains to long-run equilibrium positions of the economy, observed movements in interest rates and inflation are quite possibly a mixture of dynamic adjustment and equilibrium co-movement. This is why the simultaneity problem may be more than a simple violation of the least squares orthogonality requirement. It has been amply documented (Sargent, 1973; Summers, 1983) that much early empirical work on the Fisher effect did not recognize this problem.

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In spite of the above, a great deal of the current empirical testing of the Fisher effect has proceeded as if a single static equation is capable of uncovering the equilibrium relationship. The question of endogeneity is dealt with by applying some restrictions to the relationship so that nominal interest rates can be considered to be exogenous, as in Fama (1975), or actual inflation can be considered to be exogenous, as in Carmichael and Stebbing (1983) (CS). However, even imposing those restrictions, so that the equation is rendered econometrically correct under some null hypothesis, will not deal with the differing short-run/long-run effects on the two variables. Thus, the existence of low Durbin–Watson statistics (e.g CS) may be pointing to a problem of unit roots and spurious regression in the sense of Granger and Newbold (1977). To see how this may be important, consider the following development of single-equation tests of the Fisher effect.

Ignoring for the moment the role of taxes, the starting point is the famous Fisher identity:

$$r_t^e \equiv R_t - \pi_t^e \tag{1}$$

where r_t^e is the *ex ante* real rate of interest, R_t is the nominal rate of interest, and π_t^e is the expected inflation rate. In order to proceed to empirical testing some theory of behaviour must be postulated which will introduce stochastic variation into relation 1. Assume that inflationary expectations are unbiased:

$$\pi_t = \pi_t^e + \varepsilon_t, \tag{2}$$

where ε_t is assumed to be some constant variance, mean zero noise. Combining 1 and 2 yields

$$r_t^e = R_t - \pi_t - \varepsilon_t. \tag{3}$$

Fama's estimating equation is a rearrangement of Equation 3:

$$\pi_t = \alpha + BR_t + \varepsilon_t. \tag{4}$$

Under the joint null hypothesis of rational expectations and a constant real rate of interest, $\alpha = -\bar{r}$ and $B = 1.0$. The simultaneity problem is avoided here as under the null hypothesis all available information concerning inflationary expectations is used in the setting of R and, therefore,

$$E(R_t, \varepsilon_t) = 0. \tag{5}$$

The development of this estimating equation can be compared to the CS approach. CS begin with Equations 1 and 2. Their null hypothesis is what they term the inverted Fisher hypothesis (IFH). Under this null hypothesis the nominal interest rate is constant due to substitutability assumptions regarding money and bonds. This is introduced through the following equation:

$$r_t^e = \alpha - \pi_t^e + v_t. \tag{6}$$

This would appear to make Equation 1 a stochastic identity, with $\alpha = R_t$. Combining Equations 6 and 2 leads to the CS estimating equation

$$R_t = \alpha + B\pi_t + \eta_t \tag{7}$$

where η is a composite error which, under the null hypothesis of IFH, is orthogonal to π_t.

It is well known that the published empirical results based on the above derivations are contradictory. Part of the reason is possibly due to the fact that Fama (1975) uses pre-tax nominal interest rates, while CS use post-tax nominal interest rates. If pre-tax real interest

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rates are constant, as in Fama's results, then post-tax real interest rates will move one for one (inversely) with inflation as in the CS results. However, as demonstrated in Table 1, this does not account for the inconsistent results. What does account for this anomaly is CS use of a first-difference specification.

Table 1. *Bivariate regressions*

				$\bar{R}^2$	DW	$Q(24)$
A. *Level form*						
1.1	π_t	$= -0.0021 + 0.883$	R_t	0.464	2.02	28.27
		(0.0010) (0.108)				(0.24)
1.2	R_t	$= 0.0057 + 0.533$	π_t	0.464	1.28	95.53
		(0.0050) (0.065)				(0.00)
1.3	π_t	$= -0.0016 + 1.20$	RN_t	0.466	2.00	26.99
		(0.0009) (0.147)				(0.30)
1.4	RN_t	$= 0.0038 + 0.393$	π_t	0.466	1.23	116.35
		(0.0003) (0.048)				(0.00)
B. *First-difference form*						
1.5	$\Delta\pi_t$	$= 0.0002 - 0.433$	ΔR_t	0.007	2.86	42.79
		(0.0005) (0.346)				(0.00)
1.6	ΔR_t	$= 0.00009 - 0.048$	$\Delta\pi_t$	0.007	1.98	21.26
		(0.0001) (0.038)				(0.62)
1.7	$\Delta\pi_t$	$= 0.0002 - 0.635$	ΔRN_t	0.007	2.86	42.72
		(0.0005) (0.507)				(0.01)
1.8	ΔRN_t	$= 0.00007 - 0.033$	$\Delta\pi_t$	0.007	1.99	21.61
		(0.0001) (0.026)				(0.60)

Note: π_t is the inflation rate, R_t is the pre-tax nominal interest rate, RN_t is the post-tax nominal interest rate, $\bar{R}^2$ is the adjusted R^2, DW is the Durbin–Watson statistic, and Q is the Box–Pierce Q statistic. Standard errors are given in parentheses except for the Q statistic, where the marginal significance level for rejection of the null hypothesis of white noise residuals is given.

In Table 1 we present a series of bivariate ordinary least squares regressions. The data is US 90 day Treasury Bill rate, a marginal tax rate on interest income and the rate of change in the US CPI, quarterly from 1953.Q1 to 1971.Q4 to conform with the Fama sample period.[1] Equation 1.1 on Table 1 represents a Fama test (Equation 4 in the text). The coefficient on the pre-tax nominal interest rate is not significantly different from 1.0, the DW is very close to 2.0 and Q statistic cannot reject white-noise residuals. Equation 1.2 is the reverse regression of this, which is a CS equation in level form with a pre-tax nominal interest rate. The coefficient on inflation is more than six standard deviations from 1.0, but it is significantly different from 0.0. Notice that the DW is low and the Q resoundingly rejects white-noise residuals. Equations 1.3 and 1.4 reproduce 1.1 and 1.2 using post-tax nominal

[1] I am grateful to Jeff Carmichael for providing me with the data that was used in the CS article. All estimation was undertaken using RATS Version 4.1

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interest rates. Notice that this does not change whether or not we reject the null in each case. Although there is some change in the estimated coefficients, the inclusion of taxes is not making a qualitative difference. If this were the extent of the empirical evidence available, one would be tempted to reject the IFH in favour of the Fisher effect. However, in Equations 1.2 and 1.4 the DW is quite low and the Q quite large and, accordingly, drawing valid statistical inferences is questionable. It is possible that these regressions are candidates for the class of spurious regressions of Granger and Newbold.

Equations 1.5–8 in Table 1 reproduce 1.1–4 with all variables now expressed in first-difference form. Equation 1.8 is the CS IFH result for the period 1953–71. In terms of the null hypothesis of the Fisher effect there is very little to choose between all of the equations, no right-hand side variables are ever significantly different from 0.0. The first-difference specification appears to resoundly reject the Fisher effect. Once again the inclusion of taxes does not make a great deal of difference to the results. Clearly it is the first-difference specification which is driving the CS results. This points to the possibility that nominal interest rates and inflation contain unit roots. This is pursued in the next section.

III. INTEGRATION AND CO-INTEGRATION TESTS

The co-integration and error correction systems outlined in Granger and Engle (1987) allow long-run components of variables to obey equilibrium constraints while short-run components are allowed to have flexible dynamic specifications.[2] In this framework if nominal interest rates and inflation can each be shown to be stationary after first-differencing, denoted integrated of order one or $I(1)$, then equations of the sort given by 1.1–4 can be viewed as co-integrating regressions, rather than as an actual test of the long-run Fisher effect. If we can reject the null hypothesis of no co-integration, then equations of the sort given by 1.5–8 can be viewed as misspecified in the sense of missing (at least) the co-integration factor. In the next two subsections, we use data from the US and Australian economies to test the post-tax nominal interest rate and inflation rate for unit roots and co-integration. In order to be as general as possible, we consider the variables to be jointly endogenous.

Univariate time series of properties

Before proceeding to test for co-integration it is necessary to establish the time series properties of the individual series. If two series are integrated of different orders they cannot be co-integrated. In this paper we apply the augmented Dickey–Fuller test (ADF) as recommended by Granger and Engle (1987). This test is the t statistic on ρ in the following regression $[\Delta \equiv (1-L)]$:

$$\Delta Z_t = \rho Z_{t-1} + \sum_{i=0}^{K} \beta_i \, \Delta Z_{t-i} + \varepsilon_t \qquad (8)$$

where Z_t is the series under consideration and K is large enough to ensure that ε_t is white noise. In practice we do not know the appropriate order of the autoregression, K. In our implementation of the ADF test we follow the suggestion of Engle and Yoo (1987) and use

[2] An excellent overview is found in Hendry (1986). See also other papers in the same issue.

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the Akaike (1974) information criterion (AIC) to determine the optimal specification of Equation 8. This criterion is defined by

$$AIC(q)=N\ln(SSR/N)+2q \tag{9}$$

where N is the number of observations to which the model is fitted, SSR is the sum of squared residuals and q is the number of parameters, equal to $K+1$. By this method the appropriate order of the model is determined by computing Equation 8 over a selected grid of values of K and finding that value of K at which the AIC attains its minimum. The distribution of the ADF statistic is non-standard and, accordingly, we use the critical values tabulated by Engle and Yoo.

The results of this procedure are presented on Table 2. The upper half of the table tests the null hypothesis that the variables are $I(1)$. None of the variables are able to reject this null and the US nominal interest rate is even the wrong sign to be level stationary. Accordingly,

Table 2. *Unit root tests*

$H_0: I(1)$

$$\Delta Z_t=\rho Z_{t-1}+\sum_{i=0}^{K}\beta_i\Delta Z_{t-i}+\varepsilon_t$$

		ρ	t_ρ	$AIC(K)$
USA				
	RN	0.002	0.117	−1385.23(0)
	π	−0.044	−1.06	−1085.95(1)
Australia				
	RN	−0.002	−0.079	−750.51(0)
	π	−0.040	−0.706	−595.53(1)

$H_0: I(2)$

$$\Delta^2 Z_t=\rho\Delta Z_{t-1}+\sum_{i=0}^{K}\beta_i\Delta^2 Z_{t-i}+\varepsilon_t$$

		ρ	t_ρ	$AIC(K)$
USA				
	RN	−0.99	− 9.90[a]	−1370.79(0)
	π	−1.40	−15.93[a]	−1086.80(0)
Australia				
	RN	−1.18	−10.40[a]	−750.55(0)
	π	−1.53	−14.20[a]	−597.01(0)

[a]Denotes rejection at 1% level.

Notes: 1 The sample period is limited by availability of the marginal tax rate in each country. In levels the data period is 53, 1–78, 4 for the USA and 65, 2–81, 4 for Australia.

2 The assymptotic critical values at a 1% significance level are:
(a) for 100 observations (US) −3.73, −4.22.
(b) for 50 observations (AUS) −4.12, −4.45 for 1 and 2 right-hand side variables respectively. See Engle and Yoo (1986, Table 3).

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all variables were tested for second-order integration, the results of which are on the bottom half of Table 2. All t statistics are greater than the 1% critical values.[3] This is taken as evidence that for both the US and Australian economies, the post-tax nominal interest rate and inflation are first-difference stationary.

Co-integration tests

The tests for co-integration involve fitting the long-run relation in level form with a simple OLS regression and performing diagnostic tests on the residuals. Two formal tests of the null hypothesis of no co-integration are presented; these are the co-integrating regression Durbin–Watson (CRDW) used by Granger and Engle (1987, derived from Sargan and Bhargava, 1983) and the ADF based on the co-integrating regression residuals, with lag length again chosen by the AIC. The results of these procedures are presented on Table 3. There it can be seen that for both countries each test rejects the null hypotheses of no co-integration regardless of which variable is chosen as the dependent variable.[4] It is interesting to note that if we were to consider each of the four co-integrating regressions as static tests of the Fisher effect, only the Australian data, with inflation as the dependent variable cannot reject the null hypothesis of one-for-one adjustment.[5]

Table 3. *Co-integration tests*

A. Co-integrating regression: $RN_t = \alpha + \beta\pi_t + \varepsilon_t$				
	$\hat{\beta}$	$t_{\hat{\beta}}$	$CRDW$	$ADF(K)$
USA	0.326	12.48	1.13	−6.28(0)
Australia	0.225	5.37	0.94	−4.80(0)

B. Co-integrating regression: $\pi_t = \alpha^1 + \beta^1 RN_t + \varepsilon_t^1$				
	$\hat{\beta}^1$	$t_{\hat{\beta}^1}$	$CRDW$	$ADF(K)$
USA	1.85	12.48	1.47	−7.84(0)
Australia	1.36	5.37	1.38	−5.84(0)

Notes: 1. See Table 2, Note 1 for sample periods.
2. The ADF statistic is calculated as per Table 2, where Z_t is the residual from the co-integrating regression in each case.
3. The critical values for $CRDW$ for $N = 100$ at 1%, 5% and 10% are 0.511, 0.386 and 0.322, respectively. Note that for Australia $N = 68$ and therefore some interpretation is required. See Granger and Engle (1987, Table 2).

[3]As an additional test the ADF with four lags of the dependent variable was calculated, even though the AIC did not choose this lag length. In each case the ability of the model to reject the null hypothesis of $I(2)$ was reduced. No test, however, failed to reject above the 10% level.
[4]Once again the ADF test with four lags was implemented. For the USA this test failed to reject at the 10% level for both regressions and for Australia this test rejected at the 10% level for both regressions.
[5]This is subject to the same warning concerning valid statistical inferences with low DW statistics mentioned above.

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IV. ERROR CORRECTION MODELS

The above section presented some evidence that for both the US and Australian economies, post-tax nominal interest rates and inflation are co-integrated. The implication of this is that there should exist an error correction representation between these variables in each country. For the two-variable systems under consideration in this paper a typical error-correction model would relate the change in one variable to past equilibrium errors as well as past changes in both variables. The important implication of this is that previous vector autoregression representations of the dynamics involved in inflation and interest rates may have missed a potentially important variable if they have left out adjustment to past equilibrium errors.

One of the most recent estimates of a first-difference vector autoregression for inflation and the post-tax nominal interest rate using US data can be found in Gallagher (1986). Since the data set used in this paper is identical to that used by Gallagher, an initial specification for an error-correction model was chosen based on his results. This entails an inflation equation with twelve lags of inflation and four lags of the interest rate and an interest-rate equation with four lags of each variable. One lag of the error-correction term is added to each equation. The results of this procedure for inflation and the interest rate are found on columns 1 and 4 of Table 4, respectively. Comparison of these results with those in Gallagher (p. 248, Table 1) shows that adding the error-correction term to the inflation equation does not alter the equation in any substantial manner. Notice that the error-correction term is not statistically significant. In the interest-rate equation the error-correction term is significant and its inclusion alters the significance of the inflation lags.

Given that the error-correction term is now included in the dynamic representation, it is no longer clear what the appropriate lag lengths are. In order to investigate the dynamics further, we follow the procedure suggested in Granger and Engle (1987) of estimating the simplest error-correction model and testing for added lags. The AIC is used as a guide in this procedure. The preferred specification for inflation in column 2 has three lags of each variable and one lag of the error-correction term, which is now significant at the 6% level. Column 3 on Table 4 is an inflation specification with only one lagged error-correction term. This model is somewhat inferior to that in column 2 based on the standard error and the AIC. However, it is interesting that this most simple of dynamic specifications clearly is an improvement over Equation 1.7. For the interest rate there is very little to choose between a specification with an error-correction term and three lags of each variable and a specification with just the error-correction term.

An important empirical result to come out of inflation and interest-rate vector autoregressions using US data is that there is feedback or 'Granger-causality' in both directions (see, e.g. Sargent 1973 or Gallagher 1986). Since addition of the error-correction term has altered the estimated lag lengths, these conclusions may be altered. Consider the null hypothesis that the interest rate does not 'Granger-cause' the inflation rate. This could be tested by the hypothesis that all of the lags on the interest rate in column 2 on Table 4 are zero. A test of this constraint is provided by the statistic

$$L = N(RSSC - RSSU)/RSSU \tag{10}$$

which converges in distribution to a χ^2_c variate (see Geweke *et al.* 1983). $RSSC$ and $RSSU$ are the constrained and unconstrained sum of squared residuals, respectively, and c is the number of constraints. For the test under consideration $L = 14.41$ which is greater than

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Table 4. *Error correction models, US data*

	Dependent variable:					
	$\Delta\pi$	$\Delta\pi$	$\Delta\pi$	ΔRN	ΔRN	ΔRN
EC_{t-1}	−0.160	−0.252	−0.590	−0.166	−0.171	−0.201
	(−1.15)	(−1.94)	(−6.56)	(−2.18)	(−2.30)	(−3.56)
$\Delta\pi_t$				0.031		
				(1.15)		
$\Delta\pi_{t-1}$	−0.705	−0.493		0.059	0.027	
	(−4.16)	(−3.53)		(1.51)	(0.84)	
$\Delta\pi_{t-2}$	−0.540	−0.333		0.089	0.056	
	(−2.96)	(−2.60)		(2.33)	(1.81)	
$\Delta\pi_{t-3}$	−0.321	−0.167		0.094	0.059	
	(−1.80)	(−1.66)		(2.71)	(2.32)	
$\Delta\pi_{t-4}$	−0.078			0.042		
	(−0.50)			(1.55)		
$\Delta\pi_{t-5}$	0.121					
	(0.89)					
$\Delta\pi_{t-6}$	0.0529					
	(0.427)					
$\Delta\pi_{t-7}$	0.046					
	(0.39)					
$\Delta\pi_{t-8}$	−0.333					
	(−2.83)					
$\Delta\pi_{t-9}$	−0.464					
	(−3.76)					
$\Delta\pi_{t-10}$	−0.478					
	(−3.58)					
$\Delta\pi_{t-11}$	−0.104					
	(−0.74)					
$\Delta\pi_{t-12}$	0.036					
	(0.30)					
ΔRN_t	0.624					
	(1.57)					
ΔRN_{t-1}	0.685	0.844		−0.028	0.056	
	(1.57)	(2.06)		(−0.24)	(0.51)	
ΔRN_{t-2}	1.036	0.921		−0.243	−0.175	
	(2.31)	(2.40)		(−2.09)	(−1.68)	
ΔRN_{t-3}	1.069	1.127		−0.003	0.079	
	(2.48)	(2.91)		(−0.03)	(0.76)	
ΔRN_{t-4}	−0.488			0.008		
	(−1.12)			(0.075)		
Const.	(−0.160)	0.0001	0.0001	0.000	0.000	0.000
	(−1.15)	(0.36)	(0.18)	10.076	(0.87)	(1.08)
SEE	0.0038	0.0041	0.0044	0.0011	0.0011	0.0011
AIC	−994.32	−1076.82	−1070.54	−1336.36	−1338.49	−1338.92
DW	1.93	1.94	2.10	1.94	2.06	1.79
Q	8.06	23.98	29.46	10.90	11.40	20.80
	(0.99)	(0.63)	(0.33)	(0.99)	(0.99)	(0.79)

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$\chi^2_{30\ 01} = 11.34$. Therefore, the nominal interest rate can be said to 'Granger-cause' the inflation rate, a result that is not contrary to previous published results. However, testing the hypothesis that inflation does not 'Granger-cause' the interest rate produces $L = 6.61$, which does not reject the null. Since this is contrary to previous published results, we can only assume that it is chiefly due to the presence of the error-correction term.

The same procedure was followed using Australian data and the substantive results are presented in Table 5. The nature of these results is quite similar to those of Table 4. The inflation specification once again shows feedback from the interest rate. (The calculated L is 30.83.) In the interest-rate equation no individual lags or groups of lags were ever significant, with the exception of the error-correction term.

Table 5. *Error correction models, Australian data*

	Dependent variable:		
	$\Delta\pi$	$\Delta\pi$	ΔRN
EC_{t-1}	−0.280	−0.609	−0.264
	(−2.30)	(−5.07)	(−2.81)
$\Delta\pi_{t-1}$	−0.482		
	(−3.26)		
$\Delta\pi_{t-2}$	−0.229		
	(−1.47)		
$\Delta\pi_{t-3}$	0.043		
	(0.319)		
$\Delta\pi_{t-4}$	0.334		
	(2.83)		
ΔRN_{t-1}			
ΔRN_{t-2}			
ΔRN_{t-3}	−0.593		
	(−1.75)		
ΔRN_{t-4}	−1.56		
	(−5.18)		
Const.	0.001	0.000	0 000
	(1.27)	(0.50)	(0.72)
SEE	0.0075	0.0101	0.0030
AIC	−597.77	−567.61	−716.57
DW	1 63	1.21	2.07
Q	10.69	29.88	17.39
	(0.96)	(0.09)	(0.68)

V. SUMMARY AND CONCLUSIONS

This paper has demonstrated that for the US and Australian economies there is evidence that post-tax nominal interest rates and inflation are co-integrated and accordingly have a joint error-correction representation. This is important in that there is a large body of published empirical work concerning the existence of the Fisher effect, which appears somewhat contradictory. The econometric methodology has largely been to estimate single static equations, either in level or first-difference form, or the estimate vector autoregressions

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in first-difference form. Given the evidence of co-integration this empirical work represents a mis-specification, to the extent that error-correction considerations are omitted. Estimation of error-correction models for the US and Australian economies produces the result that the inflation rate 'Granger-causes' the post-tax nominal interest rate and that this interest rate responds only to the lagged error-correction term.

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