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Tests of the Fisher Hypothesis with International Data: Theory and Evidence

ALEX KANE, LEONARD ROSENTHAL, and GRETA LJUNG*

1. Introduction

THE FISHER HYPOTHESIS [7, 8] states that nominal interest rates adjust fully to expected inflation. For years tests of this hypothesis followed Fisher [8] in regressing nominal rates on a distributed lag of past inflation, which was used as a proxy for the expected rate of inflation. (See Roll [15] for references.) Fisher eventually concluded that nominal rates do not fully reflect expected inflation, saying that "when the cost of living is not stable, the rate of interest takes the appreciation and depreciation into account to some extent, but only slightly and in general indirectly."[1] Fisher attributed this to money illusion. In more recent years other researchers have explained the non-neutrality of inflation on grounds of general equilibrium (Sargent [16]), and taxes (Darby [3] and Feldstein [6]).

A new method of testing the Fisher hypothesis was introduced by Fama [4]. He concluded that interest rates set in U.S. financial markets are informationally efficient. Fama's work highlighted the two separate (although not always distinguishable) elements in the Fisher hypothesis: 1) the effect of expected inflation on the expected real rate, and 2) the question of whether or not the nominal rate reflects a rational forecast of expected inflation. To test the informational efficiency of nominal rates Fama [4] assumed a constant expected real rate, arguing in the later paper that the efficiency test is robust with respect to the assumption of the constancy of the expected real rate. This assumption has come under severe criticism (Carlson [2], Nelson and Schwert [14]) on the grounds that the efficiency results are not verifiable given the actual behavior of realized real rates. Summers [17] extended the study of interest rates to the period 1860 to 1979, using both Fisher's and Fama's methods, and found that the results are less favorable for the Fisher hypothesis. In a more recent study, Fama and Gibbons [5] tested the Fisher hypothesis using a procedure developed by Ansley [1] which assumes that the expected real rate follows a random walk. They concluded that nominal rates fully reflect expected inflation at the same time that the expected real rate adjusts to it.

This paper examines the question of whether international interest rates reflect expected inflation in their respective countries. In doing so, we expand on the previous work of Kane and Rosenthal [11]. We show that although Fama [4] introduced a bias into his tests, his general approach was correct. We then present simple time series tests and show that they fail for Eurocurrency rates.

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[1] Fisher [8], p. 43.

We attempt to improve on these tests by estimating expected real rates for each country using a Kalman filter. Finally, we run cross-sectional tests using Eurorate data of four different maturities. The results of these tests show that Eurorates include informationally efficient forecasts of inflation. When a stronger test of efficiency is run by looking at the time series behavior of independently generated residuals, the results are still favorable for efficiency.

II. Data Base Used in the International Tests of the Fisher Hypothesis

The international time series and cross-sectional tests analyzed in the following sections use data obtained from Data Resources Inc. The data includes monthly inflation indices and Eurocurrency rates for six countries (the United States, the United Kingdom, Switzerland, West Germany, France, and Holland) for one, three, six and twelve-month maturities. Monthly inflation rates were calculated from the respective one-month indices, and these rates were in turn compounded to arrive at three, six, and twelve-month inflation rates. The time period covered by the data is February 1974 through June 1979, a period of freely floating exchange rates.

Eurocurrency rates contain premia for default risk. Since data for default-free (Treasury) securities exist for the U.S. and the U.K., the default premia were calculated for these two Eurocurrency rates. That is $d_{jt} = R^*_{jt} - R_{jt}$, where R^*_{jt} is the Eurocurrency rate and R_{jt} is the Treasury bill rate.

III. Time-Series Tests

There are two elements to the question of whether nominal interest rates adjust to expected inflation: 1) Does the expected real rate vary, and, in particular, does it covary with expected inflation 2) Is the nominal rate set to earn a required rate of return so that it includes a premium equal to a rational forecast of inflation. Clearly any test of the adjustment of interest rates must be a joint test of the hypotheses about the two elements. Fama's [4] important contribution was in the introduction of a new methodology to test the second element. As indicated before, Nelson and Schwert [14] have refuted Fama's assumption of a constant expected real rate. Their analysis led to other approaches, as discussed below. It may be worthwhile to focus first on one of their specific arguments.

A. The Relevance of the Lack of Serial Correlation in Realized Real Rates

Fama [4] takes the lack of serial correlation in realized real rates to confirm the constancy of the expected real rate and the efficiency of nominal interest rates in incorporating expected inflation. That is, given a constant expected real rate, any variation in realized real rates must be due to inflation forecasting errors. In this case the lack of serial correlation in realized real rates indicates that nominal rates embody efficient inflation forecasts. Nelson and Schwert [14] have argued that the variance of the inflation forecasting error may conceal variation in expected real rates. They show that if the expected real rate follows a first-order autoregressive process, i.e., $\bar{r}_t = \phi\bar{r}_{t-1} + \delta_t$, and if the realized real rate contains a forecasting error, ϵ_t, so that $r_t = \bar{r}_t - \epsilon_t$, then

$$\text{Corr}(r_t, r_{t-1}) = \frac{\phi\sigma_{\bar{r}}^2}{\sigma_{\bar{r}}^2 + \sigma_\epsilon^2} \tag{1}$$

Therefore, a low serial correlation, as found by Fama, need not imply constancy of expected real rates. While their argument cannot be refuted, they failed to take something else into account. Suppose that the forecasting errors are correlated and follow an AR(1) process, $\epsilon_t = \psi\epsilon_{t-1} + \eta_t$, where ψ is the coefficient of correlation between successive forecasting errors. The correlation between present and past realized real rates is then given by

$$\text{Corr}(r_t, r_{t-1}) = \frac{\phi\sigma_{\bar{r}}^2 + \psi\sigma_\epsilon^2}{\sigma_{\bar{r}}^2 + \sigma_\epsilon^2} : \quad \begin{matrix} \text{Cov}(\delta, \eta) \overset{H}{=} 0 \\ \\ \text{Cov}(\delta, \epsilon) \overset{H}{=} 0 \end{matrix} \tag{2}$$

By looking at the components of (2) one can determine under what circumstances ψ is going to be the major factor determining the correlation in realized real rates. This will be the case if σ_ϵ^2 is large relative to $\sigma_{\bar{r}}^2$ and if ϕ is nonnegative. The research of Nelson and Schwert and Fama and Gibbons suggests that $\phi \simeq 1$ and $\sigma_\epsilon^2 \gg \sigma_{\bar{r}}^2$. Therefore if the correlation as given in (2) is near zero, then it is probably due to Ψ being equal to or close to 0. This implies that the lack of serial correlation in realized real rates is a reasonable initial indication that nominal rates incorporate expected inflation.

In this study we have performed serial correlation tests on realized real rates and default premia using the data set described above. The autocorrelation coefficients for twelve lags of the one and three-month realized real rates are presented in Table 1.The coefficients are generally small relative to their standard errors. Although the Box-Pierce Q-statistic based on all twelve lags is statistically significant at the 5 percent level in four out of the six countries for one-month rates, this appears to be mainly due to some seasonality in the data. If the Q-statistic is based on only 5 autocorrelation coefficients, its value is significant only for West Germany. The three-month rates exhibit no significant autocorrelation except for Holland, where the real rate seems to be negatively autocorrelated. The observation that expected real rates could be significantly autocorrelated and yet leave realized real rates virtually uncorrelated can be verified from data on default premia available for the U.S. and the U.K. The autocorrelation coefficients for these two series are shown in Table 2. Although the default premia of both the U.S. and the U.K. are highly serially correlated, these autocorrelations fail to show up in the observed real rates. The fact that default premia and expected real rates (as shown by Fama and Gibbons [5]) are serially correlated, while realized real rates are nearly independent supports the hypothesis that forecasting errors in real rates (which result from forecasting errors in expected inflation) are serially independent. This implies that nominal rates are efficient in incorporating inflationary expectations as suggested by Fama.

Table I

Autocorrelation of One and Three Month Real Eurorates[a]

Country	Lags 1	2	3	4	5	6	7	8	9	10	11	12	Box-Pierce Q-Stat.[b]	$\bar{\bar{r}}$	$S(r)$
–One Month Rates–															
U.S.	.13	.16	.11	−.19	−.13	−.30*	−.13	−.24*	−.03	.11	.06	.13	19.86*	−.00048	.00265
U.K.	.22	.06	.13	−.05	.10	.29*	−.03	.21	−.13	−.14	.03	.25*	20.18*	−.00219	.00820
Switzerland	.06	.25*	−.04	.12	−.07	.07	−.17	.02	−.11	.13	.01	.13	10.87	−.00000	.00448
W. Germany	.36*	.21	.07	.02	.04	−.36*	−.11	−.10	−.02	.00	.17	.39*	33.39*	.00070	.00293
France	.29*	.21	.13	−.04	.07	.02	−.04	.06	−.11	.07	−.07	.05	11.80	.00078	.00344
Holland	.20	.07	−.14	−.12	.27*	.13	.13	−.11	−.33*	−.01	.07	.41*	31.18*	.00006	.00729
–Three Month Rates–															
U.S.	.08	−.32	.00	.36	.00								5.24	−.00046	.00623
U.K.	.26	−.14	−.24	−.09	.19								4.16	−.00716	.01532
Switzerland	.20	−.12	−.03	.24	−.01								2.49	.00159	.01017
W. Germany	.32	−.02	.01	.24	−.10								3.75	.00274	.00676
France	.16	−.06	−.12	−.17	−.06								1.67	.00274	.00924
Holland	−.37*	.61*	−.47*	.50*	−.45*								26.01*	.00077	.01345

[a] Values significant at the 10% level are noted with an asterisk. The asymptotic standard error is .12 for the one month rates and .21 for the three month rates.

[b] The critical χ^2 value for the Q-statistic at the 10% level is 18.5 for the one month rates and 9.2 for the three month rates.

Table 2

Autocorrelation of One and Three Month Default Premia[a]

Country	Lags 1	2	3	4	5	6	7	8	9	10	11	12	Box-Pierce Q-Stat.[b]	$\bar{\bar{d}}$	$S(d)$
–One Month Rates–															
U.S.	.69*	.52*	.47*	.30*	.32*	.21	.13	.09	−.03	−.06	−.03	.06	80.46*	.00099	.00128
U.K.	.64*	.46*	.32*	.20	.14	.06	.08	.06	.05	.11	.18	.03	55.69*	.00178	.00201
–Three Month Rates–															
U.S.	.41	.35	.10	.05	.09								6.85	.00341	.00375
U.K.	.44*	.19	.16	.03	−.29								7.49	.00589	.00567

[a] Values significant at the 10% level are starred. The asymptotic standard error is .12 for the one month rates and .21 for the three month rates.

[b] The critical χ^2 value for the Q-statistic at the 10% level is 18.5 for the one month rates and 9.2 for the three month rates.

Further evidence of the efficiency of nominal rates is given by the autocorrelation coefficients of realized inflation rates as shown in Table 3. The autocorrelation coefficients in one and three-month inflation rates are greater than those for realized real rates for all six countries. This is consistent with the hypothesis that nominal rates adjust sufficiently to expected inflation to produce reasonably serially uncorrelated realized real rates. When one takes into account the existence of serially correlated default premia, the results are even stronger in supporting the efficiency of nominal rates.

B. Regression of the Rate of Inflation on Nominal Interest Rates

As stated earlier, tests of the Fisher hypothesis include a statement about the neutrality of inflation, i.e., its effect or, rather lack thereof, on expected real rates. Let the expected real rate be given by $\bar{r}_t = \mu\bar{i}_t + \alpha_t$, where $\bar{i}_t$ is expected inflation, μ is a measure of the impact of expected inflation on the expected real rate, and α_t is a measure of changes in the expected real rate due to factors other than inflation. The nominal interest rate can now be written as

$$R_t = \bar{r}_t + \beta\bar{i}_t = \alpha_t + (\beta + \mu)\bar{i}_t \tag{3}$$

where β is the reaction of the nominal rate to inflationary expectations. The Fisher hypothesis is then equivalent to stating that $\mu = 0$ and $\beta = 1$.

In Fama's [4] original tests, he assumed that $\mu = 0$ and $\alpha_t = \bar{r}_t$, and estimated the specification for the inflation rate as

$$i_t = \bar{i}_t + \epsilon_t = a + bR_t + e_t \tag{4}$$

The hypothesis that $b = 1$, and $a = -\bar{r}$ was then tested. If $\mu < 0$, as has been argued by Mundell [3], and if $0 < \beta < 1$, the specification of (4) would be

$$\dot{i}_t = \frac{1}{\mu + \beta}\alpha_t + \frac{1}{\mu + \beta}R_t + \epsilon_t \tag{5}$$

Thus, in testing (4), b would be expected to be greater than one for large samples. If α_t is not constant and is correlated with the nominal rate, there will be a bias in estimating the coefficient on R_t as given in (4) such that

$$\begin{aligned} E(b) &= \frac{1}{\mu + \beta}\left[1 - \frac{\text{Cov}(\alpha_t, R_t)}{\text{Var}(R_t)}\right] \\ &= \frac{1}{\mu + \beta}\left[1 - \frac{\sigma_\alpha^2}{\sigma_\alpha^2 + (\mu + \beta)^2\sigma_{\bar{i}}^2}\right] \end{aligned} \tag{6}$$

An estimated value of b greater than 1 for a sufficiently large sample, would suggest that the true value of $\mu + \beta$ is less than 1. If the true value of $\mu + \beta$ is greater than 1, then one would expect to estimate a value of b which is less than 1, while the reverse is not true. In other words, if the bias makes the estimated coefficient on R_t smaller than its true value, and if the estimated value is greater than 1, then the true value of b is expected to be greater than 1. On the other hand, if the estimated value of b is less than 1, this could well be due to the bias

Table 3

Autocorrelation of One and Three Month Inflation Rates[a]

Country	Lags 1	2	3	4	5	6	7	8	9	10	11	12	Box-Pierce Q-Stat.[b]	$\bar{i}$	$S(i)$
–One Month Rates–															
U.S.	.38*	.40*	.33*	.12	−.16	.05	.11	.02	.16	.20	.15	.15	37.63*	.00677	.00321
U.K.	.35*	.20	.22	.03	.12	.27	.04	−.12	−.03	−.05	.06	.29*	26.40*	.01221	.00880
Switzerland	.12	.35*	.13	.25*	.05	.15	−.11	.07	−.11	.06	−.04	.10	18.56*	.00263	.00455
W. Germany	.26*	.21	.09	−.05	.00	−.43*	−.09	−.06	.07	.08	.26*	.49*	41.46*	.00353	.00282
France	.15	.26*	.16	−.07	.02	.07	.01	.08	−.04	.10	−.03	.33*	17.51	.00839	.00302
Holland	.15	.03	−.24	.10	.36*	.43*	.31*	−.01	−.27	.02	.13	.52*	56.05*	.00577	.00537
–Three Month Rates–															
U.S.	.44*	.17	.17	.21	−.04								6.54	.02064	.00807
U.K.	.44*	.06	−.02	.04	.06								4.46	.03858	.01922
Switzerland	.42*	−.00	−.05	.05	−.07								4.10	.00794	.00998
W. Germany	.23	−.25	.11	.44*	−.01								7.07	.01078	.00560
France	.29	−.14	.06	.24	−.04								3.66	.02561	.00684
Holland	−.36	.73*	−.31	.64*	−.30								27.68*	.01751	.01285

[a] Values significant at the 10% level are noted with an asterisk. The asymptotic standard error is .12 for the one month rates and .21 for the three month rates.

[b] The critical χ^2 value for the Q-statistic at the 10% level is 18.5 for the one month rates and 9.2 for the three month rates.

and not because the true value is actually less than 1. In this case, the coefficient is not a good test statistic and no firm conclusion can be drawn from the tests. This is the case with Fama's [4] tests, where the coefficient b is less than unity for all T-bill maturities for the period 1953–1971. Kane and Rosenthal [11] have run time series tests of equation (4) over the period February 1974 to June 1979, using the Eurocurrency rates of six countries described in Section II. As indicated before, these rates contain default premia which may increase the bias given in (6). Estimates of (4) using one month rates are presented in Table 4a. All estimates of the coefficient, b, are below 1, and in most cases are significantly so. These estimates are most likely downward biased, impairing the value of this form of the time series tests.

C. Regression with Variable Expected Real Rates

Fama and Gibbons [5] used T-bill data for the period 1953–1977 and first estimated (4) obtaining $b = 1.36$ for one-month bills and $b = 1.32$ for three-month bills. These results would lead one to accept the hypothesis that $\beta + \mu \leq 1$. Furthermore, there was serial correlation in the residuals, which suggests that either the expected rate is serially correlated and/or the efficiency hypothesis should be rejected. Next, Fama and Gibbons used an estimation procedure developed by Ansley [1] that allows the expected real rate to follow a random walk, $\bar{r}_t = \bar{r}_{t-1} + u_t$. The expected real rate is estimated simultaneously with

$$i_t = -b\hat{\bar{r}}_t + bR_t + \epsilon_t \tag{7}$$

Since nominal rates are set as $R_t = \bar{r}_t + \beta\bar{i}_t$, the relationship in (7) can be written as

$$\dot{i}_t = -\frac{1}{\beta}\,\bar{r}_t + \frac{1}{\beta}\,R_t + \epsilon_t. \tag{8}$$

Using this method, they were able to correct for both the Mundell effect and the bias in the regression coefficient, b. Estimating (7), Fama and Gibbons obtained values of $b = 1.15$ for one-month rates and $b = .93$ for three-month rates, which are much closer to the value expected for b, i.e., 1.

Table 4(a)

Estimated Coefficients (b) for Time Series Test—Equation (4)—with One-Month Eurorates—February 1974–June 1979

Country	Nominal Rate Coefficient(b)	$S(b)$
U.S.	.89	.17
U.K.	.95	.35
Switzerland	.44	.21
West Germany	.34	.19
France	.30	.11
Holland	.24	.24

We estimated conditional expected real rates for the Eurocurrency market assuming that these rates follow a random walk. A Kalman filter was used to extract one month expected real rates and these were tested to see if they were constant over time using a procedure discussed by Ledolter [12]. For three of the six countries, we could not reject the null hypothesis that expected real rates are constant over time. For the other countries, even though the hypothesis was rejected, the variability in the respective series was relatively small.

The estimates of expected real rates discussed above were used in testing (7). The slope coefficients from these regressions are presented in Table 4b. Using the Kalman filter gets around the problem caused by the Mundell effect, lowering the estimate of the coefficient on R_t. At the same time, our method eliminates the previously discussed bias which pushes up the estimate of b. Comparing Table 4b with Table 4a, one can see that the coefficient for the U.S. drops, implying that the Mundell effect is the overriding factor affecting b. For the other five countries, the coefficient increases, implying that the bias is the overriding factor affecting b for these countries. It is worth noting that just as for the estimation procedure used by Fama and Gibbons, here too the t-statistic and the R^2s are hard to interpret.

IV. International Cross-Sectional Tests

An alternative to time series tests of the Fisher hypothesis is to perform a cross-sectional test using international interest rates. With perfect international capital markets, expected real rates will be similar across countries (see Giddy, [9]), lending power to the simple specification (4). For country j, $j = 1, \cdots, J$ the expected real rate is $\bar{r}_j = \bar{r} + v_j$, where $\bar{r}$ is an underlying expected real rate for all countries. Across country variance in real rates is given by σ_v^2. For nominal country Eurorates, R_j, it is necessary to include a default premium d_j with variance σ_d^2. Denoting country-expected inflation $\bar{i}_j$ the analogue of specification (4) becomes:

$$i_j = g + hR_j + k_j \tag{9}$$

In an international setting, $R_j = \bar{r}_j + \beta\bar{i}_j$, where β is a factor which reflects

Table 4(b)

Estimated Coefficients (b) for Time Series Test—Equation (7)—with One-Month Eurorates—February 1974–June 1979

Country	Nominal Rate Coefficient (b)	$S(b)$
U.S.	.81	.14
U.K.	1.26	.22
Switzerland	1.00	.20
West Germany	1.31	.19
France	.71	.14
Holland	.68	.25

Table 5

Average Estimates of h with Specification (12) from 65 Cross-Section Samples Over the Period 2/74–6/79

	h	$t(h - 1.0)$	R^2	DW
Average for 65 Months	1.33	1.29	.90	1.82
Average for Months 1–32	1.67	3.01	.87	1.81
Average for Months 33–65	1.01	−.37	.94	1.83

adjustment of nominal rates to expected inflation. Thus, the true relationship is

$$\dot{i}_j = -\frac{1}{\beta} r - \frac{1}{\beta}(v_j + d_j) + \frac{1}{\beta} R_j + \epsilon_j. \tag{10}$$

Estimating (9) leads to a downward bias in the estimate of h so that

$$E(h) = \frac{1}{\beta}\left[1 - \frac{\sigma_v^2 + \sigma_d^2 + \sigma_{v\bar{i}}}{\sigma_v^2 + \sigma_d^2 + \sigma_{\bar{i}}^2 + 2\sigma_{v\bar{i}}}\right], \tag{11}$$

where $\sigma_{v\bar{i}}$ is an international Mundell effect. It is assumed that default premia are quite similar across Eurorates at any point in time, and that the variance of the expected real rate is also small relative to the variance of expected inflation.[2]

The practical obstacle to this approach is the small number of countries with competitive markets. As mentioned above, we were able to obtain data for six Eurocurrency markets. To increase the sample size, Kane and Rosenthal [11] used four different maturities of Eurorates and inflation rates—one, three, six and twelve months. Denote the nominal Eurorates for these four maturities by R_{1j}, R_{3j}, R_{6j}, R_{12j}, and similarly for inflation rates i_{1j}, i_{3j}, i_{6j}, and i_{12j}. In short, let R_{tj} represent Eurorates and i_{tj} represent inflation rates where $t = 1, 3, 6, 12$ and $j = 1, \ldots, 6$. In order to prevent correlation in the residuals of the same country, due to the fact that longer maturity rates include those of shorter maturities, we subtracted from each maturity the rate of the next shorter maturity. (For a given country the one-month pair of Eurorates and inflation rates were left unchanged.) Thus for the three-month rate, we use $R_{3j}^* = R_{3j} - R_{1j}$; for the six-month rates we use $R_{6j}^* = R_{6j} - R_{3j}$, etc. The same procedure is used for the inflation rates. Since four maturities were used in each regression, three dummy intercept variables were included in each regression—D_3, D_6, D_{12}. The test analogue to equation (10) is

$$i_{tj}^* = g_0 + g_3 D_3 + g_6 D_6 + g_{12} D_{12} + h R_{tj}^* + u_t \tag{12}$$

with $4 \times 6 = 24$ observations for each cross-section. These cross-sectional regressions were run for the 65 time periods in our sample. The average statistics are presented in Table 5. The results are quite interesting. Observe that the explanatory power of nominal interest rates is very high (the average R^2 is .90 for the 65 cross-sectional regressions). In the cross-sectional tests we accept the

[2] These two assumptions seem reasonable in light of the results obtained by Fama and Gibbons [5] in their tests on expected real rates and expected inflation rates in the U.S.

Table 6

Autocorrelation Coefficients of One Month Residuals From Cross-Sectional Regressions[a]

Lags	1	2	3	4	5	6	7	8	9	10	11	12
U.S.	.081	.046	.201	−.044	.247	.156	−.047	−.111	−.100	.041	.023	.070
U.K.	.251*	.120	.222	.013	.010	.222	−.094	−.304	−.230*	−.254*	−.162	.057
Switzerland	.353*	.264	.373*	.205	.055	.101	.058	−.152	−.047	−.089	.163	.021
W. Germany	.341*	.051	−.241	−.204	−.047	−.020	−.005	−.113	−.291	−.157	.169	.348*
France	.489*	.336*	.309*	.363*	.291*	.118	.184	.262*	.135	.123	.102	.185
Holland	.442*	.109	−.052	−.046	.182	.214	.084	−.098	−.198	−.072	.086	.257*

[a] * Indicates that the autocorrelation coefficient is more than 2 standard errors from zero.

fact that the coefficient h may be large due to a strong Mundell effect, i.e., $\sigma_{v\bar{i}} < 0$, and at the same time be downward biased due to across-country variation in expected real rates. Indeed, the average value of the coefficient h for the 65 cross sections is quite large, 1.33, although not significantly different from unity. It would seem that the cross-sectional procedure using international interest rates is at least as good as the time-series tests with variable intercept coefficients.

The results above provide a clear and positive indication that nominal interest rates adjust to expected inflation. We ran a further test of nominal rate efficiency by examining the time series relationship of the residuals generated by the cross-sectional regressions. That is, a residual from each maturity and for each country is independently estimated from each regression. We generated 24 sets of residuals. Therefore, each time series contains 65 independent estimates for each residual. For example, for the U.S., we have 65 residuals arranged in a time series from February 1974–June 1979. The results of autocorrelation tests on the one-month series by country are presented in Table 6. By and large, the residuals for one-month maturity are generally not autocorrelated, or only moderately so. In particular, the one-month residuals for the U.S. are totally uncorrelated. This confirms the result of Fama and Gibbons that nominal U.S. interest rates are informationally efficient with respect to inflation. In general, then, the cross-sectional tests provide reasonable evidence of market efficiency.

V. Conclusions

This paper reviews and extends some recent tests on the efficiency of nominal interest rates with respect to expected inflation. We show that Fama's [4] time series tests were in the right spirit but these tests cannot be relied on due to possible biases in the coefficients obtained and the Mundell effect. The work done by Fama and Gibbons [5] substantially improves on Fama's earlier work and leads one to conclude that nominal interest rates are efficient. We show that cross-sectional tests using international interest rates can also provide a good alternative test of efficiency. Results of those tests provide reasonable, although not completely conclusive support for the concept of nominal interest rates adjusting to expected inflation.

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