Forecasting Inflation Using Interest-Rate and Time-Series Models: Some International Evidence

Author(s): R. W. Hafer and Scott E. Hein

Source: *The Journal of Business*, Jan., 1990, Vol. 63, No. 1, Part 1 (Jan., 1990), pp. 1-17
Published by: The University of Chicago Press
Stable URL: https://www.jstor.org/stable/2353234

R. W. Hafer
Southern Illinois University at Edwardsville

Scott E. Hein
Texas Tech University and Federal Reserve Bank of Dallas

Forecasting Inflation Using Interest-Rate and Time-Series Models: Some International Evidence*

It has been suggested that inflation forecasts derived from short-term interest rates are as accurate as time-series forecasts. Previous analyses of this notion have focused on U.S. data, providing mixed results. In this article we extend previous work by testing the hypothesis using data taken from the U.S. and five other countries. Using monthly Euro-rates and the consumer price index (CPI) for the period 1967–86, our results indicate that time-series forecasts of inflation have equal or lower forecast errors and have unbiased predictions more often than the interest-rate-based forecasts.

I. Introduction

Numerous studies have investigated the relative accuracy of alternative inflation forecasting models. One approach has been to compare the accuracy of survey respondents' inflation forecasts relative to univariate time-series models.[1] Another approach is the methodology associated with the work of Fama (1975, 1977) and recently extended by Fama and Gibbons (1982, 1984). This approach extracts from observed nominal interest rates the market's inherent expectation of inflation. Based on a univariate time-series modeling of the real interest rate, Fama and Gibbons (1984) find that the interest-rate model yields inflation forecasts with a lower error variance than a univariate model, and that the inter-

* Thanks to Gerald Dwyer, Douglas Pearce, Daniel Thornton, Roy Webb, Victor Zarnowitz, participants at the Texas Tech Finance Workshop, and an anonymous referee for useful comments and suggestions that have improved this article. We also would like to thank Dawn Peterson for her research assistance. The views expressed are the authors' and may not represent those of the Federal Reserve Bank of Dallas or of the board of governors of the Federal Reserve System.

1. Examples of such studies are Pearce (1979) and Brown and Maital (1981).

(*Journal of Business*, 1990, vol. 63, no. 1, pt. 1)

0021-9398/90/6301-0001$01.50

est-rate model's forecasts dominate those calculated from the Livingston survey.

Although a flurry of articles appeared after Fama's (1975) original article—articles that focused on Fama's assumption of a constant real rate of interest—only a few studies have examined the forecasting approach detailed in Fama and Gibbons (1984). For example, using quarterly U.S. data, Hafer and Hein (1985) compare the relative forecasting accuracies of the interest-rate model, a univariate time-series model of inflation, and forecasts taken from the American Statistical Association–National Bureau of Economic Research (ASA-NBER). Based on ex ante forecasts for the 1970–84 period, they find that the survey forecasts generally have the greater relative accuracy. Hein and Koch (1988) compare inflation forecasts derived from tax-exempt yields to those from taxable Treasury bills. Comparing monthly out-of-sample forecasts for the period 1978–86, they do not find any improvement in inflation forecasts by using the tax-exempt yields.

Evidence on the usefulness of this procedure has come exclusively from U.S. data. Our purpose in this article is to extend previous analyses by determining whether the results obtained for the United States generalize to other countries. In other words, using Fama and Gibbons's methodology, we address the question of whether nominal rates in other countries also provide accurate forecasts of their rates of inflation. As a basis of comparison, the interest-rate forecasts of inflation are pitted against forecasts generated by simple univariate time-series models using data from the United States and five other industrial countries: Belgium, Canada, England, France, and Germany. Using ex ante monthly forecasts of inflation spanning the period 1978–86, the accuracy of the time-series-model forecasts is compared with forecasts from the interest-rate model for each country.

The article is organized as follows. The time-series model for each country is constructed and estimated in Section II. Section III presents the different interest-rate models, providing estimates of the crucial real interest-rate series. Section IV discusses the accuracy of the models' forecasts. The article closes with summary remarks in Section V.

II. Time-Series Model Estimates

Univariate time-series models often are used as a basis for comparing alternative forecasts. Because these models rely only on information contained in the variable's own past, failure to improve on forecasts from these models leads to a strong rejection of the alternative forecasts. To construct the time-series models, sample autocorrelations of each country's monthly CPI inflation rate were examined.[2] The auto-

2. Data on the CPI are taken from the International Financial Statistics data tape. The inflation rates are measured as annualized changes in the logarithm of the monthly price level.

correlations for the levels and first difference of inflation rate in the six countries are reported in table 1. These estimates use data from 1967 through 1977. Because of data limitations, the data for France begin in 1970.

The autocorrelations of the inflation rates reveal a relatively slow decay for most countries, suggesting a nonstationary series. Seasonality is evidenced by large autocorrelations at the twelfth lag for England and Germany. Because the inflation-rate series do not appear stationary, first differences are examined. In every instance, the first difference of the inflation rate indicates the characteristics of a stationary series.[3] Moreover, the first-order autocorrelation coefficient always is larger than twice its standard error, suggesting a first-order moving-average (MA) model in the first difference of the inflation rate. For England and Germany, however, there remains a relatively large autocorrelation coefficient at lag 12, indicating the presence of a seasonal factor. Seasonal factors aside, the autocorrelations suggest that inflation follows a similar time-series process across the various countries studied.

Based on the autocorrelations reported in table 1, first-order moving-average models were fitted to the change in the inflation-rate series for each country. For England and Germany, a seasonal component also was estimated. The results from fitting these models are reported in table 2. In all cases the estimated MA parameters are statistically significant at the 1% level. Moreover, the reported Q-statistics indicate that the fitted models reduce the residuals to white noise. The largest Q-statistic, that for Canada, does not reject the null hypothesis of white noise residuals at the 8% level of significance. The statistical results do not reject the usefulness of the MA(1,1) specification (with seasonals where appropriate) to model inflation in the six countries used. The time-series properties of inflation thus appear to be fairly similar across our sample of countries.

An interesting aspect of the estimation results is the size of the parameter estimates relative to previous evidence based on U.S. Treasury bill rates. The estimate for the United States in table 2, 0.776, is in line with that reported by Fama and Gibbons (1984); their estimate, based on monthly data for the period 1953–77, is 0.8027. Pearce (1979) also found an MA(1,1) model to fit U.S. monthly inflation for the period 1947–75, reporting that the estimated MA parameter varied between 0.71 and 0.76 depending on the subsample of data used. Using quar-

3. An alternative approach to testing for stationarity is that of Dickey and Fuller (1979). Recently, however, Sims (1988) has argued that the Dickey and Fuller approach for detecting unit roots does not adequately distinguish between a unit root and an autocorrelation coefficient that is large and stationary. Whitt (1988) applies both the Dickey and Fuller and the Sims tests to data on the real exchange rates for five countries and finds that the test results are contradictory. Although the use of other test procedures would be a useful diagnostic on the models used here, we will employ the Box-Jenkins methodology to enhance the comparability of our results with previous work in this area.

TABLE 1 Autocorrelations of Inflation Rates: Levels and Differences, June 1967–December 1977

Variable	ρ_1	ρ_2	ρ_3	ρ_4	ρ_5	ρ_6	ρ_7	ρ_8	ρ_9	ρ_{10}	ρ_{11}	ρ_{12}
Belgium:												
$\dot{P}_t$	.581	.524	.558	.583	.535	.464	.392	.399	.382	.318	.278	.361
$(1 - B)\dot{P}_t$	−.440	−.102	.019	.075	.028	−.001	−.087	.025	.048	−.020	−.152	.196
Canada:												
$\dot{P}_t$	.417	.247	.277	.344	.270	.252	.266	.246	.353	.144	.259	.355
$(1 - B)\dot{P}_t$	−.347	−.180	−.039	.124	−.045	−.024	.025	−.106	.266	−.275	.012	.144
England:												
$\dot{P}_t$	.429	.321	.327	.191	.169	.360	.146	.087	.165	.106	.204	.492
$(1 - B)\dot{P}_t$	−.414	−.100	.129	−.092	−.193	.364	−.146	−.102	.102	−.119	−.178	.471
France:*												
$\dot{P}_t$	.548	.480	.423	.380	.362	.388	.328	.287	.306	.220	.180	.209
$(1 - B)\dot{P}_t$	−.432	−.004	−.003	−.051	−.020	.077	−.013	−.037	.091	−.067	−.092	.175
Germany:												
$\dot{P}_t$	.370	.253	.180	.009	−.103	−.060	−.099	.046	.150	.197	.370	.509
$(1 - B)\dot{P}_t$	−.407	−.042	.083	−.045	−.124	.064	−.135	.023	.043	−.100	.027	.248
United States:												
$\dot{P}_t$	.331	.418	.357	.309	.339	.303	.296	.239	.277	.256	.205	.123
$(1 - B)\dot{P}_t$	−.564	.106	−.009	−.059	.050	−.016	.032	−.065	.043	.022	.019	−.134

NOTE.—ρi (i = 1, . . . , 12) are the sample autocorrelations at lag i.
* Sample period is January 1970–December 1977.

TABLE 2 Estimated Inflation-Rate Models, June 1967–December 1977

Country	Model*	SE†	Q(df)‡
Belgium	$(1 - B)\dot{P}_t = (1 - \underset{(11.87)}{.727}B)a_t$	3.25	12.90 (11)
Canada	$(1 - B)\dot{P}_t = (1 - \underset{(18.39)}{.853}B)a_t$	4.00	17.99 (11)
England	$(1 - B)(1 - B^{12})\dot{P}_t = (1 - \underset{(9.54)}{.675}B)(1 - \underset{(9.18)}{.688}B^{12})a_t$	7.04	15.23 (10)
France§	$(1 - B)\dot{P}_t = (1 - \underset{(9.16)}{.698}B)a_t$	3.16	3.68 (11)
Germany	$(1 - B)(1 - B^{12})\dot{P}_t = (1 - \underset{(16.54)}{.845}B)(1 - \underset{(9.18)}{.658}B^{12})a_t$	3.23	9.00 (10)
United States	$(1 - B)\dot{P}_t = (1 - \underset{(13.74)}{.776}B)a_t$	3.04	5.04 (11)

* Figures in parentheses beneath each model are t-statistics.
† SE is the standard error of the estimated model.
‡ The reported Q-statistic, which is distributed as a χ^2, is used to test for white noise residuals. With 10 df, the 5% critical value is 18.3; with 11 df, it is 19.7.
§ Estimated sample is January 1970–December 1977.

terly data from 1953 through 1969 for the gross national product (GNP) deflator, Hafer and Hein (1985) also estimate an MA(1,1) model and find the coefficient to be 0.81.

The estimates for the other countries in turn are close to the U.S. estimates. For example, the smallest parameter estimate is that for England (0.675) and the largest is for Canada (0.853). The relatively small difference in estimates, and the fact that the different inflation series all can be fitted by simple MA models, indicates similar processes generating each of the respective series. Another similarity is that each country's model estimate indicates that the variance of the expected component of inflation is greater than the variance of the unexpected part. For example, the average value of the MA parameters across the countries (0.762) implies that about 24% of any given shock to inflation is absorbed by the rate-of-inflation process (Box and Jenkins 1976, p. 144). The models reported in table 2 are used to generate forecasts of inflation.

III. Interest-Rate Models

The procedure by which inflation forecasts are extracted from observed nominal interest rates is based on the Fisher equation. This familiar equation is written as

$$R^t_{t-1} = r^t_{t-1} + \dot{P}^t_{t-1}, \tag{1}$$

where R^t_{t-1} is the nominal interest rate observed at the end of period $t - 1$ that holds over period t, r^t_{t-1} is the real interest rate expected to hold over the period $t - 1$ to t, and $\dot{P}^t_{t-1}$ is the expectation at period's end in $t - 1$ for the inflation over the period $t - 1$ to t.

Fama's (1975) original work with equation (1) imposed the constraint that the real rate is constant. Evidence presented by Hess and Bicksler (1975), Fama (1976), Carlson (1977), Nelson and Schwert (1977), Garbade and Wachtel (1978) and Fama and Gibbons (1982) rejects the notion of a constant real rate, indicating that ex post real-rate series displays significant variation over time. In general, these studies (all based on U.S. data) suggest that the expected real rate behaves as a random walk. Mishkin (1984), analyzing Euro–deposit rates in seven Organization for Economic Cooperation and Development (OECD) countries for the second quarter 1967 through the second quarter 1979, also rejects the constancy of real rates.

If the real rate behaves as a random walk, then changes in the observed, ex post real interest rate $(R_{t-1} - \dot{P}_t)$ can be modeled as a simple moving-average model. In other words, if the ex post real rate can be written as

$$R_{t-1} - \dot{P}_t = r^t_{t-1} + \epsilon_t, \tag{2}$$

then changes in the real return can be captured in the time-series model

$$(R_{t-1} - \dot{P}_t) - (R_{t-2} - \dot{P}_{t-1}) = a_t - \Theta a_{t-1}, \tag{3}$$

where Θ is an estimable moving-average parameter. Using U.S. data for a 1-month Treasury bill rate observed at the end of month $t - 1$ and the CPI measure of inflation, Fama and Gibbons (1984) estimate equation (3) for the period 1953–77 and find that this model is not rejected by the data. Hafer and Hein (1985), using quarterly data, also find that the MA(1,1) model approximates the behavior of the ex post real-rate series quite well.

Adequate modeling of the real interest rate is the cornerstone to the interest-rate-model approach to forecasting inflation. To see this, simply rewrite equation (1) as

$$\dot{P}^t_{t-1} = -r^t_{t-1} + R^t_{t-1}. \tag{4}$$

A forecast of inflation derived from the observed nominal interest rate is obtained by subtracting the ex ante forecast of the real rate (from eq. [3]) from the nominal interest rate observed at the end of period $t - 1$.

Previous studies using this forecasting procedure have relied on U.S. Treasury bill data. Extending the analysis to other countries raises the problem of consistent data series over time and across countries. In this article we use 1-month Eurocurrency rates reported by the Harris Bank of Chicago. Since these data are reported on a weekly basis, that is, each Friday, we take for R^t_{t-1} that rate closest to, but not beyond, the end of the month. Although one may argue that the Euro-rate is not the optimal measure (for instance, it may incorporate a time-varying default premium), lack of comparable end-of-month data for government interest rates across our sample of countries restricts us to this series. Eurocurrency rates commonly are used in the literature dealing with the behavior of the real rate (see, inter alia, Kane and Rosenthal [1982]; Mishkin [1984]; Mark [1985]; and Cumby and Mishkin [1986]). These rates have the property that they are likely to be similar in risk across countries, are market clearing, and are not subject to direct domestic controls. Thus the 1-month Eurocurrency rate, along with the CPI measure of inflation, is used to generate the ex post real-rate series for each country.

To obtain an ex ante forecast of the real rate, appropriate time-series models are constructed. We first test whether the change in the ex post real-rate series for each country can be modeled by an MA process, as suggested in previous studies. To do this, we examine the sample autocorrelations for the level and first differences of the series. The autocorrelations of the levels data, reported in table 3, decline slowly or show little change across the 12 lags. When the series are differenced, however, each autocorrelation pattern shown in table 3 is indicative of a moving-average process.

In the two instances in which a seasonal is observed in the inflation rate series, there remains a seasonal factor in the difference real-rate series. For example, the autocorrelation at lag 12 is well over twice the standard error for England and Germany. This result indicates that the proper model includes a seasonal factor, a finding that may appear at odds with the underlying theory of efficient markets. It should be noted, however, that seasonality in the real rate comes from the inflation data: examination of the sample autocorrelations for English and German *nominal* Eurocurrency rate series reveals no seasonal in the data.[4]

Based on the autocorrelations of the changes of the different real-rate series, MA(1,1) models are fitted to the data for each country. The sample period again is 1967–77. The estimated models, reported in table 4, capture the behavior of the monthly real-rate series. The reported Q-statistics indicate that the models' residuals are not different from white noise. Moreover, the estimated coefficients all achieve significance at the 1% level. Thus, the estimated MA models in table 4 are not rejected by the data.

The results for the United States again conform with previous results based on domestic Treasury bill rates. Fama and Gibbons (1984) estimate Θ to be 0.922 using 1-month Treasury bill rates. Hafer and Hein (1985) report the value of Θ to be 0.810 based on quarterly observations. The current estimate of 0.912 suggests that the time-series prop-

4. The autocorrelations for the change in the nominal Eurocurrency rates are: England −.28, −.23, .16, .06, −.01, .01, −.11, .03, .05, −.09, −.01, and .04. For Germany, the autocorrelations are −.17, −.25, −.03, .16, .08, −.08, .14, −.18, .03, .06, −.02, and −.01. These results indicate that the seasonal in the real rates for these countries comes from the seasonal in the CPI data (see table 1).

TABLE 3 Autocorrelations of Real Interest Rates: Levels and Differences, June 1967–December 1977

Variable	ρ_1	ρ_2	ρ_3	ρ_4	ρ_5	ρ_6	ρ_7	ρ_8	ρ_9	ρ_{10}	ρ_{11}	ρ_{12}
Belgium:												
r_t	.535	.450	.513	.441	.379	.271	.236	.251	.224	.163	.144	.190
$(1 - B)r_t$	−.410	−.155	.145	−.013	.051	−.081	−.054	.047	.035	−.045	−.070	.105
Canada:												
r_t	.395	.200	.231	.275	.158	.131	.151	.158	.239	.014	.121	.240
$(1 - B)r_t$	−.336	−.192	−.009	.136	−.079	−.036	.008	−.057	.249	−.275	−.009	.154
England:												
r_t	.360	.196	.264	.138	.122	.300	.073	.036	.099	.031	.150	.437
$(1 - B)r_t$	−.376	−.181	.154	−.081	−.154	.317	−.152	−.074	.098	−.139	−.140	.456
France:*												
r_t	.319	.222	.216	.147	.095	.102	.056	.068	−.024	.090	−.052	−.057
$(1 - B)r_t$	−.425	−.061	.057	−.041	−.017	.024	−.025	.094	−.166	.168	−.105	.042
Germany:												
r_t	.414	.247	.247	.149	.069	.046	.048	.073	.087	.083	.159	.267
$(1 - B)r_t$	−.357	−.145	.086	−.015	−.047	−.025	−.015	.006	.015	−.068	−.032	.239
United States:												
r_t	.139	.236	.205	.127	.134	.110	.178	.137	.153	.145	.103	.129
$(1 - B)r_t$	−.556	.074	.028	−.050	.019	−.053	.063	−.032	.013	.021	−.041	−.074

NOTE.—ρi (i = 1, . . . , 12) are the sample autocorrelations at lag i.
* Sample period is January 1970–December 1977.

TABLE 4 Estimated Real-Interest-Rate Models, June 1967–December 1977

Country	Model*	SE†	Q(df)‡
Belgium	$(1 - B)r_t = (1 - \underset{(9.46)}{.663}B)a_t$	3.83	11.46 (11)
Canada	$(1 - B)_t = (1 - \underset{(15.93)}{.820}B)a_t$	4.28	18.10 (11)
England	$(1 - B)(1 - B^{12})r_t = (1 - \underset{(9.85)}{.682}B)(1 - \underset{(8.62)}{.673}B^{12})a_t$	6.97	5.44 (10)
France§	$(1 - B)r_t = (1 - \underset{(10.09)}{.730}B)a_t$	3.81	3.30 (11)
Germany	$(1 - B)(1 - B^{12})r_t = (1 - \underset{(10.90)}{.716}B)(1 - \underset{(11.73)}{.772}B^{12})a_t$	3.83	8.38 (10)
United States	$(1 - B)r_t = (1 - \underset{(24.95)}{.912}B)a_t$	3.02	2.67 (11)

* Figures in parentheses beneath each model are t-statistics.
† SE is the standard error of the estimated model.
‡ The reported Q-statistic, which is distributed as a χ^2, is used to test for white noise residuals. With 10 df, the 5% critical value is 18.3; with 11 df, it is 19.7.
§ Estimated sample period is January 1970–December 1977.

erties of the Eurocurrency rates and the domestic Treasury bill rate series are comparable.[5] The parameter estimates also show that the unanticipated component in the real rate always accounts for less than one-half of the observed variation in the ex post real return. The U.S. results indicate that about 9% of the unexpected component of the real rate in the last period is incorporated into the forecast for the current period. At the other extreme, the estimation results for Belgium indicate that about 34% of the unexpected component of the real rate in period $t - 1$ is incorporated into the expected rate for the current period.

Based on the model estimates found in table 4, forecasts of next period's real rate are generated. Using the U.S. estimates as an example, the ex ante real-rate forecast is given by

$$\hat{r}^t_{t-1} = (R_{t-2} - \dot{P}_{t-1}) - 0.912\, a_{t-1}. \tag{5}$$

This forecast then is used in equation (4) along with the observed nominal interest rate to produce ex ante inflation forecasts for each country.

IV. Forecast Results

The interest-rate and time-series model forecasts of inflation are compared in this section. In all cases, the model's parameter estimates are generated using data only through December 1977. Based on these

5. One reason for the slight difference may be the changing premium of the Euro-rate over the Tresury bill rate. See Mishkin (1984) for a discussion of this.

TABLE 5 Forecast Summary Statistics of In-Sample Inflation Forecasts

	Summary Statistics*			
	MAE		RMSE	
Country	Interest Rate	Time Series	Interest Rate	Time Series
Belgium	2.809	2.547	3.818	3.239
Canada	3.400	3.144	4.267	3.981
England	5.388	5.336	6.909	6.983
France	2.850	2.221	3.794	3.148
Germany	2.914	2.443	3.795	3.199
United States	2.377	2.266	3.010	3.028

NOTE.—The sample period for Belgium, Canada, and the United States runs June 1967–December 1977. For England and Germany, the period is June 1968–December 1977 due to the estimation of the seasonal component. Because of data limitations, the sample for France is January 1970–December 1977.

* MAE represents the mean absolute error; RMSE is the root mean-squared error.

estimates, out-of-sample, 1-month-ahead forecasts for the 1978–86 period are made.[6]

Before considering the out-of-sample forecasts for the different countries, it is useful to briefly examine the models' in-sample forecasting properties. To do this, summary statistics based on the in-sample forecast errors are calculated. The statistics found in table 5 indicate that the monthly in-sample forecast errors generally range from 2% to 3%, based on the mean absolute error (MAE). The performance for England stands out as the worst, and that for the United States, one of the best. The MAE statistics indicate that the in-sample time-series forecasts always are more accurate than those from the interest-rate model. The results in table 5 also indicate that, based on a root-mean-squared error (RMSE) criterion, the time-series model generally produces more accurate forecasts. For example, the interest-rate model forecasts increase the RMSE relative to the time-series model by large amounts for Belgium (18%), Canada (7%), France (20%), and Germany (19%). Only for England and the United States do we find that the interest-rate model produced more accurate in-sample inflation forecasts relative to the simple time-series model.

6. A more flexible procedure would be to allow the estimates to evolve across the forecast period. Even so, our procedure of fixing the coefficient estimates still allows the models to incorporate recent changes in inflation and nominal interest rates. As noted by Fama and Gibbons (1984), the interest-rate model uses the nominal interest rate which is allowed to vary, reflecting the market's changing perceptions about inflation. Consequently, fixing the estimation period from which the models' coefficients are based actually should give an edge in terms of forecast accuracy to the interest-rate model over the simple time-series model.

TABLE 6 Forecast Summary Statistics of Out-of-Sample Inflation Forecasts, January 1978–December 1986

	Summary Statistics*			
	MAE		RMSE	
Country	Interest Rate	Time Series	Interest Rate	Time Series
Belgium	3.516	3.214	4.631	4.167
Canada	3.385	3.191	4.356	4.009
England	4.784	4.768	6.911	6.881
France	3.938	2.350	7.057	3.012
Germany	2.261	2.207	2.942	2.912
United States	2.721	2.448	3.559	3.420

* See table 5 for details.

A. *Summary Statistics: Out-of-Sample Forecasts*

The relative accuracy of the two forecasting procedures are compared by studying one-step-ahead ex ante forecasts of monthly inflation for each country. Summary statistics for this forecasting exercise over the January 1978–December 1986 period are summarized in table 6. Comparing the mean absolute errors or RMSEs indicates that the ex ante time-series forecasts *always* are more accurate than the interest rate models' forecasts. The statistics are quite close for some countries, such as England, Germany, and the United States. In the case of France, however, the reduction in forecast error is dramatic: the RMSE based on the interest-rate model's forecast (7.06) is more than twice that from the time-series model (3.01).[7]

The results presented in table 6 indicate that the time-series forecasts are relatively more accurate than those from the interest-rate model.[8]

7. The explanation for this difference stems in part from several observations during early 1983 when France's Euro-rate rose dramatically. To see how this episode influences the interest-rate forecasts of inflation, we excluded the observations from April to July 1983 and recalculated the statistics in table 6. Now the MAE and RMSE for the interest-rate model are 4.250 and 3.124, respectively. The statistics for the time-series forecasts are 2.314 (MAE) and 2.987 (RMSE). Thus, while this period influences the quantitative results, it does not affect the qualitative conclusion that the time-series model more accurately forecasts inflation. Moreover, this result raises concerns that, because nominal interest rates reflect more than just changes in inflation expectations, forecasts of inflation from interest rates also are influenced by these factors.

8. This conclusion also is supported by subperiod results. To determine if our results were sensitive to the sample, forecast-error statistics were generated for the periods January 1978–December 1982 and January 1983–December 1986, approximately splitting the sample period. Although the relative size of the difference changes, in all instances except one we found that the RMSE for the time-series forecast errors were less than those from the interest-rate model. The only exception is for England during the 1983–86 period. In that instance, the RMSE for the time-series model was 4.29, compared with 3.97 for the interest-rate model's forecast-error series.

These results do not tell us, however, whether the two respective forecasts are statistically different. To consider this issue, a procedure outlined in Ashley, Granger, and Schmalensee (1980) was used, the null hypothesis of this test procedure being that the forecasts have equal mean-squared errors (MSE). This test was conducted for each of the six countries in our sample. The null hypothesis was rejected at the 5% level of significance for Belgium, Canada, and France. In these cases, the time-series model produces a significantly lower MSE than the interest-rate model. For the other three countries—England, Germany, and the United States—the null hypothesis was not rejected. In these three instances, there is no statistical difference.

B. Bias Test

Why does the interest-rate model do so poorly? Some insight is gained by considering another property of the forecasts, namely, their unbiasedness. This is done by estimating the regression

$$\dot{P}_t = \alpha_0 + \beta_1 \hat{\dot{P}}_t^i + \eta_t, \tag{6}$$

where $\dot{P}_t$ is the actual rate of inflation, and $\hat{\dot{P}}_t^i$ is the rate forecast by the ith model. Unbiasedness of the forecast is not rejected if the joint hypothesis that $\alpha_0 = 0$ and $\beta_1 = 1.0$ is not rejected by the data. Moreover, the error series (η_t) should be characterized by a lack of any significant serial correlation.[9] Estimating equation (6) using the forecasts for the 1978–86 period for each of the countries considered yields the relevant estimates and test statistics reported in table 7. When the interest-rate model's inflation forecasts are used, the calculated F-statistics are all significant at better than the 0.01% level of significance.[10] In every instance, rejection of the null hypothesis comes about because both the estimated constant terms are much larger than zero and the estimated slope coefficients are significantly less than unity. This indicates a tendency of the interest-rate models to overpredict inflation. Moreover, the reported Durbin-Watson (DW) statistics for France, Germany, and the United States all reject the hypothesis of no autocorrelation in the residuals at the 5% level of significance.[11]

9. Webb (1987) has argued that bias tests by themselves do not necessarily indicate whether the forecast errors follow a predictable process.

10. Given the changes in U.S. monetary policy procedures in 1979 and again in 1982, it is possible that the associated interest-rate effects both here and abroad may be reflected in the results reported in table 7. To see whether those results differ across the forecast period, the tests also were conducted using forecasts for the 1978–82 and 1983–86 subperiods. These results, while showing some qualitative changes, do not change the conclusions reached on the basis of the full sample.

11. The extremely large F-statistic for France's interest-rate model in table 7 again reveals the impact of the surge in the Euro-rate during early 1983. Deleting observations for the period April–July 1983, eq. (6) was reestimated. The estimated constant term is 3.991 (0.583), and the slope coefficient is 0.529 (0.056), both of which reject the null. The calculated F-statistic is 34.90.

TABLE 7 Results of Bias Tests of Out-of-Sample Forecasts, January 1978–December 1986

	Estimated Coefficients		Test Statistics	
	α_0	β_1	D-W	*F*
Interest-rate model:				
Belgium	3.143	.403	1.71	10.25
	(.839)	(.132)		(.00)
Canada	2.695	.602	2.05	8.309
	(.841)	(.099)		(.00)
England	2.824	.659	1.75	9.580
	(.884)	(.078)		(.00)
France	6.094	.288	1.15*	138.610
	(.518)	(.043)		(.00)
Germany	1.018	.680	1.56*	8.970
	(.363)	(.076)		(.00)
United States	1.373	.726	1.20*	11.463
	(.528)	(.061)		(.00)
Time-series model:				
Belgium	1.607	.681	1.88	2.050
	(.972)	(.161)		(.13)
Canada	.526	.893	2.11	.551
	(1.108)	(.139)		(.58)
England	2.743	.682	1.74	7.329
	(.913)	(.083)		(.00)
France	.420	.936	1.82	.398
	(.785)	(.084)		(.67)
Germany	.790	.738	1.47*	4.497
	(.399)	(.088)		(.01)
United States	.273	.937	1.19*	.342
	(.662)	(.088)		(.71)

NOTE.—Coefficients from estimating equation (6); Standard errors are reported in parentheses. The *F*-statistic is distributed with (2,106) degrees of freedom. The null hypothesis is $\hat{\alpha}_0 = 0$ and $\hat{\beta}_1 = 1.0$.
* Autocorrelation at the 5% level for the Durbin-Watson (D-W) statistic.

In contrast to the evidence from the interest-rate model forecasts, unbiasedness is rejected only for the time-series forecasts in England and Germany.[12] As the estimates in table 7 show, both the constant terms are greater than zero and the slope coefficients significantly less than unity for these two countries. It should be noted, however, that even though we cannot reject the joint hypothesis that $\alpha_0 = 0$ and $\beta_1 = 1$ using the U.S. data, the reported Durbin-Watson statistic (1.19) indicates positive serial correlation at the 5% level of significance. Consequently, the hypothesis of unbiasedness also is doubtful for the time-series model forecasts of U.S. monthly inflation.

C. *Forecast Combinations*

It is possible to compare the forecasts' relative informational content by combining the forecasts from each model. This is suggested by

12. We again examined the subperiods and found that unbiasedness is rejected in both.

TABLE 8 **Inflation Forecast Results, Individual and Combined Models, Postsample Weighting, January 1978–December 1986**

	RMSE			Weights[a]		
Country	Interest Rate	Time Series	Combined	Constant	Interest Rate	Time Series
Belgium	4.631	4.167	4.132	1.53	−.20	.90*
Canada	4.356	4.009	4.016	.92	.24	.61*
England	6.911	6.881	6.450	2.81*	.04	.03
France	7.057	3.012	2.876	.62	.13*	.78*
Germany	2.942	2.912	2.738	.84*	.48*	.25
United States	3.559	3.420	3.260	1.01	.59*	.20

[a] Weights from estimating eq. (7) using postsample forecasts.
* Significant at the 5% level.

Granger and Ramanathan (1984) as one approach to obtain forecasts that incorporate information from each model. A similar approach is discussed in Lupoletti and Webb (1986).

Using each model's forecasts, we estimate the regression

$$\dot{P}_t = \alpha_0 + \beta_1 \dot{P}_t^{IR} + \beta_2 \dot{P}_t^{TS} + \epsilon_t, \tag{7}$$

where $\dot{P}$ is the actual rate of inflation, $\dot{P}_t^{IR}$ represents the inflation forecast from the relevant interest-rate model, and $\dot{P}_t^{TS}$ represents the inflation forecast calculated from the time-series model. The relative significance of the estimated coefficients provides some information on the "marginal explanatory power" of each forecast in the presence of the other. If the estimated coefficient on the time-series model forecast became insignificant with the inclusion of the interest-rate model forecast, then one could argue that the time-series forecast contains no marginally useful information over and above that already embedded in the interest rate. Of course, an opposite result would suggest that forecasts of inflation embedded in nominal rates do not efficiently utilize information in past information rates.

Equation (7) was estimated using the out-of-sample forecasts from the two models for each country. The time period is January 1978–December 1986. The outcome is presented in table 8. For convenience, we report the forecast RMSE for the two individual models, the "combined"model—that is, equation (7)—and the weights (the estimated β's) assigned to each forecast. The forecasts from the combined models indicate that the improvement in forecasting accuracy is greater relative to the interest-rate forecasts than the time-series model. Even after eliminating the 59% improvement in the combined forecast for France relative to the interest-rate model, the average improvement in the forecast RMSE is 8% when compared to the interest-rate model and 4% when compared to the time-series model forecasts.

TABLE 9 Inflation Forecast Results, Individual and Combined Models, In-Sample Weighting, January 1978–December 1985

	Weights[a]			
Country	Constant	Interest Rate	Time Series	RMSE
Belgium	.93	.15	.72*	4.158
Canada	1.31	−.07	.90*	4.026
England	2.39*	.37	.42*	6.460
France	2.36*	.24*	.49*	3.014
Germany	1.78*	.02	.63*	2.887
United States	1.21	.50*	.32	3.233

[a] Weights from estimating eq. (7) using in-sample model forecasts through December 1977.
* Significant at the 5% level.

An interesting aspect of table 8 is the reported weights for each model. The weights for the two forecasts using the U.S. data indicate that forecasts from the interest-rate model are weighted relatively more than those from the time-series model. The evidence for Germany also indicates that the weight on the interest-rate forecast (0.48) is significantly different from zero and much larger than that on the time-series forecast (0.25). For Belgium and Canada, however, the evidence suggests that the interest-rate model adds little information to that already contained in the time-series forecasts of inflation. The results for France show that the time-series model's forecasts receive a greater weight than the interest-rate forecast, although both contain marginally significant information. Finally, the results for England suggest that neither series significantly improves on the accuracy of the other.

Granger and Ramanathan (1984) suggest that a more competitive evaluation of the two models is to compare the forecasts from equation (7), where the weights are the coefficients found in estimating equation (7) using the *in-sample* inflation forecasts available through December 1977. The summary RMSEs generated by this procedure, along with the weights are reported in table 9.

Note the shift in weights in table 9. Estimates of equation (7) using this weighting procedure generally give insignificant weight to the inflation forecasts from the interest-rate model. In fact, only the weight for France and the United States is significant at the 5% level. The results in table 9 show that, in four of the six countries, combining the interest-rate and time-series inflation forecasts using the pre-1978 weights led to a reduction in the RMSE generated by either model individually. Only for Canada and France do we find that the combined model's forecast RMSE is larger than the time-series model's RMSE reported in table 8. When one compares the RMSEs in table 9 to the time-series models' RMSEs in table 8, the time-series model yields forecasts that are essentially as accurate as the combined model for Belgium, Canada, France, and Germany. Only for England and the United States is the combined models' RMSE less than that from the time-series model.

The evidence in tables 8 and 9 indicates that combining the individual model forecasts generally leads to an improvement in the forecast accuracy. This result, and the fact that the weights generally are not zero, suggests that the best model to forecast inflation is neither the time-series approach nor the interest-rate procedure. In relative terms, however, the marginal improvement in forecasting accuracy clearly is greater when the time-series model forecasts are combined with the interest-rate model forecasts, and not vice versa.[13]

V. Conclusion

What have we learned from the evidence presented? One important result is that, based on data from several countries, inflation forecasts generated from observed nominal interest rates do not dominate those from a univariate time-series model. This conclusion contrasts with that reached by Fama and Gibbons (1984), who focused solely on U.S. data. Based on the evidence from Belgium, Canada, England, France, and Germany, we find that time-series forecasts of inflation have equal or lower forecast error and produce unbiased forecasts more often than the interest-rate model. The interest-rate model, it should be noted, generally overpredicted inflation for the 1978–86 period for all countries studied.

The other important finding from this study is that the interest-rate model forecasts may provide marginally useful information that allows one to improve on the time-series inflation forecasts. This is true for France, Germany, and the United States. This result suggests that the best inflation forecast is one that combines the information inherent in both the time-series and interest-rate models.

References

Ashley, R.; Granger, C. W. J.; and Schmalensee, R. 1980. Advertising and aggregate consumption: An analysis of causality. *Econometrica* 48 (July): 1149–67.

Box, G. E. P., and Jenkins, G. M. 1976. *Time Series Analysis: Forecasting and Control.* San Francisco: Holden-Day.

Brown, B. W., and Maital, S. 1981. What do economists know? An empirical study of experts' expectations. *Econometrica* 49 (March): 491–504.

13. The performance of the time-series model relative to the interest-rate model may reflect the influence of varying default premiums in the interest-rate series. While this is a possibility, the results for the United States are very similar to those found previously using the 1-month Treasury-bill rate. An avenue of future research is to investigate the sensitivity of the results based on alternative interest-rate series. See Mishkin (1984) for a related discussion.

Carlson, J. A. 1977. Short-term interest rates as predictors of inflation: Comment. *American Economic Review* 67 (June): 469–75.
Cumby, R. E., and Mishkin, F. S. 1986. The international linkage of real interest rates: The European-U.S. connection. *Journal of International Money and Finance* 5 (March): 5–23.
Dickey, D. A., and Fuller, W. A. 1979. Distribution of the estimators for autoregressive time series with a unit root. *Journal of the American Statistical Association* (June), pp. 427–31.
Fama, E. F. 1975. Short-term interest rates as predictors of inflation. *American Economic Review* 65 (June): 269–82.
Fama, E. F. 1976. Inflation uncertainty and expected returns on Treasury bills. *Journal of Political Economy* 84 (June): 427–48.
Fama, E. F. 1977. Interest rates and inflation: The message in the entrails. *American Economic Review* 67 (June): 487–96.
Fama, E. F., and Gibbons, M. R. 1982. Inflation, real returns, and capital investment. *Journal of Monetary Economics* 9 (May): 297–324.
Fama, E. F., and Gibbons, M. R. 1984. A comparison of inflation forecasts. *Journal of Monetary Economics* 13 (May): 327–48.
Garbade, K., and Wachtel, P. 1978. Time variation in the relationship between inflation and interest rates. *Journal of Monetary Economics* 4 (November): 755–65.
Granger, C. W. J., and Ramanathan, R. 1984. Improved methods of combining forecasts. *Journal of Forecasting* 3 (April–June): 197–204.
Hafer, R. W., and Hein, S. E. 1985. On the accuracy of time-series, interest rate, and survey forecasts of inflation. *Journal of Business* 58 (October): 377–98.
Hein, S. E., and Koch, T. W. 1988. Interest rates, taxes and short-term inflation forecasts. Mimeographed. Lubbock: Texas Tech University, May.
Hess, P. J., and Bicksler, J. L. 1975. Capital asset prices versus time series models as predictors of inflation: The expected real rate of interest and market efficiency. *Journal of Financial Economics* 2 (December): 341–60.
Kane, A., and Rosenthal, L. 1982. International interest rates and inflationary expectations. *Journal of International Money and Finance* 1:97–110.
Lupoletti, W. M., and Webb, R. H. 1986. Defining and improving the accuracy of macroeconomic forecasts: Contributions from a VAR model. *Journal of Business* 59 (April): 263–85.
Mark, N. C. 1985. Some evidence on the international inequality of real interest rates. *Journal of International Money and Finance* 4:189–208.
Mishkin, F. S. 1984. The real interest rate: A multi-country study. *Canadian Journal of Economics* 17 (May): 283–311.
Nelson, C. R., and Schwert, G. W. 1977. Short-term interest rates as predictors of inflation: On testing the hypothesis that the real rate of interest is constant. *American Economic Review* 67 (June): 478–86.
Pearce, D. K. 1979. Comparing survey and rational measures of expected inflation: Forecast performance and interest rate effects. *Journal of Money, Credit, and Banking* 11 (November): 447–56.
Sims, C. A. 1988. Bayesian skepticism on unit root econometrics. Discussion paper no. 3 (May). Minneapolis: Institute for Empirical Macroeconomics, Federal Reserve Bank of Minneapolis.
Webb, R. H. 1987. The irrelevance of tests for bias in series of macroeconomic forecasts. *Federal Reserve Bank of Richmond Economic Review* 73 (November/December): 3–9.
Whitt, J. A., Jr. 1988. The long-run behavior of the real exchange rate: A reconsideration. Unpublished manuscript. Atlanta: Federal Reserve Bank of Atlanta (October).