Table 1—Sample Autocorrelations of One-Month *Ex Post* Real Rate, January 1953–June 1971

Lag	Autocorrelation	Lag	Autocorrelation
1	.10	7	−.08
2	.12	8	.04
3	−.02	9	.10
4	−.01	10	.09
5	−.02	11	.03
6	−.01	12	.18

Note: Standard Error = .07; Box-Pierce $Q = 19.4$ (χ^2 with 12 degrees of freedom).

Suppose that the *EARR* rather than being constant is a stochastic process with first-order serial correlation coefficient ϕ. Denoting the market's forecasting error for the rate of inflation by ϵ_t, the *EARR* by i_t, and the EPRR by r_t, we have

$$r_t \equiv i_t - \epsilon_t \tag{1}$$

It is easy to show that in an efficient market the first-order serial correlation coefficient for the *EPRR* will be related to ϕ by

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

where σ_i^2 and σ_ϵ^2 denote the variances of the *EARR* and the forecast error, respectively. It should be clear that first-order autocorrelation in the *EPRR* may be considerably less than ϕ if the variance of forecast errors is large relative to the variance of the *EARR*.

TABLE 2—VALUES OF ϕ AND σ_w^2 AND CORRESPONDING VARIANCES AND STANDARD DEVIATIONS FOR THE *EARR* AS AN ANNUAL PERCENTAGE RATE CONSISTENT WITH $\mathrm{Corr}(r_t, r_{t+1}) = .10$ and $\sigma_r^2 = 4.32$

ϕ	Variance of *EARR*, σ_i^2	Standard Deviation of *EARR*, σ_i	σ_w^2
.4	1.44	1.20	1.21
.5	1.08	1.04	.806
.6	.864	.924	.547
.7	.720	.852	.374
.8	.619	.780	.230
.9	.547	.732	.102
.95	.508	.708	.049
.99	.485	.696	.010
.999	.481	.696	.001

Note: The statistic σ_w^2 is computed under the assumption: $i_t = \phi i_{t-1} + w_t$.

The autocorrelation structure of the *CPI* monthly inflation series for January 1953 through July 1971 suggests that the series may reasonably be represented as a first-order moving average process in its first differences, implying nonstationary behavior in the rate of inflation. The estimated model for the 2/53–7/71 period is

$$(4) \qquad (1 - B)\rho_t = \underset{(.0179)}{.0222} + (1 - \underset{(.029)}{.894}B)e_t$$

$$\hat{\sigma}_e = 2.408$$

where e_t is a sequence of residuals which are the one-step-ahead forecast errors for this model and B is the lag operator. The forecast $\hat{\rho}_t$ of ρ_t implied by this model may be written apart from a constant as

$$(5) \qquad \hat{\rho}_t = .11\rho_{t-1} + (.89)(.11)\rho_{t-2} + \dots + (.89)^j(.11)\rho_{t-j-1} + \dots = \sum_{j=0}^{\infty} \theta^j(1 - \theta)\rho_{t-j-1}$$

where $\theta = .89$. Note that the weight given to ρ_{t-1} in this forecast is very small; interestingly, it is the same as the regression coefficient of ρ_{t-1} in Fama's regressions of ρ_t on R_t and ρ_{t-1} for monthly data.[11] Also

TABLE 3

α	β	γ	$\hat{\sigma}_e$	R^2	$D.W.$
A. Composite Predictors of the Rate of Inflation: 2/53–7/71					
−.775	.969		2.347	.292	1.81
(.358)	(.102)				
−.641	.651	.383	2.322	.310	1.93
(.359)	(.165)	(.158)			
B. Composite Predictors of the Change in the Rate of Inflation: 2/53–7/71					
−.774	.889		2.333	.458	2.07
(.167)	(.065)				
−.546	.633	.317	2.320	.466	2.03
(.206)	(.152)	(.170)			

Note: The prediction equations are:

A. $\rho_t = \alpha + \beta R_t + \gamma \hat{\rho}_t + e_t$

B. $(\rho_t - \rho_{t-1}) = \alpha + \beta(R_t - \rho_{t-1}) + \gamma(\hat{\rho}_t - \rho_{t-1}) + e_t$

(Standard errors in parentheses.)

tive regressions were too small to alter our basic conclusions. Ideally, the composite weights should be estimated simultaneously with the parameters of the time-series model by specifying a dynamic regression (transfer function) model which includes pure interest rate and pure time-series prediction as special cases. The resulting point estimates and their standard errors would reflect appropriately the information contained in the data. Taking this approach we embedded the interest rate in the time-series model for inflation and obtained the following results for the 2/53–7/71 period:

$$(6) \quad (1 - B)\rho_t = \underset{(.0211)}{.0199} + \underset{(.262)}{.577}(1 - B)R_t + (1 - \underset{(.032)}{.876}B)e_t$$

$$\hat{\sigma}_e = 2.430$$

If we are willing to assume further that the covariance between ρ^* and i is zero (no "Mundell Effect"), then the covariance between R and ρ is simply the variance of ρ^* and thus

$$(8) \qquad \hat{\sigma}_i^2 = \hat{\sigma}_R^2 - \hat{\sigma}_{R\rho}$$
$$\hat{\sigma}_\epsilon^2 = \hat{\sigma}_\rho^2 - \hat{\sigma}_{R\rho}$$

The same relations hold with regard to variances and covariances of changes in the rate of inflation $(\rho_t - \rho_{t-1})$ and predicted changes in the rate of inflation $(R_t - \rho_{t-1})$ and $(\rho_t^* - \rho_{t-1})$. Since there is some doubt about the stationarity of ρ_t and R_t and therefore about the existence of variances and covariances, computations in terms of changes in the rate of inflation and $(R_t - \rho_{t-1})$ may be preferable. Sample moments for these variables from the data used by Fama imply

$$(9) \qquad \hat{\sigma}_i^2 = .642$$
$$\hat{\sigma}_\epsilon^2 = 4.85$$

It is interesting to note that this estimate of σ_i^2, the variance of the *EARR*, is quite close to those presented in Table 2 which are associated with a fairly strongly autocorrelated *EARR*.