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The Time-Varying-Parameter Model for Modeling Changing Conditional Variance: The Case of the Lucas Hypothesis

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The main econometric issue in testing the Lucas (1973) hypothesis in a time series context is estimation of the forecast-error variance conditional on past information. The conditional variance may vary through time as monetary policy evolves and agents are obliged to infer its present state. Under the assumption that a monetary policy regime is continuously changing, a time-varying-parameter model is proposed for the monetary-growth function. Based on Kalman-filtering estimation of recursive forecast errors and their conditional variances, the Lucas hypothesis is tested for the U.S. economy (1964:1–1985:4) using monetary growth as aggregate demand variable. The Lucas hypothesis is rejected in favor of Friedman's (1977) hypothesis—the conditional variance of monetary growth affects real output directly, not through the coefficients on the forecast-error term in the Lucas-type output equation.

KEY WORDS: Conditional variance; Kalman filter; Lucas hypothesis.

A person's uncertainty about the future arises not simply because of future random terms but also because of uncertainty about current parameter values and of the model's ability to link the present to the future. (Harrison and Stevens 1976, p. 208)

1. INTRODUCTION

The Lucas (1973) hypothesis predicts a negative relationship between the variance of nominal shocks and the magnitude of the output response to nominal shocks. Lucas (1973), Froyen and Waud (1980), Alberro (1981), Kormendi and Meguire (1984), and others have examined the Lucas hypothesis using cross-country data, assuming that policy regimes did not change within countries, different policy regimes being represented by different countries. The assumption of a constant variance of nominal shocks within a country over time does not seem realistic, however. The variance conditional on information available at the time of forecasting may be time-varying due, for example, to a continuously changing policy regime that is represented by evolutionary regression coefficients. This conditional variance, rather than the unconditional variance that is based on the whole sample, is what really matters for the behavior of economic agents, as Engle (1982) pointed out. A more powerful test of the Lucas hypothesis may, therefore, result from modeling variation in conditional variance through time.

This article discusses some of the implications of a continuously changing monetary-policy regime (of the U.S. Federal Reserve) for tests of the Lucas hypothesis and proposes a test in the time series context using money growth ($M1$) as an aggregate demand variable. The key issue involved in testing the Lucas hypothesis in the time series context is that of estimating the conditional variance that changes over time. Calculating a moving variance based on several past observations (e.g., Froyen and Waud 1984; Lawrence 1983) is one way of capturing the changing conditional variance of time series data, and the autoregressive conditional heteroscedasticity (ARCH) modeling introduced by Engle (1982) is another way, but neither method specifies the source of changing conditional variance.

Kalman filtering, which was first introduced in the engineering literature, is applied here to estimate the time-varying coefficients of the time-varying-parameter (TVP) model. [See Anderson and Moore (1979), Harvey (1981), Los (1984), and Rosenberg (1973) for the discussion of Kalman filtering and the TVP model.] As by-products of the Kalman-filtering estimation of the TVP model, recursive forecast errors and their conditional variances that can be used to test the Lucas hypothesis are obtained. A nice thing about Kalman filtering is that it gives us insight into how a rational economic agent would revise his estimates of the coefficients of the model in a Bayesian fashion when new information is available, especially under a changing policy regime. Benjamin Friedman (1979) argued that "what is typically missing in rational expectation mechanisms is a clear outline of the way in which economic agents derive their knowledge which they use to formulate expectations" (p. 24). By referring to the informational availability assumption of the rational-expectations hypothesis, he proposed least-squares learning as an optimal learning process. Under a continuously changing policy regime, however, the least-squares learning is no longer optimal. In this circumstance, an application of the Kalman filter is an appropriate way of proxying the rational agents' learning process.

An alternative to the Lucas hypothesis is Milton Friedman's (1977) hypothesis, which states that the increased variability of the inflation rate causes a reduction in the allocative efficiency of the price system, causing a reduction in natural level of output. In money-oriented Phillips-curve models such as that of Barro (1976), unanticipated inflation is posited as an intermediate link in the causal chain connecting unanticipated money growth to real variables in the system, so the conditional variance of monetary forecast errors is used as a proxy for the variability of inflation rate. By allowing the natural level of output to depend on the conditional variance of the monetary-forecast error, Milton Friedman's hypothesis is also tested in this article.

The organization of the article is as follows. In Section 2, a monetary-growth function based on the stability-test results of the regression coefficients is specified and estimated, assuming that the regression coefficients follow random walks. In Section 3, empirical tests of the Lucas hypothesis and Friedman's hypothesis are performed based on the Kalman-filtering estimation of forecast errors and their conditional variance from Section 2. Both two-step and joint estimation procedures are considered. Section 4 concludes.

2. STABILITY TESTS AND KALMAN-FILTERING ESTIMATION OF THE TVP MONETARY-GROWTH FUNCTION

McNees (1986) stated, "A policy reaction function is likely to be a fragile creature. Over time, . . . the importance attached to conflicting objectives [of the policy] may change, [policy makers'] views on the structure of the economy may change" (p. 7). In this section, the TVP model is applied to a monetary-growth function. Stability tests for the regression coefficients provide a rationale for TVP modeling.

Two different stability tests were performed on the monetary-growth function to be specified in (2.1), one against the alternative hypothesis of *unstable regression coefficients* and the other against the alternative hypothesis of *random-walk coefficients*. The former test is based on a homogeneity test (or Chow test), as proposed by Brown, Durbin, and Evans (1975). By splitting the entire sample into some arbitrary nonoverlapping subsamples, the between-group-over-within-groups ratio of mean squares was calculated. Under the null hypothesis of stable coefficients, the test statistic is distributed as $F(kp - k, T - kp)$, where k is number of regressors, p is number of nonoverlapping subsamples, and T is the whole sample size. Test results rejected the null hypothesis in favor of unstable coefficients at the 5% significance level in all cases.

The actual structural form of the time-varying regression coefficients, which have rejected the stability of regression coefficients, is of further interest. Engle and Watson (1985) suggested unit roots for the coefficients in cases of structural change in which agents adjust their estimation of the state only when new information becomes available. Therefore, a stability test can be performed against the alternative hypothesis that the coefficients follow random walks.

Under the alternative hypothesis of random-walk coefficients, residuals from an ordinary least squares (OLS) regression have a particular heteroscedastic form, which depends on $t * x_t^2$ (x_t is a vector of regressors). Following Breusch and Pagan (1979), one half times the explained sum of squares from a regression of $\hat{u}_t^2/\sigma_u^2$ on $t * x_t^2$, where $\hat{u}_t^2$ is the vector of residuals from an OLS regression of, say, (2.1), is distributed as $\chi^2(k)$ under the null hypothesis of stable coefficients. The test results show that the null hypothesis of stable coefficients is rejected for coefficients on $DINT_{t-1}$, INF_{t-1}, and the intercept term, and the joint hypothesis that all of the coefficients are stable (against the alternative hypothesis of random walks) is rejected at the 1% significance level.

Based on the preceding stability-test results, we assume that each of the regression coefficients of the model follows a random walk. The TVP model for a monetary-growth function with estimated innovation variances for U.S. quarterly data 1964:1–1985:4 is

$$\begin{aligned}
DM1_t &= \beta_{0t} + \beta_{1t}DINT_{t-1} + \beta_{2t}INF_{t-1} \\
&\quad + \beta_{3t}SURP_{t-1} + \beta_{4t}DM1_{t-1} + e_t \\
\beta_{it} &= \beta_{it-1} + v_{it}, \qquad i = 0, 1, \ldots, 4 \\
\sigma_e^2 &= .126284 \\
\sigma_{v0}^2 &= .012133 \\
\sigma_{v1}^2 &= .000896 \\
\sigma_{v2}^2 &= .074544 \\
\sigma_{v3}^2 &= .000683 \\
\sigma_{v4}^2 &= .001184,
\end{aligned} \tag{2.1}$$

where $DM1$, $DINT$, INF, and $SURP$ stand for the quarterly $M1$ growth rate, the change in the interest rates on three-month treasury bills, inflation measured by the Consumer Price Index (CPI), and the full-employment budget surplus, respectively. A description of the data and their sources can be found in the Appendix. The specification of the model (i.e., the choice of the right-side variables) is motivated by Mishkin (1982) and Weintraub (1980). The OLS regression results of the model are very similar to theirs, except that a lagged inflation term is added in our model.

In estimating the preceding non-time-varying parameters of the model, the likelihood value obtained from the Kalman-filter algorithm was maximized using the scoring method proposed by Engle and Watson (1981) and Watson and Engle (1983). A nice feature of the scoring method is that we need only the first derivatives. The computer program was written in the GAUSS programming language, and the numerical derivatives were used, as proposed by Engle and Watson (1981) and Watson and Engle (1983).

Given estimates of the innovation variances, the next step is to estimate the evolutionary coefficients of the model based on past information ($\beta_{it/t-1}$, $i = 0, 1, \ldots, 4$). At this step, the Kalman filter is run again with the preceding estimates of σ_e^2 and σ_{vi}^2's for given initial values of $\beta_{i0/-1}$ ($i = 0, \ldots, 4$) and their variance–covariance matrix. (These initial values were estimated at the expense of the first 16 observations.) As by-products of running the Kalman-filtering algorithm, the one-step-ahead forecast errors [$\eta_{t/t-1} = DM1_t - x_t\beta_{t/t-1}$, where x_t is vector of regressors in (2.1)] and their conditional variances ($H_{t/t-1}$) can be estimated. The graphs showing the estimates of the evolutionary coefficients of the model, one-step-ahead forecast errors, and their conditional variances are available from us on request.

Once we estimate the model, we need to see whether the model is correctly specified. One useful way of checking the appropriateness of the specified model is to check for whiteness or lack of serial correlation in the one-period-ahead forecast errors ($\eta_{t/t-1}$), as suggested by Engle and Watson (1981). The null hypothesis of no serial correlation in the forecast errors was not rejected at the 5% significance level, which shows the appropriateness of the model specification. Box–Pierce test statistics for the heteroscedasticity-adjusted one-period-ahead forecast errors ($H_{t/t-1}^{-1/2}\, \eta_{t/t-1}$) were $Q(12) = 10.1$, $Q(24) = 21.9$, and $Q(36) = 29.6$.

When ARCH tests were performed based on a fixed-coefficient version of the monetary-growth function, very strong evidence of an ARCH effect was found. To determine if the source of the ARCH effect could be varying coefficients of the model, we checked to see if the serial correlation still remains in the squared forecast-error terms (from the TVP specification) after adjusting them for the conditional heteroscedasticity ($H_{t/t-1}$) as implied by the TVP modeling. If we find any serial correlation in the squares of heteroscedasticity-adjusted forecast-error terms ($H_{t/t-1}^{-1/2}\eta_{t/t-1}$) from the TVP specification, we may suspect that the ARCH effect in the OLS regression may be due to reasons other than evolutionary coefficients of the model, but we could not reject the null hypothesis that there is no serial correlation in the $(H_{t/t-1}^{-1/2}\eta_{t/t-1})^2$ terms, implying that the existence of the ARCH in the monetary-growth function could be due to the changing regression coefficients of the model. The test statistics are available from us.

It is of interest that ARCH and TVP's can produce measures of conditional variances that often move together. Figure 1 compares the conditional variance of the forecast-error terms from the two specifications. Note that the TVP estimation results show higher uncertainty about future monetary policy during the oil-shock period around 1974, but ARCH conditional variance rises less. In the ARCH model, insofar as forecast errors in the near past are small, implied uncertainty about the future is small. In the TVP framework, however, conditional uncertainty is directly related to observed variables that contain information about current monetary policy. This is apparent from the following equation for the conditional variance of forecast errors from the Kalman-filtering algorithm:

$$H_{t/t-1} = x_t P_{t/t-1} x'_t + \sigma_e^2, \tag{2.2}$$

where $P_{t/t-1}$ is the variance–covariance matrix of $\beta_{t/t-1}$, σ_e^2 is the variance of the disturbance term e_t, and x_t is the vector of regressors in (2.1). This equation tells us that, under a continuously changing policy regime, there are two sources of uncertainty, uncertainty that arises because of future disturbances and uncertainty that arises because of evolutionary regression coefficients. In 1974, unusually large movements in the inflation rate caused the conditional variance to be unusually large.

3. EMPIRICAL TESTS OF THE LUCAS HYPOTHESIS AND FRIEDMAN'S HYPOTHESIS

This section presents empirical results from a Lucas-type reduced-form output equation, which also allows conditional uncertainty to have direct effect on output, as hypothesized by Friedman. To test Friedman's (1977) hypothesis as well as the Lucas hypothesis, we assume that the natural level of output is affected by conditional variance of a nominal shock, a proxy for inflation uncertainty; that is, following Froyen and Waud (1984), we assume that

$$y_t = y_{nt} + y_{ct} \tag{3.1}$$

and

$$y_{nt} = \delta_0 + \delta_1 t + \delta_2 H_{t/t-1}, \tag{3.2}$$

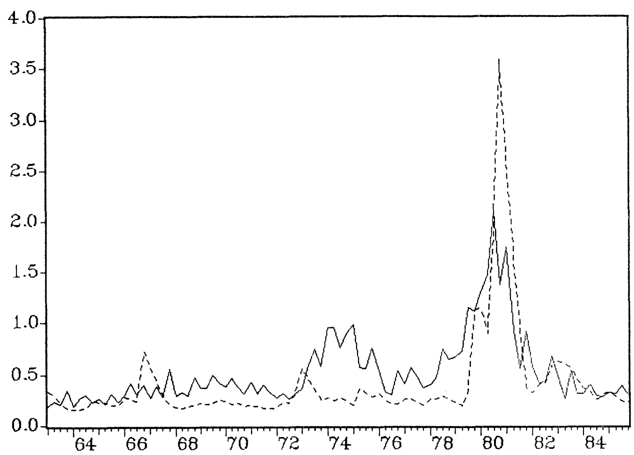

Figure 1. Conditional Variances of Forecast Errors: ——, TVP; ---, ARCH.

Table 1. Test of the Lucas Hypothesis (level), Uncorrected for Serial Correlation

Estimation	α_0	γ_0	γ_1	α_2	α_3	R^2	DH
Two-step	.00053	.0026	—	—	.954	.9188	3.29
	(.485)	(2.049)			(31.944)		
	.00053	.0026	−.00009	—	.954	.9198	3.28
	(.418)	(2.030)	(−.009)		(31.761)		
	.0069	.0035	—	−.013	.967	.9334	2.20
	(3.823)	(2.962)		(−4.23)	(35.098)		
	.00798	.0038	.0027	−.014	.968	.9355	2.20
	(4.200)	(3.199)	(1.67)	(−4.59)	(35.489)		
Joint	.0070	.0042	.0032	−.013	.966		
	(3.100)	(2.961)	(1.75)	(−2.62)	(34.49)		

NOTE: t values are in parentheses. $cy_t = \alpha_0 + (\gamma_0 + \gamma_1 \ln H_{t/t-1})\eta_{t/t-1} + \alpha_2 H_{t/t-1} + \alpha_3 cy_{t-1} + u_t$, cy_t = time-detrended real output, $\eta_{t/t-1}$ = recursive forecast error, and $H_{t/t-1}$ = conditional forecast-error variance.

where y_t, y_{nt}, and y_{ct} represent the log of real output, natural output, and cyclical output, respectively. Therefore, the time-detrended output cy_t is the sum of cyclical output and the portion of natural output that is dependent on the inflation uncertainty. The recursive forecast errors ($\eta_{t/t-1}$) and the conditional variance ($H_{t/t-1}$) estimated in Section 2 are used to test the Lucas and Friedman hypotheses (a two-step procedure). Thus the output equation to be tested is reduced to

$$cy_t = \alpha_0 + \alpha_{1t}\eta_{t/t-1} + \alpha_2 H_{t/t-1} + \alpha_3 cy_{t-1} + \mu_t \quad (3.3)$$

and

$$\alpha_{1t} = f(H_{t/t-1}), \quad (3.4)$$

where cy_t is time-detrended output, $\eta_{t/t-1}$ is a recursive forecast error in monetary growth, and $H_{t/t-1}$ is the conditional variance of the forecast error. The functional form of the time-varying coefficient α_{1t} in Equation (3.4) was chosen in such a way as to minimize the multicolinearity problem in OLS regression of Equation (3.3). Thus the specification $\alpha_{1t} = \gamma_0 + \gamma_1 \ln(H_{t/t-1})$ was chosen. Testing the Lucas hypothesis is equivalent to testing $\gamma_0 > 0$ and $\gamma_1 < 0$, and testing Friedman's hypothesis is equivalent to testing $\alpha_2 < 0$. Test results based on the OLS regression of Equation (3.3) are reported in Tables 1–4.

In most of the regressions, γ_0 is positive and significantly different from 0. The estimated values of γ_1, on the contrary, have the wrong signs and are not significantly different from 0. [When a $\eta_{t/t-1} * \ln(H_{t/t-1})$ term is added to the regression, there is little improvement in R^2.] So the Lucas hypothesis is not supported by the data, but when the conditional variance term ($H_{t/t-1}$) is included directly in the regression equation, the R^2 increased significantly, and its coefficient (α_2) is significantly different from 0 with a negative value. This suggests that the conditional variance of forecast error affects the output level directly and not through the coefficient on the forecast error ($\eta_{t/t-1}$) term.

Coefficients on lagged detrended output in Tables 1 and 2 are close to 1 (range from .912 to .968). This could be due to downward bias in the coefficients when there is a unit root, as explained by Dickey and Fuller (1979), so the regressions were performed by using the first difference of detrended output as a dependent variable (Tables 3 and 4). The test results were not significantly different from those using detrended output in levels. The R^2's from regression using differenced data range from .1456 to .2775 after correcting for serial correlation. This suggests that monetary shocks and their conditional variances alone cannot provide a complete picture of short-term business fluctuations.

Mishkin (1982) and Pagan (1984) pointed out problems associated with a two-step procedure; that is, the regressors in the output equation, $\eta_{t/t-1}$ and $H_{t/t-1}$, are generated regressors, and this might influence the distribution of test statistics, invalidating the inference made previously. Consideration was given to this possibility. Therefore, joint estimation of the TVP monetary-growth function with the real-output equation was carried out.

Table 2. Test of the Lucas Hypothesis (level), Corrected for Serial Correlation

Estimation	α_0	γ_0	γ_1	α_2	α_3	R^2	DH
Two-step	.00009	.0027	—	—	.912	.9292	−.520
	(.538)	(2.592)			(18.022)		
	.00087	.0026	.00056	—	.910	.9293	−.546
	(.547)	(2.559)	(.39)		(17.821)		
	.00674	.0036	—	−.012	.949	.9368	−.546
	(3.208)	(3.425)		(−3.56)	(25.623)		
	.00798	.0037	.0025	−.014	.951	.9355	−.112
	(3.641)	(3.575)	(1.72)	(−4.01)	(26.240)		
Joint	.0074	.0040	.0028	−.014	.947		
	(2.785)	(3.058)	(1.65)	(−2.55)	(25.30)		

NOTE: t values are in parentheses. $cy_t = \alpha_0 + (\gamma_0 + \gamma_1 \ln H_{t/t-1})\eta_{t/t-1} + \alpha_2 H_{t/t-1} + \alpha_3 cy_{t-1} + u_t$, cy_t = time-detrended real output, $\eta_{t/t-1}$ = recursive forecast error, and $H_{t/t-1}$ = conditional forecast-error variance.

Table 3. Test of the Lucas Hypothesis (differenced output), Uncorrected for Serial Correlation

Estimation	α_0	γ_0	γ_1	α_2	R^2	DW
Two-step	.00015	.00276	—	—	.0502	1.3649
	(.137)	(2.179)				
	.00014	.00276	−.00013	—	.0502	1.3673
	(.134)	(2.158)	(−.078)			
	.00684	.00364	—	−.0131	.2186	1.5788
	(3.778)	(3.102)		(−4.38)		
	.00793	.00392	.00277	−.0151	.2436	1.5750
	(4.169)	(3.342)	(1.70)	(−4.74)		
Joint	.00719	.00414	.00284	−.0127		
	(2.982)	(2.911)	(1.560)	(−2.620)		

NOTE: t values are in parentheses. $Dcy_t = \alpha_0 + (\gamma_0 + \gamma_1 \ln H_{t|t-1})\eta_{t|t-1} + \alpha_2 H_{t|t-1} + u_t$, $Dcy_t = cy_t - cy_{t-1}$, cy_t = time-detrended real output, $\eta_{t|t-1}$ = recursive forecast error, and $H_{t|t-1}$ = conditional forecast-error variance.

Estimates from the two-step procedure were used as initial values in the joint estimation. Test results based on the joint estimation are reported on the last rows of Tables 1–4. Point estimates were not very different from those from two-step estimation, and standard errors of the estimates were larger than those from two-step estimation. This is because the joint estimation procedure does not ignore the nonzero off-diagonal elements in the information matrix of the joint estimates. Although the t values are lower in general for the joint estimates, inferences made from the two-step procedure are still valid.

4. CONCLUSION

A conventional fixed-coefficient model of a monetary-growth function understates the degree of learning by economic agents. A typical test of the rational-expectations hypothesis is performed based on the whole sample, but at time t agents do not have information on $t+1, \ldots, T$. Especially when the policy regime changes continuously, the correct specification of the learning process by agents is crucial to the test result and its interpretation.

A TVP model was proposed in modeling a monetary-growth function, and the Kalman-filtering technique was applied to estimate the model. The Kalman filter shows how rational economic agents would combine past information and new information to form a new expectation. The algorithm also provides recursive forecast errors and their conditional variance at each point in time.

Based on the estimated forecast errors and conditional variances from the Kalman-filtering estimation of a monetary-growth function, the Lucas hypothesis (in a time-series context) was tested. Consideration was also given to the problems associated with the generated regressors as raised by Mishkin (1982) and Pagan (1984); that is, joint estimations were performed as well as two-step estimation. Test results rejected the Lucas hypothesis, but the conditional variance of monetary shocks itself played an important role in explaining the business cycle in the U.S. economy 1964:1–1985:4. In other words, the conditional variance affected real output directly, not through the coefficient on the monetary (or nominal) shock. This result supports Friedman's (1977) view that natural output is negatively dependent on inflation uncertainty.

APPENDIX: DATA

Data used in this article consist of *DM*1 (log differences of *M*1 money stock multiplied by 100), *DINT* (changes in the three-month treasury-bill rate), *INF* (log differences of the CPI multiplied by 100), *SURP* (full-employment budget surplus), and *LNY* [log of real gross national product (GNP)]. The data are described in Table A.1. The computer program, as well as the data, is available from Kim on diskette.

The real GNP series is from the *Survey of Current*

Table 4. Test of the Lucas Hypothesis (differenced output), Corrected for Serial Correlation

Estimation	α_0	γ_0	γ_1	α_2	R^2	DW
Two-step	.00017	.0029	—		.1456	2.0623
	(.113)	(2.640)				
	.1768	.0029	.00048	—	.1465	2.0647
	(.119)	(2.640)	(.3155)			
	.00668	.0038	—	−.012	.2535	2.0132
	(3.164)	(3.499)		(−3.76)		
	.00788	.0039	.0026	−.015	.2775	2.0109
	(3.577)	(3.639)	(1.712)	(−4.16)		
Joint	.00719	.0041	.0028	−.014		
	(2.797)	(3.127)	(1.64)	(−2.59)		

NOTE: t values are in parentheses. $Dcy_t = \alpha_0 + (\gamma_0 + \gamma_1 \ln H_{t|t-1})\eta_{t|t-1} + \alpha_2 H_{t|t-1} + u_t$, $Dcy_t = cy_t - cy_{t-1}$, cy_t = time-detrended real output, $\eta_{t|t-1}$ = recursive forecast error, and $H_{t|t-1}$ = conditional forecast-error variance.

Table A.1. Data

Time	LNY	DM1	DINT	INF	SURP
1959:1	7.38175	—	—	—	−.05600
1959:2	7.40062	.82831	.22667	.19175	.02000
1959:3	7.39603	.75135	.54000	.49685	−.04200
1959:4	7.40452	−.86926	.69000	.60814	−.02900
1960:1	7.42154	−.25989	−.35667	.11361	.21700
1960:2	7.41866	−.16574	−.88000	.60378	.15900
1960:3	7.41962	.94340	−.63333	.03761	.11300
1960:4	7.41101	−.04696	−.05333	.63730	.12600
1961:1	7.42136	.53871	.04333	.22396	.02600
1961:2	7.43373	.76789	−.04667	−.03729	−.01000
1961:3	7.44793	.60088	.00000	.40945	−.00300
1961:4	7.47017	.94027	.15667	.11138	.02900
1962:1	7.48319	.63709	.26333	.40733	−.17400
1962:2	7.49354	.70065	−.01000	.36887	−.13800
1962:3	7.50279	−.11268	.12667	.25740	−.11600
1962:4	7.50114	.56212	−.02667	.25674	−.09000
1963:1	7.51458	1.02611	.09333	.32913	−.03700
1963:2	7.52833	.94976	.03000	.18238	.05900
1963:3	7.54565	1.00613	.35667	.58140	−.01200
1963:4	7.55281	.95301	.20333	.28944	−.04600
1964:1	7.57492	.68744	.03333	.43259	−.19300
1964:2	7.58345	.70400	−.05333	.14378	−.30500
1964:3	7.59347	1.62365	.02000	.17944	−.18500
1964:4	7.59775	1.28846	.18667	.50072	−.12100
1965:1	7.61918	.74059	.20667	.32057	−.01800
1965:2	7.63356	.59262	−.01667	.63807	−.07100
1965:3	7.64936	1.15468	−.00667	.28229	−.31600
1965:4	7.67211	1.81624	.30000	.52715	−.40800
1966:1	7.69170	1.70609	.44333	.94192	−.36200
1966:2	7.69430	1.06374	−.02333	.89873	−.31600
1966:3	7.70450	−.28899	.45667	.85661	−.45300
1966:4	7.70940	.26975	.16667	.81550	−.51000
1967:1	7.71503	1.01465	−.69667	.33784	−.67800
1967:2	7.72099	1.39968	−.85333	.53818	−.66900
1967:3	7.73530	2.19217	.64000	1.03455	−.71400
1967:4	7.74093	1.58603	.45333	.95821	−.67000
1968:1	7.75246	1.34737	.29667	1.07932	−.60100
1968:2	7.76934	1.69855	.47000	.97120	−.72300
1968:3	7.77708	1.92865	−.32333	1.28043	−.48000
1968:4	7.77612	2.11144	.39000	1.23283	−.34500
1969:1	7.79008	1.78681	.50667	1.21781	−.10300
1969:2	7.79144	.80772	.10333	1.50935	−.05100
1969:3	7.79696	.42595	.82667	1.45678	−.14000
1969:4	7.79297	.76541	.33000	1.43586	−.10300
1970:1	7.78680	1.04899	−.14333	1.62030	−.08000
1970:2	7.78593	.73577	−.53333	1.39294	−.28600
1970:3	7.79803	1.34549	−.34667	1.11765	−.33600
1970:4	7.78896	1.77658	−.97667	1.38675	−.32600
1971:1	7.81545	1.71520	−1.51333	.83964	−.37800
1971:2	7.81537	2.10344	.41000	.94314	−.43000
1971:3	7.82048	1.70995	.76000	1.01635	−.40000
1971:4	7.82044	.94553	−.78000	.65378	−.30700
1972:1	7.84212	1.93568	−.79333	.89202	−.16800
1972:2	7.86138	1.63384	.33333	.59029	−.40300

Business, as stored in the Citibase Economic Data Base; *M*1 money stock and the three-month treasury-bill rates are from the Board of Governors of the U.S. Federal Reserve System, as stored in the Citibase Economic Data Base; the CPI series is from the U.S. Bureau of Labor statistics, as stored in the Citibase Economic Data Base; and the *SURP* series is from Gordon (1987).

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Table A.1 (continued)

Time	LNY	DM1	DINT	INF	SURP
1972:3	7.87173	2.04653	.45000	.87895	−.19500
1972:4	7.89032	2.46058	.64333	1.00265	−.56700
1973:1	7.91352	1.99650	.83667	1.51123	−.37500
1973:2	7.91608	1.17128	.90333	2.02234	−.34500
1973:3	7.91509	1.18329	.72000	2.03193	−.18600
1973:4	7.92400	1.24525	−.82333	2.45291	−.18900
1974:1	7.91841	1.66556	.11667	2.93682	−.16000
1974:2	7.92125	.89021	.53667	2.60117	−.25300
1974:3	7.90813	.89456	.03667	2.86993	−.06200
1974:4	7.89930	1.20052	−.83000	3.04949	−.20200
1975:1	7.87956	.70867	−1.61000	2.05311	−.32800
1975:2	7.88968	1.54397	−.35667	1.22030	−1.24400
1975:3	7.90651	1.83341	.93667	1.98764	−.64300
1975:4	7.92034	.80673	−.70333	1.84833	−.62800
1976:1	7.93894	1.33404	−.71000	1.14058	−.56700
1976:2	7.94339	1.62897	.24000	.83218	−.46700
1976:3	7.94754	1.06408	−.00667	1.54678	−.52800
1976:4	7.95746	1.94350	−.47667	1.42749	−.56100
1977:1	7.97109	2.29821	−.04333	1.84083	−.33100
1977:2	7.98708	1.71822	.21000	1.77062	−.44100
1977:3	8.00697	1.68919	.65667	1.37604	−.68100
1977:4	8.00440	2.09215	.61333	1.44720	−.59900
1978:1	8.01318	1.92180	.28333	1.69179	−.57200
1978:2	8.04427	2.19320	.08333	2.28691	−.44000
1978:3	8.05281	2.02392	.83667	2.26953	−.42900
1978:4	8.06514	1.77144	1.25667	2.31812	−.41700
1979:1	8.06517	1.22160	.81333	2.45881	−.22300
1979:2	8.06423	1.82728	−.00667	3.19749	−.12300
1979:3	8.07322	3.40151	.29667	3.14371	−.30500
1979:4	8.07131	1.12003	2.17000	3.12107	−.33500
1980:1	8.08129	1.08220	1.51000	3.86071	−.43500
1980:2	8.05738	−1.49429	−3.73667	3.32302	−.45400
1980:3	8.05804	4.78721	−.46333	1.83516	−.45800
1980:4	8.07066	2.67028	4.46000	2.79517	−.39800
1981:1	8.08982	.43025	.77667	2.69385	−.24300
1981:2	8.08647	1.64798	.51667	2.06684	−.18100
1981:3	8.09089	1.65203	.14667	2.75129	−.29800
1981:4	8.07683	1.23086	−3.30333	1.61516	−.46600
1982:1	8.06161	1.80714	1.06333	.89793	−.38400
1982:2	8.06461	.41701	−.39333	1.40156	−.34600
1982:3	8.05659	2.13212	−3.10333	1.73617	−.66900
1982:4	8.05811	3.93690	−1.41000	.31865	−1.01600
1983:1	8.06671	2.53423	.20000	.10221	−.90700
1983:2	8.08896	2.79038	.29000	1.15112	−.84700
1983:3	8.10362	2.73398	.74333	.99370	−1.01400
1983:4	8.12121	1.61283	−.34000	.95091	−1.14200
1984:1	8.14662	1.53723	.37000	1.29020	−1.12900
1984:2	8.15995	1.58167	.62667	.99450	−1.25500
1984:3	8.16639	1.07117	.52333	.88880	−1.38500
1984:4	8.17053	1.05981	−1.51667	.81754	−1.48700
1985:1	8.17996	2.66302	−.62000	.82140	−1.37700
1985:2	8.18510	2.57107	−.72333	1.11593	−1.89800
1985:3	8.19506	3.49415	−.35333	.57911	−1.74500
1985:4	8.20273	2.67452	.06000	.95439	−1.99600

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