Searching for the Darby Effect in Tax Exempt and Taxable Interest Rate Data

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Abstract

The relationship between tax exempt and taxable interest rates during periods of changing inflation has been used as evidence of a Darby effect—of interest rates rising and falling more than inflation. This paper adjusts for residual autocorrelation in the fixed intercept model used in a previous analysis and finds that this reverses the ordinary least squares (OLS) finding. The relationship also is estimated with two versions of models that allow a variable risk premium in the return on tax exempt bonds. One allows for the intercept to follow a random walk; the other allows the risk premium to respond to recent volatility in the relationship between the interest rates. The results of the variable-risk premium analysis are mixed. Consequently, these findings preclude a firm conclusion that interest rate data confirm the existence of a Darby effect.

Introduction

Irving Fisher (1930) argued that there is a reason for the nominal interest rate to rise and fall about the same amount as the rate of inflation, so that the real interest rate remains approximately unchanged.[1] When Fisher's book was published in 1930, U.S. income tax rates were not high enough to be of much significance. In recent decades, however, tax rates have been too high to ignore. Darby (1975) noted that due to the U.S. income tax treatment of interest, the nominal interest rate might be expected to respond to inflation by an amount larger than the one-for-one adjustment envisioned by Fisher. In a related study, Tanzi (1976) made a similar argument, and a paper by Feldstein (1976) provided a sophisticated theoretical basis for what has become known as the *Darby effect.*

Subsequently there have been several papers that have analyzed the adjustment of interest rates to inflation under a variety of tax systems. It has been shown that when there are separate corporate and personal income taxes, the adjustment of the interest rate depends on how debt and equity returns are treated under the two taxes, how capital gains are taxed, and how depreciation deductions

[1]Although this has become known as the *Fisher effect,* Fisher noted that he was not the originator of the concept. Also, he reported evidence contrary to the effect.

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are computed. Examples of such papers include Feldstein, Green, and Sheshinski (1978) and Miles (1983).

There have been many empirical studies of the relationship between inflation and interest rates, but most have found less than a Fisher effect. In spite of little reported evidence of a Darby effect, Darby and Melvin (1986, p. 85) state:

> While there have been numerous studies with conflicting results, two recent studies may settle the issue: Michael Melvin, "Expected Inflation, Taxation, and Interest Rates: The Delusion of Fiscal Illusion," *American Economic Review* (September 1982, pp. 841-845), and Robert Ayanian, "Expectations, Taxes, and Interest: The Search for the Darby Effect," *American Economic Review* (September 1983, pp. 762-765).

Melvin's paper (1982) argued that the Darby effect may be present in a structural model, even if it is not apparent in a reduced form estimated equation in which the interest rate is on the left side and the expected inflation rate is included in the right side variables. He concluded that his results were consistent with a Darby hypothesis embedded in a structural model, although his estimate of the regression coefficient of the expected inflation rate parameter was less than

$$\frac{1}{1-T},$$

where:

T = an income tax rate.

A cynic might say that Melvin argued that a Darby effect may exist even if it cannot be seen. In any case, results that only can be said to be not inconsistent with a Darby effect do not settle the issue. Therefore, much of the case must rest on Ayanian's results.

This paper notes, however, that Ayanian's results do not settle the issue. When the data used by Ayanian are adjusted for autocorrelation with standard methods, the results are reversed—the Darby effect withers away. Moreover, the model used by Ayanian does not account for a time-varying risk premium on municipal bonds relative to Treasury bills. Therefore, contrary to Darby and Melvin's assertion that the issue is settled, empirical support for the presence of the Darby effect remains tenuous.

This paper demonstrates that standard methods of adjusting Ayanian's data for autocorrelation reverse his results. Due to the time-varying risk premium in municipal securities, it is necessary to estimate the relationship between municipal and Treasury bills (of the same maturity) with a random intercept model.

Ayanian's Fixed Intercept Model

Ayanian (1983) used the following approach in his empirical analysis. The argument begins with the Fisher equation, which states that the nominal interest rate (i) exceeds the expected real interest rate (r) by the expected inflation rate (π), i.e.,

(1) $i = r + \pi$

As Darby, Feldstein, and Tanzi have noted, with the U.S. income tax structure, the relationship becomes

(2) $i[1\text{-}T] = r' + \pi$

where:

r' = the expected after-tax real interest rate, and
T = the marginal lender's marginal tax rate.

The finance literature (e.g., Miller, 1977) implies that the relevant rate is the corporate income tax rate. Ayanian notes the difficulty of measuring the expected after tax real interest rate and the expected inflation rate and suggests that their sum can be measured by the nominal yield on tax exempt bonds (i_X). If so, the relationship between (i_X) and the rate of taxable bonds (i_T) would be

(3) $i_T = \left(\frac{1}{1\text{-}T}\right) i_X.$

Ayanian searched for the Darby effect by regressing (i_T) on (i_X) to see whether the coefficient of (i_X) exceeds one.

Ayanian began with monthly observations of the yields on one year Treasury bills and on one year prime grade municipal bonds, took quarterly averages of the monthly observations (for unstated reasons), regressed them, and reported the results without any mention of a Durbin-Watson statistic (D-W) or an autocorrelation coefficient. For his sample period and two of its subsamples, his procedure produced coefficients significantly greater than one.

A problem with Ayanian's fixed intercept model results is that the original monthly data appear to be characterized by severe autocorrelation. His method of dealing with autocorrelation (taking quarterly averages) is inappropriate and inadequate. When standard procedures for dealing with autocorrelation are applied to the data, the Darby effect disappears. Using the standard Cochrane-Orcutt (CORC) corrective procedure for first order autocorrelation produces coefficients that are significantly less than one, as does regressing first differences of the original observations.

The empirical results from the constant intercept model are summarized in Table 1. To make the current results comparable to Ayanian's, the same sample period (1952:1 through 1979:4) and two subperiods were used. The first three rows show the results of using ordinary least squares (OLS) with the monthly observations. The coefficients appear to be significantly greater than one, but the Durbin-Watson statistics suggest that highly positive autocorrelation is present in all the sample periods. OLS estimates of the regression coefficients are unbiased in the presence of autocorrelation, but OLS seriously underestimates the sampling variance of regression coefficients and unduly inflates the t-statistics for the coefficients. The results in the first three rows were not reported by Ayanian. He took quarterly averages of the monthly observations and reported results almost identical to the second three rows, but he reported neither D-W statistics nor autocorrelation coefficients. A comparison of the D-W statistics shows that autocorrelation appears to be worse with quarterly averages than with monthly observations.

The results of the standard CORC procedure are reported in the third three rows. The regression coefficients are significantly less than one in every sample period. A less elegant, but commonly used, method of dealing with high autocorrelation is to regress first differences of the original observations. The last three rows show that this method also produces regression coefficients significantly below one in every sample period except 1/1952 through 12/1965, for which it is not significantly different from one. Thus, the evidence of a Darby effect in these data disappears when standard methods of dealing with autocorrelation are applied.[2]

The empirical results of the standard fixed intercept regression model are econometrically inefficient and demonstrably unstable. In short, the choice of the estimation technique (OLS or CORC) and the type of data (monthly or quarterly averages of monthly observations) are the main determinants of the apparent

[2]Clouding the evidence still further, tests for higher than first order autocorrelation indicate that it is present. Estimates of the coefficient with adjustments for nth order autocorrelation in the monthly data produce mixed results. In the 1/1952 to 12/1979 sample period, adjustments for nth order autocorrelation produce coefficients significantly less than one for n = 1, 2, coefficients greater than one for n = 8, 9, . . . , 13 and coefficients not significantly different from one for n = 3, 4, 5, 6, 7, 14, 15. These results suggest that the regression equation is incomplete or misspecified--see Judge *et al.* (1982).

presence or absence of the Darby effect. Moreover, the strong evidence of autocorrelation in each sample period may imply that the conditional variance of the error term changes over time. The results of a Lagrangian multiplier (L-M) test indicate that the null hypothesis of a constant conditional variance is rejected at the 1 percent significance level. Nonconstant conditional variance may indicate that municipal bonds have time-varying risk premia. If so, the standard constant intercept regression model is not appropriate, and Ayanian's model is misspecified.

Darby Effect with Stochastic Intercept

Ayanian's paper ignores a body of research in the finance literature (e.g., Miller and Fama, 1977; Trzcinka, 1982) on the relationship between taxable and tax exempt interest rates. These papers make no presumption that the ratio of tax exempt and taxable interest rates sheds any light on the existence of the Darby effect. Some of their observations, however, suggest that if the behavior of this ratio is to be examined for consistency with the Darby effect, then the estimation needs to account for more complications than just autocorrelation.

Trzcinka argues that municipal securities contain elements of risk that are not present with corporate securities of the same ratings and that *a fortiori* would not be present with federal government securities. Trzcinka cites arguments by Hempel (1972) concerning the difficulty of seizing the assets of a bankrupt municipality, by Zimmerman (1977) regarding the questionable usefulness of information supplied by municipalities, and by Fama concerning the relative unpredictability of local government decisions. To these can be added the observation that municipal bonds are subject to a considerable risk of tax reform. These risk factors vary over time, and variation of these presumably explains why municipal bond rates occasionally have exceeded the rate on federal bonds despite the tax advantage of municipal bonds. Trzcinka argues that because the risk factors specific to municipal bonds vary over time, a random intercept model is needed to estimate the relationship between taxable and tax exempt interest rates accurately.

Trzcinka's focus on the Miller hypothesis concerning the relationship between taxable interest rates and municipal bond rates suggested a regression equation with the tax exempt interest rate on the left side and an intercept and a taxable interest rate on the right side. In terms of the current research, Trzcinka's equation (3) is

(4) $i_X - \lambda = (1-T)\, i_T$

where

λ = the stochastic differential risk premium on municipal bonds.

Ayanian's approach to searching for the Darby effect has the taxable rate on the left side, so Trzcinka's formulation must be reversed to be comparable to Ayanian's. The preceding equation can be rearranged as:

$$(5)\quad i_T = \frac{-\lambda}{1-T} + \frac{1}{1-T} i_x.$$

Letting $-\lambda/(1-T) = \delta$, $1/(1-T) = \beta$, and adding time subscripts along with the normal classical error term (u_t), the preceding equation becomes

$$(6)\quad i_{T,t} = \delta_t + \beta i_{x,t} + u_t;\ E(U_t^2) = \sigma^2 .$$

Because δ_t is not a constant, this equation cannot be estimated by standard regression techniques.

To test the Miller hypothesis in the context of a random intercept model, Trzcinka used a technique suggested by Cooley and Prescott (1973). Cooley and Prescott (C&P) argue that in standard regression models it is likely that the intercept changes are due to stochastic factors. In regression models in which the error term follows first order autocorrelation, a common assumption is that the residual dependence is only transitory. Whether this is an appropriate formulation depends on the structure of the model being estimated. Inevitably, structural changes (e.g., changes in technology or preferences) play an important role in many regression models, and the above simplifying assumption is not justifiable. Therefore, in the absence of a built-in mechanism to account for permanent shocks, the regression model is misspecified.

The C&P procedure allows for the possibility of two different types of disturbances embedded in the error term. One type is a disturbance of a temporary nature (cyclical); this can be treated with the usual methods of dealing with autocorrelation. A second type of disturbance is a permanent shock that is secular (lasts over several business cycles) and thus does not respond favorably to autocorrelation correction. These longer-term disturbances may be modeled with a regression model in which the intercept follows a random walk. The hypothesized random walk not only minimizes the impact of permanent shocks and the resulting model misspecifications, but also improves the model's forecasting power immensely. With periodic changes in the riskiness of municipal bonds, the random intercept δ_t is likely to follow a random walk of the following form:

(7) $\delta_t = \delta_{t-1} + v_{t-1}$,

where the error terms of equations (6) and (7) (u_t and v_t) are normally and identically, but independently, distributed with zero means and finite variance σ_u^2 and σ_v^2, respectively. Incorporating the above modifications into equation (6) and reexpressing δ_t for the immediate postsample period

$$(\delta_{n+1} = \delta_t + \sum_{s=t}^{n} V_s),$$

yields

(8) $i_{T,t} = \delta_{n+1} + \beta i_{x,t} + W_t$,

where

$$W_t = u_t - \sum_{s=t}^{n} V_s;\ t = 1, \ldots , n.$$

Because the composite error term is autocorrelated and heteroscedastic, the standard regression techniques are inapplicable. To estimate the model appropriately, C&P suggest a reparametrization of σ_u^2 and σ_v^2 as follows:

$$\sigma_u^2 = (1-\gamma)\,\sigma^2; \quad \sigma_v^2 = \gamma\sigma^2; \quad \text{and } 0 < \gamma < 1.$$

If $\gamma = 0$, then all the variation is transitory and is due to σ_u^2, i.e., equation (8) is a standard regression model with a constant intercept. Conversely, if $\gamma = 1$, the variation is permanent and is caused by σ_v^2. In the absence of either of these two extreme cases, the conditional covariance matrix $\omega(\gamma)$ can be defined as (note that $E(WW') = \sigma^2\omega(\gamma)$)

$$\omega(\gamma) = (1 - \gamma)\, I + \gamma R$$

where:

I = an identity matrix; and
R = a n x n matrix whose elements are

$$r_{ij} = \min\,(n-i + 1, n-j + 1)$$

$\omega(\gamma)$ is predominantly dependent upon the numerical value of (γ), which is not known *ex ante*. C&P contend, however, that the (γ) that can maximize the concentrated and conditional maximum likelihood function (say, $\hat{\gamma}$) should be used to obtain the required Omega matrix. To avoid the tedious job of inverting $\omega(\gamma)$ for each value of (γ) over the interval 0 and 1, C&P introduced a transformation procedure.[3] The empirical results of estimating equation (8) with the C&P procedure are reported in Table 2.

These results are consistent with the existence of the Darby effect. With the monthly data, the estimated value of β is significantly greater than one for the sample period as a whole and for both subsamples. This is also true of the data combined into quarterly averages of monthly data. (Quarterly averages of monthly data are reported for comparison with Ayanian.)[4]

The results with this random intercept model differ from those with the simpler estimation techniques reported in Table 1—evidence of a Darby effect is detected, regardless of the choice of sample period and whether monthly observations are combined into quarterly averages.

There is reason not to take the Cooley and Prescott results as the final word on this relationship, however. The numerical value of the variance share $(\hat{\gamma})$ that maximizes the conditional maximum likelihood function $\omega(\hat{\gamma})$ is about $(\hat{\gamma} = 0.10)$. This implies that a relatively small portion of the composite error term is attributable to a random walk of the intercept. Consequently, the relationship was estimated with an alternate specification of the time-varying risk premium on municipal bonds.

Engle, Lilien, and Robins (1987) introduced a model that has become known as the ARCH-M model (autoregressive conditional heteroscedasticity in

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[3]The proof of the different properties of this transformation is beyond the scope of this paper. Nevertheless, a brief discussion may be useful. For the R-matrix, there exists an orthogonal eigenvectors matrix (p) by which the variables in equation (8) can be transformed easily irrespective of (γ), as follows:

$$pi_{T,t} \sim N\left\{ P\beta i^*_{x,t}, \sigma^2 P[(1-\gamma)I + \gamma R]P' \right\}$$

or

$$\sim N[P\beta i^*_{x,t}, \sigma^2 D(\gamma)]$$

The resulting conditional Omega matrix (ω^*_γ) is diagonal $[D(\gamma)]$, and the elements on the main diagonal are the eigenvalues. The inversion of ω^*_γ is easier than the original $\omega(\gamma)$ and enhances the search of $(\hat{\gamma})$. Upon estimation of $(\hat{\gamma})$, the transformed Omega matrix (ω^*_t) is known, and generalized least squares (GLS) can be applied to the model--see Cooley and Prescott (1973, p. 367).

[4]Aggregation over time (e.g., quarterly averages of monthly data) causes autocorrelation among individual disturbances of the resulting observations, even if they originally were uncorrelated--see Johnston (1984, pp. 287-291).

the mean). This model is designed for estimation in the presence of a time-varying risk premium. They used the ARCH-M model to analyze the changing risk premium in the term structure of interest rates. The main feature of the model is that the variable risk premium (proxied by the conditional variance of the error term) is a significant determinant of the excess holding yield of a risky asset. Furthermore, it efficiently and consistently demonstrates that the above risk factor is a function of its past squared surprises.

Engle, Lilien, and Robins modeled a changing risk premium as a function of the conditional variance in the excess holding yield on a riskier asset. The idea of ARCH-M is that high variance in the recent past adds to the risk premium paid at the present time. The concepts underlying the ARCH-M process are derived from the familiar return (mean)-risk (variance) analysis of Tobin (1958), according to which an asset with (now undiversifiable) riskiness, as measured by the variance of its returns, will have a higher average return to compensate a risk-averter for its higher risk.

The ARCH-M approach suggests that the risk premium at a particular time is a function of surprises (conditional variance) in the excess holding yield in the recent past. Consequently, the risk premium on municipals would be higher following a period of irregularity in the relationship between taxable and tax exempt yields than it would be following a relatively tranquil period. The Engle, Lilien, and Robins formulation formally stipulates that the conditional variance of the excess-holding yield in one period is positively related to its recent past. They found that for the term structure of interest rates, the ARCH-M mechanism was a significant aspect of the risk premia on longer-term Treasuries relative to shorter-term Treasuries.

To determine whether the ARCH-M is appropriate for use in estimating a regression equation, a L-M test appears to be necessary. A simple version of the L-M test is suggested by Engle: One begins with a standard regression model estimated with OLS. The squared residuals of the model are regressed on their lagged values P times, and the coefficient of determination (R^2) is obtained. The number of observations (n) multiplied by R^2 ($n*R^2$) is distributed as a chi-square (X^2) with P degrees of freedom. If the numerical value of $n*R^2$ is greater than the tabled $X^2{}_{d.f.p}$, then the null hypothesis of a constant conditional variance is rejected at a given significance level. The results of the L-M test for equation (3) on the basis of a fourth order ARCH-disturbance are reported below:

Data	Sample Period	X^2
Monthly	January 1952-December 1979	119
Quarterly	1952:1-1979:4	53.76

With $X^2_{d.f.4} = 13.277$, the alternative hypothesis of a volatile conditional variance is accepted at the 1 percent significance level.

To put the analysis of this paper in a framework suitable for estimation with the ARCH-M procedure, it is noted that the (λ) parameter in equation (4) varies over time. Instead of following a random walk as it did in the Cooley and Prescott approach, however, (λ) is determined by the conditional variance of the error term in equation (6). With reference to equation (5), note also that the risk premium (λ) is preceded by a negative sign.

In terms of the symbols of the current research, the ARCH-M model is

$$i_{T,t} = \alpha + \beta i_{x,t} + \delta \text{Log } h_t + \varepsilon_t$$

$$\varepsilon_t/\text{past information} \approx N(0,h_t^2)$$

$$h_t^2 = \gamma_0 + \gamma_1 \sum_{i=1}^{4} \omega_i \varepsilon_{t-i}^2$$

where:

$i_{T,t}$ = Treasury bill rate;
$i_{x,t}$ = Municipal bond rate;
α = the intercept;
β = the Darby coefficient;
δ = the ARCH-M coefficient;
h_t^2 = the conditional variance of the error term - $E(\varepsilon_t^2)$;
ω_i = a weight parameter ($\omega_i = (5-i)/10$; $i = 1, 2, 3, 4$);
γ_0 = the intercept;
γ_1 = the ARCH-coefficient;
h_t = the conditional standard error.

The empirical findings are reported in Table 3.

As can be seen, the evidence of ARCH is strong, which implies that the conditional variance (h_t^2) changes over time. The varying conditional variance (i.e., the conditional heteroscedasticity of the error term) is indicative of a time-varying risk premium associated with municipal bonds. The ARCH-M coefficients, however, are not statistically significant for the entire sample period and the two subsamples, regardless of the type of data used.

The significance of ARCH in this context has further implications. First, the forecasted value of (λ) can be improved because its variation is determined by past squared surprises (especially if small variations tend to be grouped together and large variations also are generally grouped together). Stated differently, (λ)

tends to be more volatile after radical disturbances than following periods of economic tranquility. Second, because $\hat{\varepsilon}^2_{t-i}$ serves as a proxy for the missing variables $i_{T(t-i)}$ and $I_{X(t-i)}$, the ARCH formulation may capture specification errors in standard regression models. In short, it is more desirable to minimize the mean impact of the omitted variables by inclusion of the ARCH instead of conducting an endless search for all the relevant explanatory variables. Third, unlike the classical case in which the unconditional mean of the error term is zero and the variance is finite, the conditional mean and variance are time variant. Consequently, risk averters can readjust the expected value of (λ) continuously in an attempt to optimize the overall risk exposure. Finally, when the logarithm of the conditional standard error term is incorporated into equation (6), the statistical significance of the Darby coefficients drops considerably. With the quarterly observations over the two subsamples and with the monthly data over the later part of the sample period, the regression coefficient is not significantly different from one.

Conclusion

There is no consistent evidence of a Darby effect in these data. The OLS results indicate a Darby effect, but these results are reversed when the fixed intercept model is adjusted for first order autocorrelation. Moreover, the variable risk premium models provide mixed empirical results. In the Cooley and Prescott model in which the risk premium follows a random walk, the Darby coefficient is significantly greater than one. In the ARCH-M model, the Darby coefficient is estimated to be greater than one, but not significantly above one in three of the four subsamples. Therefore, the overall results do not seem to provide strong support for the presence or absence of a Darby effect. With these empirical findings, the matter seems to be open to debate.

There was a theoretical basis for the Darby effect during the sample period for these data. When most nominal interest was taxed and most nominal capital gains were tax exempt, there was reason for the Darby effect to exist. The existence of a Darby effect in the 1/1952-12/1979 period has not been disproven. These data are demonstrably sensitive to the method of estimation and, therefore, do not seem to provide a consistent basis for a strong conclusion.

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Table 1
Results from Regression of Taxable Interest Rate on Tax Exempt Interest Rate

	Intercept	Regression Coefficient	D-W	R^2
		OLS—Monthly Data		
1/1952- 12/1979	0.181 (0.068)	1.62* (0.023)	0.51	0.93
1/1952- 12/1965	-0.04 (0.074)	1.75* (0.043)	0.58	0.90
1/1966- 12/1979	0.465 (0.236)	1.54* (0.061)	0.50	0.79
		OLS—Quarterly Averages		
1952:1- 1979:4	0.137 (0.110)	1.63* (0.037)	0.40	0.94
1952:1- 1965:4	-0.066 (0.119)	1.74* (0.069)	0.90	0.92
1966:1- 1979:4	0.263 (0.393)	1.60* (0.103)	0.33	0.81
		CORC—Monthly Data		
1/1952- 12/1979	3.59 (1.28)	0.545 (0.060)	1.89	0.98
1/1952- 12/1965	1.54 (0.332)	0.807 (0.106)	1.91	0.96
1/1966- 12/1979	4.69 (0.907)	0.516 (0.082)	1.88	0.95
		First Differences—Monthly Data		
1/1952- 12/1979		0.581 (0.061)	1.93	0.22
1/1952- 12/1965		0.98** (0.110)	2.02	0.38
1/1966- 12/1979		0.494 (0.082)	1.89	0.18

Numbers in parentheses are standard errors. Regression coefficients with an asterisk (*) are significantly greater than one. All others are significantly less than one at the 5 percent significance level, except those marked with two asterisks (**). Data on the taxable interest rate and the tax exempt prime grade municipal bond rate are drawn from the *Federal Reserve Bulletin* and *Solomon Brothers Bond Market Roundup*, respectively

Table 2
Regression Results of the Random Intercept Model

	Intercept	Regression Coefficient	D-W	R^2
		Monthly Data		
1/1952-12/1979	0.0149	1.67	2.00	0.98
	(0.0265)	(0.010)		
1/1952-12/1965	0.015	1.75	1.98	0.99
	(0.023)	(0.013)		
1/1966-12/1979	-0.0049	1.70	2.00	0.98
	(0.061)	(0.014)		
	Quarterly Averages of Monthly Observations			
1952:1-1979:4	0.0185	1.67	2.00	0.99
	(0.043)	(0.014)		
1952:1-1965:4	0.284	1.69	1.99	0.98
	(0.041)	(0.025)		
1966:1-1979:4	0.16	1.67	2.00	0.99
	(0.101)	(0.022)		

Numbers in the parentheses are standard errors. All the regression coefficients (β) are significantly greater than one at the 1 percent significance level

Table 3
Regression Results of the ARCH-M Model

Sample Period	α	β	δ	γ_0	γ_1	L
			Monthly Data			
1/1952-	-0.22	1.73*	-0.092	0.14**	0.59**	120.9
12/1979	(0.43)	(0.18)	(0.15)	(0.012)	(0.081)	
1/1952-	-0.36	1.82*	-0.11	0.099**	-0.075**	234.39
12/1965	(0.48)	(0.33)	(0.17)	(0.008)	(0.0092)	
1/1966-	0.77	1.48	0.063	0.21**	0.73**	28.09
12/1979	(2.20)	(0.59)	(0.37)	(0.034)	(0.14)	
		Quarterly Averages of Monthly Observations				
1952:1-	-0.31	1.75*	-0.12	0.19**	0.29**	43.65
1979:4	(0.78)	(0.31)	(0.30)	(0.027)	(0.13)	
1952:1-	-0.38	1.87	-0.086	0.066**	0.13	88.73
1965:4	(0.96)	(0.60)	(0.36)	(0.014)	(0.17)	
1966:1-	0.46	1.56	0.031	0.34**	0.19	5.26
1979:4	(4.22)	(1.11)	(0.80)	(0.06)	(0.14)	

Numbers in parentheses are standard errors

* The regression coefficient is significantly greater than one at the 5 percent significance level

** The regression coefficient is significantly different from zero at the 5 percent significance level

The ARCH-M model also was estimated with h_t in place of log h_t and different numerical values for ω_i, but the empirical findings were similar to those reported

Because h_t^2 is variable over time, the estimator in the ARCH-M model is the maximum likelihood (M-L) procedure, which allows for the abnormality of the error term. The distinctive features of the M-L have been demonstrated by Engle, Lilien, and Robins. The log likelihood is depicted by (L) which is clearly larger with the monthly observations than with the quarterly averages