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A Cointegration Analysis of the Relationship Between Underwriting Margins and Interest Rates: 1930-1989

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ABSTRACT

This article presents evidence that 90-day Treasury bill rates and underwriting margins for stock property-liability insurers were negatively related and cointegrated over the period 1930-1989. These results imply that underwriting margins fluctuated around a long-term equilibrium relationship with interest rates in a way that is consistent with financial models of insurance prices.

Introduction

The property-liability insurance industry has long been characterized by profit cycles. The severe downturn in insurer profitability and coverage availability which began in 1983 and continued into 1985 increased the attention that the underwriting cycle has received in the literature. A portion of the literature has focused on measurement of the cycle (Venezian, 1985; Fields and Venezian, 1989; Doherty and Kang, 1988; Berger, 1988; and Cummins and Outreville, 1987). The prevailing opinion of this literature is that the cycle has an approximate length of six years. Another portion of the literature has focused on determining a fair underwriting margin for insurers (Hill, 1979; Fairley, 1979; Kraus and Ross, 1982; Doherty and Garven, 1986; Myers and Cohn, 1987; and Cummins, 1990, 1991). This literature provides a theoretical base from which an equilibrium premium rate is derived. All of the models developed in this literature are extensions of financial asset pricing models. The three most well-known models are the insurance capital asset pricing model (CAPM), the discounted cash flow model (DCF), and the option pricing model

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(OPT).[1] A general result of these financial pricing models is that the equilibrium price of insurance is negatively affected by the movement of interest rates.

The objective of this article is to empirically verify the relationship between premiums and interest rates and also to determine whether this relationship possesses the characteristics of equilibrium behavior. Doherty and Kang (1988) and Smith (1989) analyzed the effect interest rates have on underwriting margins and premium rates. Doherty and Kang used a traditional supply and demand construct to evaluate whether the underwriting cycle is the result of the industry making rational adjustments toward equilibrium. They conclude that the behavior of capital market rates provides a continuing series of exogenous shocks to which the industry is unable to make instantaneous adjustments. The slow adjustments manifest themselves as a cycle. Smith shows that, from 1950 through 1982, property-liability premium rates reflected taxable bond yields, and he thereby concludes that such a result is indicative of competitive behavior.

The purpose of this article is to enhance the results of Doherty and Kang and Smith by providing more robust estimations using the cointegration methodology. The Doherty and Kang analysis used only differenced data, so, at best, it captures only short-term dynamics. Cointegration captures long- and short-term dynamics. Smith's conclusion that the reflection of taxable bond yields in premium rates is indicative of competitive behavior is made more substantial with cointegration analysis, because cointegration analysis explicitly determines whether the short-term behavior of the underwriting margin is stationary. Stationarity is of significance because a stationary variable possesses the tendency to return to a particular value, and such a tendency is a characteristic of equilibrium behavior.

Verification of equilibrium dynamics has two important implications. First, the existence of equilibrium dynamics suggests that the persistence of the underwriting cycle is not due to irrationality on the part of insurers, a finding that would be consistent with Cummins and Outreville (1987). Second, an industry that responds to market conditions by moving toward competitive equilibrium requires different types of regulatory attention than an industry that does not.

The time period of analysis—1930 through 1989—encompasses many changes in the property-liability insurance industry. Cointegration analysis is robust enough to permeate these changes because cointegration separates long-term and short-term phenomena. Any structural changes that took place in the industry from 1930 to 1989 were long-term adjustments.[2] Hakkio and Rush (1991) conclude that the length of a data set in a cointegration study is of greater importance than the number of observations.

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[1] Cummins (1990,1991) gives very good reviews of the financial asset pricing models as they apply to insurance pricing.

[2] It will be discussed below that for a variable to be considered a candidate for cointegration it must be integrated of order one, which dispels the notion of needing a constant mean or variance over the entire time period.

The remaining four sections of this article provide, first, a brief overview of financial asset pricing models that have been applied to insurance pricing; second, a description of the cointegration methodology; third, the empirical results of the cointegration analysis; and, fourth, conclusions and closing remarks.

Financial Models of Insurance Pricing

The fair level of profit for insurers is earned from two sources, investment activity and underwriting activity. In its simplest form this relationship can be written as

$$I = r_a A + r_u P, \tag{1}$$

where I = fair level of profit,
r_a = investment income as percent of mean assets,
A = mean assets of the insurer,
r_u = underwriting income as percent of premiums earned, and
P = premiums earned.

Cummins (1990, 1991) shows that equation (1) can be expressed as return on equity, r_e:

$$r_e = r_a + s(r_a k + r_u), \tag{2}$$

where s = the premiums-to-surplus ratio, and
k = the liabilities-to-premiums ratio (funds generating coefficient).

From equation (2), it is clear that a firm's return on investment has direct influence on what level of underwriting margin is considered acceptable by an insurer.

Insurance Capital Assets Pricing Model

The insurance CAPM was the earliest application of financial asset pricing to be applied to insurance pricing.[3] The insurance CAPM determines the equilibrium underwriting margin to be

$$E(r_u) = -kr_f + \beta_u(E(r_m) - r_f), \tag{3}$$

where r_f = the risk-free rate,
$E(r_m)$ = the expected market return, and
$\beta_u = Cov(r_u, r_m)/Var(r_m)$, or the underwriting beta.

The insurance CAPM shows that, in equilibrium, the underwriting return consists of a reward for risk-bearing in the amount of $\beta_u(E(r_m)-r_f)$, less an interest credit to policyholders for the use of borrowed funds, $-kr_f$. As long as the risk-premium is relatively stable, equation (3) predicts an inverse relationship between underwriting margin and interest rates.

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[3] For derivation of the insurance CAPM and a complete discussion of all its implications, see Fairley (1979), Hill (1979), and Cummins (1990).

Discounted Cash Flow Models

Discounted cash flow insurance pricing models are based on the net present value capital budgeting concepts of corporate finance.[4] These models attempt to determine the equilibrium price of insurance by discounting the expected loss payment flows with a properly risk-adjusted discount rate. The premium rate for the basic one-year model, in which all premiums are collected at the beginning of the year with all losses being paid at the end of the year, can be expressed as

$$p = \frac{(1+\pi)L}{(1+i)(1+r_f^*+\lambda)}, \tag{4}$$

where L = the amount of losses,
i = the general inflation rate,
π = the insurance inflation rate,
r^*_f = the real risk-free rate of interest, and
λ = a risk premium.

Under the Fisher hypothesis, $(1+i)(1+r^*_f) = (1+r_f)$, and again the pricing model predicts an inverse relationship between asset returns and premiums.

Option Pricing Models

One of the most recent applications of financial asset pricing models has been option pricing.[5] Assuming an insurer begins with equity in the amount of E, and again assuming all premiums, P, are collected at the beginning of the year, and all losses, L, are paid at the end of the year, the insurance premium can be expressed as the riskless present value of anticipated loss payments, less the value of the insolvency put option (Cummins, 1991):

$$P = Le^{-r_f^*\tau} - Put[P+E;\tau,L]. \tag{5}$$

The first derivative of equation (5), with respect to r^*_f is

$$\frac{\partial P}{\partial r_f^*} = -\tau L e^{-r_f^*\tau}[1 - \int_{d_2}^{\infty}\frac{1}{\sqrt{2\pi}}e^{\frac{-x^2}{2}}dx] < 0, \tag{6}$$

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[4] For a discussion of net present value decisions, see Copeland and Weston (1983) or almost any corporate finance textbook. For derivation and discussion of DCF models, see Myers and Cohn (1987) and Cummins (1990, 1991).

[5] For derivation and discussion of the application of OPT to insurance pricing, see Doherty and Garven (1986) and Cummins (1988, 1990, 1991).

which connotes an inverse relationship between underwriting margins and interest rates.

Each insurance pricing model has its own peculiarities, but each of them concludes that the movement of interest rates in the economy has a negative influence on the level of equilibrium underwriting margins. It is this relationship that is the focus of the cointegration study presented here.

Cointegration Methodology

The time series methodology employed here is cointegration analysis.[6] The underlying premise of cointegration analysis is that, while time series variables may contain secular trends (thus resulting in the variables having a unit root), sometimes, due to a unique economic relationship between variables, a linear combination of these variables does not contain a unit root.[7] This linear combination represents the long-term relationship between the variables, a relationship that may be regarded as equilibrium, or an attractor. In the two-variable case, which is used here, the linear combination forms a line that connects the long-term equilibrating pairs of values of the two variables. The deviations from the line represent the short-term movements around equilibrium.

Cointegration analysis proceeds by first determining if the variables under consideration are individually integrated of order one, denoted by I(1). An I(1) variable will contain one unit root and will need to be differenced once to become stationary, that is, I(0). If it is determined that the variables are I(1), then a test for cointegration can be performed by determining whether a linear combination of the variables can be found which will be I(0) and have a zero mean. Consider the simplest case of two variables, x_t and y_t, which are both distributed I(1). These variables are cointegrated if the linear combination $(x_t - Wy_t)$ is distributed I(0). In an application, this linear combination is known as the cointegrating equation, and W, the cointegrating coefficient, can be estimated using ordinary least squares (see Engle and Granger, 1991).

The value of W can be interpreted as the long-term relationship between the variables x_t and y_t. The short-term behavior of the variables is revealed with further evaluation of $(x_t - Wy_t) = z_t$. To do this, two error-correction models are

[6] For a thorough introduction to cointegration analysis, see Engle and Granger (1991).

[7] If a variable contains a unit root, the variable is said to be nonstationary. A nonstationary variable has neither a fixed mean nor a constant variance. A nonstationary variable, such as a random walk, will not fluctuate about a certain value. A (strictly) stationary variable, by contrast, has a fixed mean and a constant variance. The behavior of a stationary variable, when plotted over time, is characterized by numerous fluctuations about the mean. White noise, for instance, is a stationary process. A stationary variable does not contain a unit root. As discussed below, it is sometimes necessary to distinguish between mean stationarity and variance stationarity.

estimated, one for each I(1) variable under consideration. These models take the form

$$\begin{aligned} \Delta x_t &= \alpha_1 + \rho_1 z_{t-1} + \sum_{i=1}^{g} \delta_{1i} \Delta x_{t-i} + \sum_{i=1}^{g} \eta_{1i} \Delta y_{t-1} + \varepsilon_1 \\ \Delta y_t &= \alpha_2 + \rho_2 z_{t-1} + \sum_{i=1}^{g} \delta_{2i} \Delta x_{t-i} + \sum_{i=1}^{g} \eta_{2i} \Delta y_{t-i} + \varepsilon_2, \end{aligned} \tag{7}$$

where the number of lags, g is chosen arbitrarily, at least one of ρ_1, ρ_2 is nonzero, and $(\varepsilon_1, \varepsilon_2)$ is bivariate white noise.

The coefficients ρ_1 and ρ_2 measure the short-term behavioral dynamics by determining the relationship between the deviations from the attractor line, z_{t-1}, and the changes in the cointegrated variables, Δx_t and Δy_t.

Empirical Results

The cointegration analysis performed here covers all stock property-liability insurers from 1930 through 1989. Annual data are used, and the two variables are the underwriting margin and the 90-day Treasury bill rate for new issues (risk-free rate).[8]

Unit Root Tests

The first step of cointegration analysis is determining whether the variables being considered are indeed distributed I(1). These variables are the underwriting margin, r_u, and the risk-free interest rate, r_f. To test for unit roots, an Augmented Dickey-Fuller test (ADF) was performed on each variable. The results of these tests are shown in Table 1. I conclude that both variables are I(1).

Cointegration Estimation

The cointegration technique employed here, as developed by Engle and Granger (1987), is a two-step method. The first step is ordinary least squares (OLS) estimation of the relationship between the risk-free interest rate and the underwriting margin. The regression is formed with the risk-free interest rate as the independent variable. The fitted line measures the long-term equilibrium path that exists between these two variables. The deviations from the line (error terms) represent the short-term movements about equilibrium. If it is determined that the deviations about the fitted line are stationary, the cointegration

[8] The underwriting margin data were collected from A. M. Best (1959, 1990). The figure used is the statutory underwriting income divided by earned premiums. The risk-free rate was collected from U.S. Department of Commerce (1955, 1969, 1989).

Table 1
Augmented Dickey-Fuller Tests for Units Roots

	T-Values	Number of Observations	Lags
Levels			
Risk-Free Rate, r_f	-0.987	56	3
Underwriting Margin, r_u	-1.017	56	3
First Differences			
Risk-Free Rate, r_f	-4.596	55	3
Underwriting Margin, r_u	-4.509	55	3

Sources: A. M. Best (1959, 1990), U.S. Department of Commerce (1955, 1969, 1989).
Note: The null hypothesis of these tests is that the variables contain a unit root, while the alternative is that the variables are I(0). If the reported t-value is greater than the critical t-value, the null hypothesis is accepted. The 5 percent and 10 percent critical values for these unit root tests are -2.93 and -2.60, respectively (see Fuller, 1976, p. 373). The data are annual industry figures for stock property-liability insurers for 1930 through 1989.

hypothesis is upheld because stationary error terms possess the tendency to return to the line.

The estimated cointegrating equation is

$$\underset{}{r_u} = \underset{(4.50)}{3.752} - \underset{(-6.89)}{1.146}r_f, \tag{8}$$

where the t-values are reported below the estimates, the adjusted $R^2 = 0.44$, and the F-ratio = 47.5. The magnitude of these estimates appears reasonable, and the sign of the slope coefficient is negative as expected.[9]

An ADF unit root test was performed on the error terms from the estimation of the cointegrating equation. The t-value of the test was -4.124, which rejects the presence of a unit root at a significance level of five percent.[10] Since the error terms are stationary, the cointegration hypothesis is upheld.

In an effort to test for the presence of heteroskedasticity, a Goldfeld-Quandt test was performed on the error terms of the cointegrating equation. The null hypothesis of homoskedasticity was rejected at a significance level of one percent, in deference to the alternative hypothesis which held that the variance

[9] The OLS estimates of the cointegrating equation are not to be interpreted directly from any of the pricing models (say, the CAPM). The objective of the cointegration analysis is to confirm the existence of an attracting equilibrium. To accomplish this objective, the first-step estimation need only produce an unbiased, consistent estimator of the cointegrating coefficient. A consistent first-step estimator will result in asymptotically efficient estimates in the second step (see Engle and Granger, 1991, chap. 1).

[10] There were two lags used in this ADF test, which left 57 observations for computation. The null hypothesis of this test is that the variable contains a unit root, while the alternative is that the variable is I(0). If the reported t-value is greater than the critical t-value, the null hypothesis is accepted. The 5 percent and the 10 percent critical values for the ADF test are -3.29 and -2.90, respectively (see Engle and Yoo, 1987). No intercept term was used for these unit root tests.

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of the error terms was larger in the later years of the sample.[11] This finding indicates that the series of error terms are nonstationary in variance. Such a result, though, does not undo the cointegration hypothesis. Hansen (1992) shows that OLS estimates of heteroskedastic cointegration models are consistent, and he concludes that the dominant property (requirement) of the cointegrating residuals is mean stationarity. As a result of the presence of heteroskedasticity, the reported t-values for the cointegrating equation estimates are of dubious inference quality.

The second step of the cointegration estimation is OLS estimation of two error-correction models (ECM). These models are

$$\begin{aligned} \Delta r_{u_t} &= \alpha_1 + \rho_1 z_{t-1} + \sum_{i=1}^{4} \delta_{1i} \Delta r_{f_{t-i}} + \sum_{i=1}^{4} \eta_{1i} \Delta r_{u_{t-i}} + \varepsilon_1 \\ \Delta r_{f_t} &= \alpha_2 + \rho_2 z_{t-1} + \sum_{i=1}^{4} \delta_{2i} \Delta r_{f_{t-i}} + \sum_{i=1}^{4} \eta_{2i} \Delta r_{u_{t-i}} + \varepsilon_2, \end{aligned} \tag{9}$$

where z is the vector of error terms from the estimation of the cointegrating equation. The results of the ECM estimation are reported in Table 2.

The important feature of the ECM estimation is the error term variable, z_{t-1}. This variable is significantly different from zero in only the underwriting equation, that is, Δr_u. This nonzero coefficient on z_{t-1} in the Δr_u equation indicates that premium rates are the primary focus of equilibrium adjustments. The magnitude of the coefficient, -0.48, suggests that the equilibrium adjustments are partial adjustments. The sign of the coefficient is negative as expected. The negative coefficient confirms that when the property-liability industry is operating above the equilibrium line, there is a tendency to reduce rates.

Figures 1A, 1B, and 1C display the nonstationary cointegrated variables, r_f and r_u, along with the mean stationary error terms, z_t. It can be seen in these figures that as the risk-free rate has trended upward, the underwriting margin, with its characteristic oscillations, has trended downward. When the effect of the risk-free rate is removed from the underwriting margin, the mean stationary series in Figure 1C is the result. The mean stationary series reveals that the short-term deviations from the cointegrating equation consistently return to the long-term relationship. The larger variability of errors in the later years (as confirmed by the Goldfeld-Quandt test) can be interpreted as a more volatile equilibrium situation, but it can possibly, and more technically, be described

[11] For the Goldfeld-Quandt test, the ten central observations were dropped, leaving 25 observations for each regression. The test value was 3.063, and the $F_{24,24}$ critical value is 2.66. The sum of squares for the early years regression was 254.564, while the sum of squares for the later years was 779.591.

Table 2
Estimation Results of Error-Correction Models

	Δr_u		Δr_f	
	Coefficient	T-Value	Coefficient	T-Value
Intercept	-0.093	-0.279	0.243	1.611
z_{t-1}	-0.480	-2.833	0.032	0.414
Δr_{ft-1}	0.157	0.405	-0.014	-0.079
Δr_{ft-2}	0.862	2.279	-0.336	-1.970
Δr_{ft-3}	-0.201	-0.532	-0.046	-0.269
Δr_{ft-4}	-0.198	-0.532	-0.017	-0.103
Δr_{ut-1}	0.550	3.341	0.091	1.223
Δr_{ut-2}	0.0154	0.104	0.044	0.664
Δr_{ut-3}	-0.325	-2.365	0.103	1.666
Δr_{ut-4}	0.253	1.718	0.104	1.568

Sources: A. M. Best (1959, 1990), U.S. Department of Commerce (1955, 1969, 1989).
Note: For Δr_u, F = 8.286 and adjusted R^2 = 0.548. For Δr_f, F = 2.516 and adjusted R^2 = 0.202. The data are annual industry figures for stock property-liability insurers for 1930 through 1989. The 5 percent critical t-values for n = 60 are ± 2.00 (standard t distribution).

as a bi-integrated process in which larger residual variances are associated with larger values of the (risk-free interest rate) regressor. Hansen (1992) has developed a general model of this type.

Impulse Response Function

As an addendum to the cointegration analysis, an impulse response function has been derived in an attempt to measure the response of the underwriting margin to a one-time shock in the risk-free interest rate. Impulse response functions are derived from vector autoregressions (VAR) and are common tools of economic analysis.

The ECMs estimated above can be classified as a vector autoregression, but the presence of the z_{t-1} terms does not allow the direct computation of the impulse response function. Lütkepohl and Reimers (1992) show how impulse response functions can be derived from cointegration systems.

The ECMs must first be written in the following form:

$$e\Delta X_t = \mu_t + \Gamma_1 \Delta X_{t-1} + \Gamma_2 \Delta X_{t-2} + \Gamma_3 \Delta X_{t-3} + \Gamma_4 \Delta X_{t-4} + \Pi X_{t-1} + \varepsilon_t, \qquad (10)$$

where the X matrix is a stacked vector equal to $[r_u\ r_f]'$, and ΔX_t is equal to the first differences of this stacked vector.

Ordinary least squares estimation of equation (10) is used to form the following more conventional VAR model:

Figure 1A

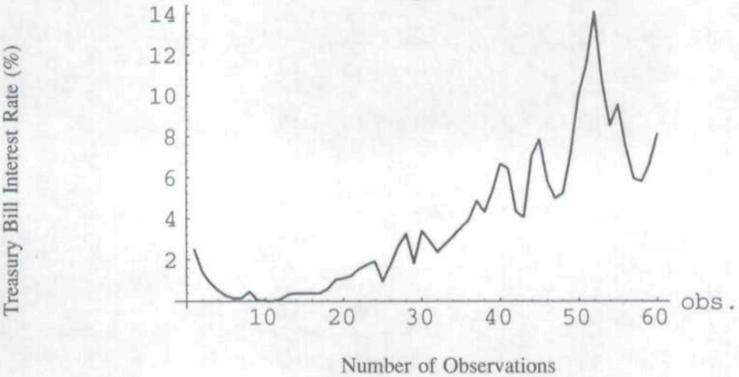


Figure 1B

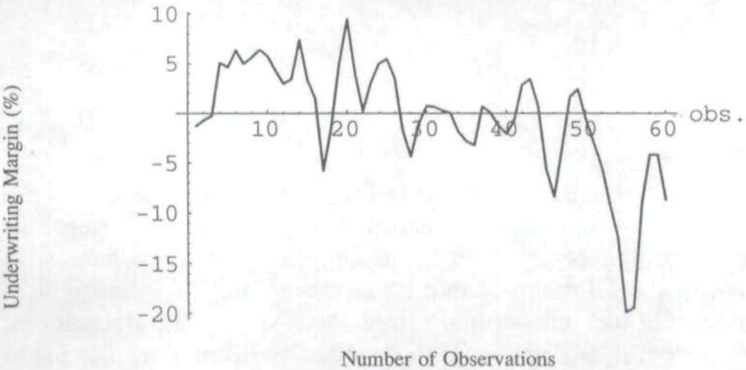
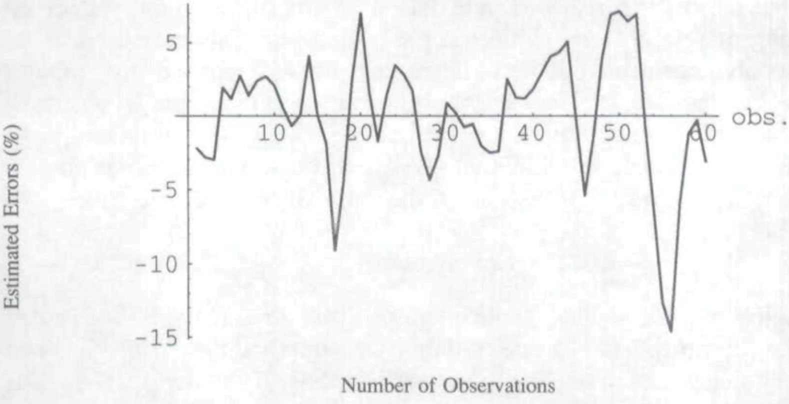


Figure 1C



The data are annual industry figures for stock property-liability insurers, for 1930-1989.

$$X_t = \mu_t + \Theta_1 X_{t-1} + \Theta_2 X_{t-2} + \Theta_3 X_{t-3} + \Theta_4 X_{t-4} + \Theta_5 X_{t-5} + \varepsilon_t, \tag{11}$$

where
$$\begin{aligned} \Theta_1 &= I_{(2x2)} + \hat{\Gamma}_1 + \hat{\Pi}, \\ \Theta_2 &= \hat{\Gamma}_2 - \hat{\Gamma}_1, \\ \Theta_3 &= \hat{\Gamma}_3 - \hat{\Gamma}_2, \\ \Theta_4 &= \hat{\Gamma}_4 - \hat{\Gamma}_3, \\ \Theta_5 &= -\hat{\Gamma}_4. \end{aligned}$$

From the Θ_i coefficients, the impulse responses are computed as follows:

$$\phi_n = \sum_{j=1}^{N} \phi_{n-j} \Theta_j, \qquad n = 1,2,... \tag{12}$$

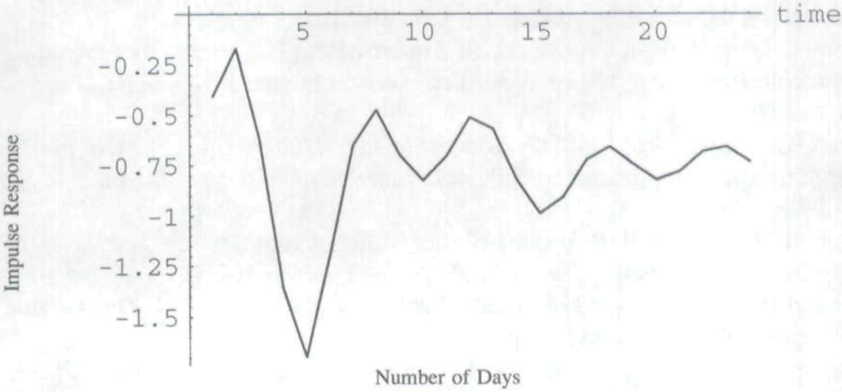
where $\phi_0 = I_{(2x2)}$, and
$\Theta_j = 0$ for $j > p$.

Figure 2 plots the impulse responses on the underwriting margin from a positive one-time unit shock (one standard deviation) to the risk-free rate. Interpretation of the plot begins with the assumption that the system is in equilibrium and that equilibrium is located at the origin. The plotting then shows that the level of the underwriting margin, after some sharp fluctuations, settles, *ceteris paribus*, at a new equilibrium that is lower than the initial equilibrium. The distance between the new equilibrium and the initial equilibrium is equal to 75 percent of the unit shock. This means that a positive one-time shock on the risk-free rate has a permanent negative effect on underwriting margin, as long as no other shocks occur. This may seem to be an unreasonable conclusion at first glance, but the VAR basis of the impulse responses is equation (11). The variables in equation (11) are the levels of the risk-free rate and the underwriting margin and are, therefore, nonstationary. As such, and in the absence of additional shocks, these variables do not possess the tendency to return to a prior level in the wake of the one-time shock.

Conclusion

The major conclusion that can be drawn from this analysis is that the underwriting margin of stock property-liability insurers and the risk-free interest rate were cointegrated over the time period of 1930 through 1989. The cointegrating relationship is negative, as theory suggests, and that a one-time shock to interest rates has, *ceteris paribus*, a permanent, negative effect on underwriting margins. Also, there is some evidence, although preliminary, that equilibrium activity becomes more volatile at higher interest rates.

Figure 2



Responses of property-liability industry underwriting margin to a unit shock in the 90-day Treasury Bill rate.

Although these results do not prove that the industry is guided by competitive behavior, they are nevertheless consistent with rational, competitive behavior (Granger, 1986; Campbell and Schiller, 1987). In addition, the evidence presented here does not eliminate the possible existence of aggregation bias (Fields and Venezian, 1989). But the results do suggest that the long-term equilibrium prices are dependent upon the interest rate and that prices move toward equilibrium.

References

A. M. Best Company, 1959, 1990, *Best's Aggregates and Averages, Property-Casualty* (Oldwick, N.J.: A.M. Best).

Andrews, Donald W. K., 1987, Least Squares Regression with Integrated or Dynamic Regressors Under Weak Error Assumptions, *Econometric Theory,* 3: 98-116.

Berger, Lawrence A., 1988, A Model of the Underwriting Cycle in the Property/Liability Insurance Industry, *Journal of Risk and Insurance,* 55: 298-306.

Campbell, John Y. and Robert J. Shiller, 1987, Cointegration and Tests of Present Value Models, *Journal of Political Economy,* 95(5).

Copeland, Thomas E. and J. Fred Weston, 1983, *Financial Theory and Corporate Policy,* Second Edition (Reading, Mass.: Addison-Wesley).

Cummins, J. David, 1988, Risk-Based Premiums for Insurance Guaranty Funds, *Journal of Finance,* 43: 823-839.

Cummins, J. David, 1990, Asset Pricing Models and Insurance Ratemaking, *Astin Bulletin,* 20(2): 125-166.

Cummins, J. David, 1991, Statistical and Financial Models of Insurance Pricing and the Insurance Firm, *Journal of Risk and Insurance,* 58: 261-302.

Cummins, J. David and Scott E. Harrington, eds., 1987, *Fair Rate of Return in Property-Liability Insurance* (Boston: Kluwer-Nijhoff).

Cummins, J. David and François Outreville, 1987, An International Analysis of Underwriting Cycles, *Journal of Risk and Insurance,* 54: 246-262.

Dickey, David A. and Wayne A. Fuller, 1979, Distribution of the Estimators for Autoregressive Time Series with a Unit Root, *Journal of the American Statistical Association,* 74: 427-431.

Doherty, Neil A. and James R. Garven, 1986, Price Regulation in Property-Liability Insurance: A Contingent-Claims Approach, *Journal of Finance,* 41: 1031-1050.

Doherty, Neil A. and Han Bin Kang, 1988, Interest Rates and Insurance Price Cycles, *Journal of Banking and Finance,* 12: 199-214.

Engle, Robert F. and C. W. J. Granger, 1987, Co-Integration and Error Correction: Representation, Estimation, and Testing, *Econometrica,* 55: 251-276.

Engle, R. F. and C. W. J. Granger, eds., 1991, *Long-Run Economic Relationships, Readings in Cointegration* (Oxford: Oxford University Press).

Engle, Robert and Byung Sam Yoo, 1987, Forecasting and Testing in Cointegrated Systems, *Journal of Econometrics,* 35: 143-159.

Fairley, William B., 1979, Investment Income and Profit Margins in Property-Liability Insurance: Theory and Empirical Results, *Bell Journal of Economics,* 10: 192-210.

Fields, Joseph A. and Emilio C. Venezian, 1989, Interest Rates and Profit Cycles: A Disaggregated Approach, *Journal of Risk and Insurance,* 56: 312-319.

Fuller, Wayne A., 1976, *Introduction to Statistical Times Series* (New York: John Wiley).

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Granger, Clive, 1986, Developments in the Study of Cointegrated Economic Variables, in: R. F. Engle and C. W. J. Granger, eds., *Long-Run Economic Relationships, Reading in Cointegration* (Oxford: Oxford University Press).

Hakkio, Craig S. and Mark Rush, 1991, Cointegration: How Short Is the Long Run? *Journal of International Money and Finance,* 10: 571-581.

Hansen, Bruce E., 1992, Heteroskedastic Cointegration, *Journal of Econometrics,* 54: 139-158.

Hill, Raymond D., 1979, Profit Regulation in Property-Liability Insurance, *Bell Journal of Economics,* 10: 172-191.

Jarrow, Robert A. and Andrew Rudd, 1983, *Option Pricing* (Homewood, Ill.: Richard D. Irwin).

Judge, George G. et al., 1985, *The Theory and Practice of Econometrics,* Second Edition (New York: John Wiley).

Kraus, Alan, and Stephen A. Ross, 1982, The Determination of Fair Profits for the Property-Liability Insurance Firm, *Journal of Finance,* 53: 1015-1028.

Lütkepohl, Helmut and Hans-Eggert Reimers, 1992, Impulse Response Analysis of Cointegrated Systems, *Journal of Economic Dynamics and Control,* 16: 53-78.

Myers, Stewart C. and Richard A. Cohn, 1987, A Discounted Cash Flow Approach to Property-Liability Insurance Rate Making, in: J. David Cummins and Scott E. Harrington, eds., *Fair Rate of Return in Property-Liability Insurance* (Boston: Kluwer-Nijhoff).

Smith, Michael L., 1989, Investment Returns and Yields to Holders of Insurance, *Journal of Business,* 62: 81-98.

Stewart, Barbara D., 1981, The Economics of the Underwriting Cycle, *Journal of Insurance,* 42: 2-6.

U.S. Department of Economics Analysis, 1955, 1969, 1989, *Business Statistics, A Supplement to the Survey of Current Business* (Washington, D.C.: U.S. Government Printing Office).

Venezian, Emilio C., 1985, Ratemaking Methods and Profit Cycles in Property and Liability Insurance, *Journal of Risk and Insurance,* 52: 477-500.

Wilson, William C., Jr., 1981, The Underwriting Cycle and Investment Income, *CPCU Journal,* 34: 225-232.

Winter, Ralph A., 1988, The Liability Crisis and the Dynamics of Competitive Insurance Markets, *Yale Journal on Regulation,* 5: 455-499.

Winter, Ralph A., 1991, The Liability Insurance Market, *Journal of Economic Perspectives,* 5(3): 115-136.