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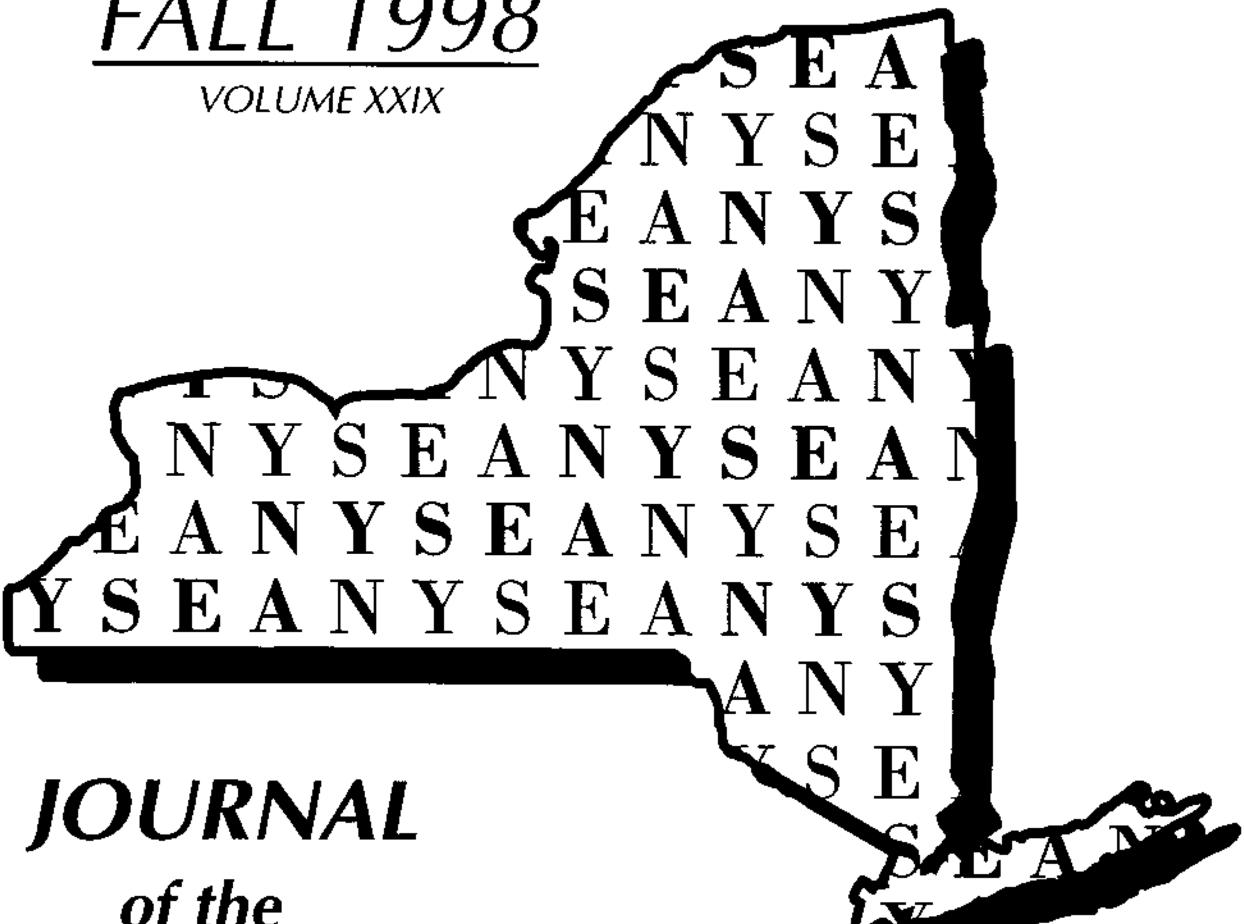

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RELATIVE MAGNITUDES OF RISK AVERSION AND PRUDENCE

Joseph G. Eisenhauer*

INTRODUCTION

Risk aversion and prudence are fundamental descriptions of preferences in expected utility theory. Whereas risk aversion measures the propensity to avoid risk, prudence measures the propensity to take precautions when faced with risk. The modern literature on behavior under uncertainty is largely built around these concepts, and as Eeckhoudt and Schlesinger (1994a, p. 53) have recently observed, "Given the many applications of prudence that are appearing in expected utility models, the relationship between prudence and risk aversion becomes quite important." Empirical measurement, however, has lagged well behind theoretical advances and applications in this area. The present paper estimates the relative magnitudes of risk aversion and prudence for the average U.S. household using an expected utility model of life insurance demand. The theoretical model is developed after a brief introduction to the literature. The empirical results follow, and the paper ends with a brief conclusion.

LITERATURE

The measures of absolute and relative risk aversion originated in the work of John Pratt (1964) and Kenneth Arrow (1965), who independently postulated the hypotheses of decreasing absolute risk aversion (DARA) and increasing relative risk aversion (IRRA). According to these hypotheses, the willingness to wager a fixed sum increases, but willingness to wager a given proportion of wealth declines, as financial resources rise. Arrow (1965) provided a mathematical foundation for IRRA, but simply asserted without evidence (p. 35) that DARA "certainly seems supported by everyday observation." And as Campbell (1980 p. 1166) pointed out, "Most researchers seem to agree that the most plausible theoretical form of absolute risk aversion is a decreasing function of its argument." Indeed, according to Laffont (1989, p. 24), some researchers even dispute the need to test the proposition, arguing that, "since we must assume that absolute risk aversion decreases with wealth to obtain results that accord with intuition and observations of rational behavior...we can infer that agents must satisfy this assumption in general."

Shortly after the seminal contributions of Pratt and Arrow appeared, Hayne Leland (1968) observed that risk aversion is insufficient to induce precautionary saving, and demonstrated that convex marginal utility is the requisite condition. Although as Deaton (1992) points out, many researchers find the convexity difficult to justify intuitively, Kimball (1987, 1990 p. 54) used the convexity to establish the measure of prudence, choosing that term to indicate "the propensity to prepare and forearm oneself in the face of uncertainty." The role of prudence has since become increasingly prominent in applied work. Blanchard and Mankiw (1988), for example, show how the intertemporal consumption function depends on the magnitude of prudence, and Gollier, Jullien, and Treich (1997) show that prudence must be twice as great as risk aversion for the "precautionary principle" embodied in recent international environmental

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treaties to be efficient.[1] Moreover, as Kimball (1990, p. 65) observed, "absolute prudence is greater than or less than absolute risk aversion depending on whether absolute risk aversion is decreasing or increasing." Thus, the DARA hypothesis implies that prudence exceeds risk aversion.

Yet, while several studies have estimated the magnitude of relative risk aversion (Hansen and Singleton, 1982 and 1983; Szpiro, 1986), empirical research on prudence and absolute risk aversion is rather sparse. Studies by Grossberg (1991) and Kuehlwein (1991), for example, failed to observe precautionary saving and by implication, failed to find evidence of prudence. Similarly, the few empirical studies which have examined the DARA hypothesis have yielded mixed results. Due to the difficulty of obtaining accurate data, Wolf and Pohlman (1983, p. 843) explain that frequently, "efforts to classify an individual's risk preferences have been confined to direct assessments in hypothetical environments." A recent example is the study by Levy (1994) which found evidence of DARA among MBA students in a classroom investment experiment. But of course, graduate students may not be a representative subset of the general population. Moreover, such experiments may have limited relevance to real-world applications, especially when participation is voluntary and therefore subject to self-selection bias. Indeed, if the most risk averse individuals elect not to participate, the remaining participants are the least risk averse of the population at hand, and may be less risk averse than prudent; a voluntary experiment is therefore likely to be biased in favor of excessive prudence.

In a non-experimental study, Wolf and Pohlman (1983, p. 847) examined the bidding behavior of one securities broker and concluded, "As indicated by the likelihood ratio statistics...the hypotheses that the bidding data were produced by either increasing or constant absolute risk aversion cannot be rejected." It is, of course, difficult to draw general conclusions from the behavior of a single individual. But in a contemporaneous study, Szpiro (1983a) compared aggregate purchases of non-life insurance across countries and found only limited support for the DARA hypothesis. The cross-section approach, however, invokes a rather dubious assumption of similar preference structures across nations, and requires internationally consistent units of measurement for each of the variables involved. Moreover, it is often difficult to distinguish income from wealth in non-life insurance contexts. To overcome these difficulties, Eisenhauer (1997) advocates the use of time series data on life insurance within a single population.

The present research extends that work in several ways. First, we indicate the relative magnitudes of prudence and risk aversion implied by the DARA hypothesis, and argue (on the basis of IRRA) that DARA is unlikely to hold at high levels of financial resources. An explicit theoretical solution to the problem of optimal life insurance coverage is then derived and estimated using labor income, wealth, a pricing factor, and several additional control variables. The results are used to determine whether the average household is prudent, and if so, whether it is as prudent as implied by the DARA hypothesis. Both linear and nonlinear frameworks are utilized as a means of investigating the robustness of the results. The findings suggest that the average household exhibits a positive degree of prudence, but contrary to the prevailing theory, is less prudent than risk averse.

THEORY

Consider a household's concave von Neumann-Morgenstern utility function $U(X)$, such that $U'(X) > 0$ and $U''(X) < 0$. The Pratt-Arrow measure of absolute risk aversion is given by $A(X) = -U''(X)/U'(X)$, and Kimball's measure of absolute prudence is given by $\eta(X) = -U'''(X)/U''(X)$. Simple differentiation of the former reveals that

$$\text{(1)} \qquad dA(X)/dX = A^2 - A\eta \gtreqless 0 \text{ as } A \gtreqless \eta .$$

Thus, the DARA hypothesis, $dA(X)/dX < 0$, implies that prudence exceeds risk aversion ($\eta > A$). This is of course an empirical question, but even on theoretical grounds one might question its plausibility as a general proposition. At the same time that he proposed DARA, Arrow (1965) posited an increasing relative risk aversion (IRRA) on mathematical grounds, where the coefficient of relative risk aversion is $R(X) = -XU''(X)/U'(X)$, or $R(X) = XA(X)$. Of course, an increasing $R(X)$ may coexist with increasing, constant, or decreasing $A(X)$. Importantly, however, these are measures of risk aversion "in the small" as Pratt (1964) termed them; that is, measures applicable to infinitesimal change. Thus, if an infinitesimal increase in X raises the product $XA(X)$ at the same time that it reduces $A(X)$, the reduction in $A(X)$ must be "less than infinitesimal," or virtually nonexistent. More precisely, differentiation of $R(X)$ and some simple algebra reveals

(2) $$dR(X)/dX = A + X(A^2 - A\eta \gtreqless 0 \text{ as } A + 1/X \gtreqless \eta .$$

Thus, while DARA implies $A < \eta$, IRRA implies $\eta < A + 1/X$. When X is large, therefore, prudence and absolute risk aversion must be nearly equal if both the relative and absolute hypotheses are to hold simultaneously. Indeed, in the extreme case of infinite X, IRRA would require absolute risk aversion to exceed absolute prudence, contradicting the DARA hypothesis. Thus, we might anticipate rejection of the DARA hypothesis for extremely wealthy households.

To obtain an empirically measurable model of absolute risk aversion and prudence, we consider a life insurance problem. Let Y represent the present discounted value of expected future earnings, conditional on there being no loss. Assume there is a probability p of the breadwinner's death causing the loss of Y, but the accumulated stock of nonhuman assets, or current wealth (W) is not subject to the same risk. Life insurance coverage (V) is available at a premium rate, or per-dollar cost of coverage m, where $m = \lambda p$, $0 < m < 1$, and $\lambda \geq 1$; the total premium mV is paid regardless of the state of nature. The household chooses V to maximize the expected utility function

(3) $$E(U) = (1\text{-}p)U_N(W+Y\text{-}mV) + pU_L(W+V\text{-}mV)$$

where the subscripts N (no loss) and L (loss) denote favorable and unfavorable states, respectively. This conventional single-period model is adopted for simplicity, but because a rational, forward-looking consumer takes all known information into account in the initial purchase decision, the central results extend to multiple periods; see Campbell (1980) for a comparison of models. In addition, the ability to distinguish clearly between wealth and income makes the life insurance model somewhat unique among single period models, and largely obviates the need for multiple periods.[2]

To derive an explicit, albeit approximate, solution to the problem of optimal coverage, it is necessary to utilize a Taylor series expansion of expected utility.[3] Expanding (3) around the point of full coverage ($V = Y$) and writing $X = W + Y - mY$ gives

(4) $$E(U) \approx (1\text{-}p)U(X)\text{-}m(1\text{-}p)(V\text{-}Y)U'(X)+(1/2)m^2(1\text{-}p)(V\text{-}Y)^2U''(X)$$
$$+pU(X)+(1\text{-}m)p(V\text{-}Y)U'(X)+(1/2)1\text{-}m)^2p(V\text{-}Y)^2U''(X)$$

where third-order and higher terms are dropped. Differentiating with respect to V yields the first-order condition

(5) $$[m^2(1\text{-}p)+(1\text{-}m)^2p](V^*\text{-}Y)U''(X)=(m\text{-}p)U'(X)$$

and the concavity of the utility function ensures that the second-order condition holds. By rearranging (5), an explicit solution for optimal coverage can be obtained as

(6) $V^* = Y - \theta/A$

where $\theta = (m-p)/[m^2(1-p) + (1-m)^2p] \geq 0$.

Equation (6) expresses optimal life insurance coverage as an explicit function of the parameters of the model and has a clear intuitive interpretation. For any given value of expected future income, Y, life insurance coverage increases with absolute risk aversion ($\partial V^*/\partial A > 0$). Moreover, with an actuarially fair premium ($m = p$), full coverage ($V^* = Y$) would be optimal; however, optimal coverage is less than the present discounted value of potential future income ($V^* < Y$) if the insurance premium is loaded ($\lambda > 1$, or equivalently, $m > p$). In the data below, $1.6 < \lambda < 2.3$.[4] Thus, partial coverage inevitably prevails, and resources are greater in state N (no loss) than in state L (loss), even with the optimal level of insurance coverage in place. Furthermore, for a household exhibiting constant absolute risk aversion ($dA/dX = 0$), life insurance coverage would be strictly proportional to income ($\partial V^*/\partial Y = 1$).[5]

Most importantly for present purposes, differentiating (6) with respect to W and applying (1) yields

(7) $\partial V^*/\partial W = \theta(1-\eta/A)$

from which we obtain

(8a) $\partial V^*/\partial W \gtreqless \theta \text{ as } 0 \gtreqless \eta$

and

(8b) $\partial V^*/\partial W \gtreqless 0 \text{ as } A \gtreqless \eta$

Thus, by (8a) the marginal effect of wealth on life insurance coverage exceeds, equals, or is less than the pricing parameter θ as the policyholder exhibits imprudence, nonprudence, or prudence, respectively. By (8b), life insurance coverage increases, remains constant, or decreases with wealth as absolute risk aversion exceeds, equals, or is less than absolute prudence, respectively. Hence, $\partial V^*/\partial W > 0$ implies $\eta < A$ (or increasing absolute risk aversion), $\partial V^*/\partial W = 0$ implies $\eta = A$ (or constant absolute risk aversion), and $\partial V^*/\partial W < 0$ implies $\eta > A$ (or decreasing absolute risk aversion). Moreover, estimation of $\partial V^*/\partial W$ and calculation of θ allows an explicit estimation of η/A, the relative magnitudes of prudence and risk aversion.[6] Note that since relative prudence is given by $\rho = X\eta$, the ratio of relative prudence to relative risk aversion is equal to the ratio of absolute prudence to absolute risk aversion: $\rho/R = \eta/A$. Thus, although our analysis is conducted for the absolute measures, it is equally applicable to the relative measures.

DATA AND EMPIRICAL RESULTS

Although time series data disaggregated by household and covering each of the variables specified by the model would clearly be the most suitable basis for an empirical test of this sort, such data are not readily available at present. Consequently, most recent studies of life insurance demand have utilized either aggregate cross-section data (Browne and Kim, 1993), cross-section micro data (Fitzgerald, 1987), or aggregate time series data (Truett and Truett, 1990). As noted above, however, cross-section research assumes interpersonal consistency of utility whereas time series studies invoke a less problematic assumption of intertemporally consistent utility. Thus, the present empirical test uses aggregate U.S. time series data for 1970-1992. The data were obtained from the American Council of Life Insurance (1994), the Council of Economic Advisors (1994), and the annual *Statistical Abstract of the United States*. Here, as in much of the literature on consumer behavior, "aggregate consumption is treated as if it had been generated by the decision processes of a single agent" (Deaton, 1992, p. 37). Because the relevant variables are assumed to be known to all households, consistent aggregation follows if we assume stable preference patterns over time. As Deaton (1992, p. 42) notes, aggregate behavior approximates individual behavior "provided macro variables are included in the individual information sets and provided...that the dying households are replaced by look-alike young ones."[7] Moreover, because individuals are often reluctant to reveal their personal finances, aggregate data may yield an efficiency gain; as Langbein and Lichtman (1978, p. 33) note, "If individual level data are less complete or less reliable than a complementary set of aggregate data, using the aggregate data may be preferable, even though individual level behavior is at issue."[8]

For consistency with the theoretical model, and to control for the number of households, the aggregate data were reduced to per-household units. Thus, the dependent variable LIFEINS measures inflation-adjusted life insurance coverage per household, while the independent variables INCOME and WEALTH are inflation-adjusted, per household measures of disposable personal income and financial wealth, respectively. Assuming that the growth rate of income equals the intertemporal discount rate, the present value of future income is simply a multiple of current income. Using current income as a proxy for the present value of future income therefore inflates the regression coefficient but does not alter its sign.[9] The premium loading factor (LOAD) captures price effects and should therefore exert a negative influence on the demand for insurance. The loading factor is used to avoid collinearity between the premium rate and the age-adjusted mortality rate. Inspection of the correlation matrix reveals no severe collinearity between income and wealth. Because the number of dependents (and thus, the size of the household) might be expected to exert a positive influence on the amount of life insurance purchased, the average number of persons per household (HHSIZE) is included as a control variable.[10] Likewise, life insurance is more likely to be purchased by working-age individuals than by children or the elderly. The percentage of the population aged 25 to 64 (AGE) is therefore included as an additional control variable, and its coefficient is expected to be positive.

Summary statistics for each variable are provided in Table 1. While inspecting Table 1, it is important to bear in mind that the distributions of both income and wealth exhibit a strong positive skewness in the U.S. (as they do elsewhere). Consequently, the average household should not be interpreted as a representative unit, being in fact relatively wealthy in comparison with the median and modal households. The mean values of LIFEINS and INCOME provided in the table indicate that the average American household carries an amount of life insurance equal to twice its current annual income. This is clearly less than the present discounted value of potential future income, and is therefore consistent with the partial coverage implication of the model above, given the average loading factor of 2.0 over the period in question.

The appropriate empirical specification for estimating the partial derivative in equation (7) is simply a linear regression. To investigate the robustness of the results, however, log-linear and semi-logarithmic models were also estimated, with corrections for first-order autoregressive serial correlation where necessary. The semi-log model regressed LIFEINS on the natural logs of the independent variables. The regression results are given in Table 2, where two-tailed prob-values are shown in brackets below the estimated coefficients. Despite their simplicity, each of the models explains more than 95 percent of the variation in life insurance coverage, and the results are consistent in terms of the signs and significance of the parameter estimates. The coefficients of INCOME, LOAD, HHSIZE, and AGE all have the expected signs, although the loading factor is not statistically significant and is therefore removed from the second iteration of each model. The positive effect of INCOME suggests that life insurance is a normal good, while the insignificance of LOAD suggests that life insurance demand is essentially insensitive to price changes.

TABLE 1
SUMMARY STATISTICS

VARIABLE	MINIMUM	MAXIMUM	MEAN	STANDARD DEVIATION
LIFEINS	$50825.10	$75980.00	$60308.13	$8607.43
INCOME	$28808.30	$33741.40	$31136.41	$1458.62
WEALTH	$68230.90	$91128.30	$79116.18	$7020.50
LOAD	1.59591	2.293370	2.0267935	0.2478925
HHSIZE	2.6647	3.2377	2.8749435	0.1834585
AGE	0.44023	0.51239	0.4782204	0.0249927

TABLE 2
REGRESSION RESULTS FOR LIFE INSURANCE DEMAND*

VARIABLE	LINEAR MODEL		LOG-LINEAR MODEL		SEMI-LOG MODEL	
CONSTANT	-377466	-381644	-2.6038	-2.93625	-776271	-802195
	[.0003]	[.0001]	[.3559]	[.2673]	[.0007]	[.0003]
INCOME	1.99303	2.01436	0.99551	1.00781	60606.07	61511.4
	[.0012]	[.0007]	[.0020]	[.0012]	[.0031]	[.0022]
WEALTH	0.23400	0.24370	0.33417	0.35533	21280.92	23086.8
	[.0113]	[.0037]	[.0084]	[.0019]	[.0093]	[.0020]
LOAD	-816.84		-0.06087		-6447.92	
	[.8583]		[.6899]		[.5159]	
HHSIZE	44664.3	44438.4	2.18947	2.21898	124788	128886
	[.0019]	[.0013]	[.0047]	[.0030]	[.0092]	[.0062]
AGE	481857	485493	3.69438	3.83944	212992	230091
	[.0005]	[.0001]	[.0016]	[.0004]	[.0031]	[.0008]
R^2	.9731	.9737	.9692	.9691	.9659	.9638
Adj. R^2	.9635	.9667	.9582	.9609	.9532	.9541
DW	2.0295	2.0785	2.0221	2.0524	2.0406	2.0116

*Two-tailed prob-values given in brackets below estimated coefficients.

Most importantly, the coefficient of WEALTH is positive and highly significant; both $\eta > A$ and $\eta = A$ are therefore rejected in favor of $\eta < A$ at the one percent level of significance. Thus, the evidence of the U.S. life insurance market suggests that the average household exhibits increasing absolute risk aversion and is therefore more risk averse than prudent, at least within the range of real wealth values observed over the period, \$68,230 to \$91,128 in 1983 prices. To be more precise, the wealth effect from the linear model is approximately 0.24, and our mean value of $\lambda = 2$ implies $m = 2p$, or equivalently, $\theta = 1$ on average. Since $\partial V^*/\partial W < \theta$, we know the average household exhibits prudence ($\eta > 0$), and substituting these values into equation (7) generates a point estimate of $\eta/A = .76$. Thus, the average household appears to exhibit only three-fourths as much prudence as risk aversion.

Moreover, the outcome is robust: the linear, logarithmic, and semi-logarithmic specifications yield similar results. In particular, the log-linear model yields income, wealth, and price elasticities of 1.00, 0.33, and -0.06 respectively, although as noted above the price elasticity is not statistically significant. In addition, these findings are consistent with empirical results obtained elsewhere in the life insurance literature. Although previous studies of life insurance demand were not designed to estimate the relative magnitudes of risk aversion and prudence, several studies which included the value of noncontingent financial assets as an independent variable found positive wealth effects (Ferber and Lee, 1980; Fitzgerald, 1987).

Of course, most households also possess claims to contingent wealth in the form of Social Security retirement and survivor benefits. The former are approximately three times as great as the latter; during the period in question, retirement benefits routinely constituted three-quarters of all Old Age and Survivors Insurance (OASI) annuities, while survivor benefits constituted the remaining quarter. Indeed, Feldstein (1974, p. 905) has suggested that "For the great majority of Americans, the most important form of household wealth is the anticipated social security retirement benefits." But retirement benefits are based on earnings and contingent upon the worker's survival; Feldstein himself estimated the present value of Social Security benefits as a function of current disposable income and the probability of survival. Consequently, what is often called "Social Security wealth" in the literature is actually consistent with the definition of income (Y) in the model above, and fundamentally different from wealth (W) as defined by the stock of noncontingent assets. Thus, if retirement benefits were included in the regression equation, we could expect them to have a positive coefficient, but they would almost surely exhibit collinearity with income and thereby distort the empirical results. On the other hand, Social Security survivor benefits are contingent upon the breadwinner's death, and serve as a direct substitute for privately purchased life insurance (Fitzgerald, 1987). It is not surprising, therefore, that studies which used survivor benefits as a proxy for wealth obtained negative wealth effects (see, for example, Lewis, 1989).[11]

CONCLUSION

The results of this study suggest that the average household exhibits a positive degree of prudence, but less prudence than implied by the DARA hypothesis; specifically, we estimate that the average household is approximately one and one-third times more risk averse than prudent. Several caveats are, however, in order. The present data do not distinguish between different types of life insurance policies, such as term and whole life.[12] Similarly, while the number and average size of households are controlled, the present study does not incorporate changes in the composition of households (such as differences in education or the head-of-household gender) over time. Future research with disaggregated data might profitably explore these aspects of the issue. There is, however, no *a priori* reason to believe that such considerations would significantly alter the results obtained here. Indeed, the present results are largely consistent with those of Gandolfi and Miners (1996), who distinguished between individual and group policies and between purchases by husbands and wives.

Additionally, the findings are quite reasonable given that the mean wealth levels observed in this study are relatively high in comparison with the median and modal levels of wealth. However, because they pertain to a household of mean size with mean income and wealth, the results do not necessarily apply to households elsewhere in the socioeconomic spectrum. Thus, although the average household appears to exhibit greater risk aversion than prudence, these data do not permit such a conclusion to be drawn for the population in general.

The importance of this research is suggested by the growing body of applied and often policy-oriented work which relies on assumptions regarding absolute risk aversion and prudence. The rejection of DARA, for example, could justify the employment of quadratic utility functions which are useful in financial mean-variance analyses, but which have fallen out of favor precisely because of their inconsistency with DARA. Of course, no single test can be considered conclusive; but the present findings suggest that the question is not yet settled. Certainly, then, further empirical testing with nonexperimental and preferably disaggregated data is needed. As Eeckhoudt and Schlesinger (1994, p. 5) have argued, "As the concept of prudence finds a wider range of applications over time, a knowledge of its relationship to risk aversion should prove more and more useful."

ENDNOTES

1. The precautionary principal involves action to prevent future damage (as to the environment) under uncertainty. See O'Riordan and Cameron (1995).
2. An alternative (though unfashionable) justification for a single period model is simple myopia: consumers may simply have an extremely short planning horizon caused, for example, by bounded rationality or liquidity constraints; see, for example, Kurz (1990). In that case, Y would be interpreted as the conditional expectation of next-period income, rather than the conditional expectation of all future income.
3. Eisenhauer (1997) presents an exact but implicit solution, from which the η/A ratio cannot be estimated. A derivation similar to the one used here is presented by Szpiro (1986) for property-liability insurance, and Szpiro (1983b) shows that the resulting expressions can be aggregated.
4. Because loading factors were not reported directly, data from the American Council of Life Insurance (1994) on premiums, coverage, and age-adjusted mortality rates were used to calculate the loading factors as the ratio of premiums to expected claims: $\lambda = (mV)/pV$. The mean and standard deviation are given in Table 1. The range of computed factors is consistent with the range reported by Lewis (1989), $1.2 < \lambda < 2.7$.
5. Note that with disaggregated data on expected future earnings, mortality rates, premium rates, and insurance coverage, the magnitude of absolute risk aversion could be computed directly from (6).
6. A similar analysis could, in principal, be conducted on the basis of the income effect. However, the empirical estimation uses current income as a proxy for potential future income and thus distorts the magnitude of the income effect as described in footnote 8.
7. Moreover, the increasing prevalence of group policies, such as those purchased by employers, provides some conceptual justification for beginning with aggregate data and averaging across households. While individual purchases are still predominant, group coverage expanded from 30.6 percent of life insurance in 1961 to 40.6 percent in 1991.
8. Similarly, Grunfeld and Griliches (1960, p. 1) have argued that "aggregation of economic variables can, and in fact frequently does, reduce these specification errors. Hence, aggregation does not only produce an aggregation error, but may also produce an aggregation gain."
9. To see this, let INC, g, r, and Y be current income, its (constant) growth rate, the discount rate, and the present discounted value of potential future income, respectively. Then if g = r,

   $$Y = \sum_{t=1}^{n} [INC(1+g)^t/(1+r)^t] = nINC \ .$$

   Thus, replacing βY by bINC (=bY/n) implies $b = n\beta$; the regression coefficient is inflated by a factor of n, but its sign is unchanged.
10. Goldsmith (1983, p. 40) contends, "On theoretical grounds, the impact of household size on [life insurance purchases] is ambiguous, because the increase in expected household earnings due to an additional dependent is opposed by an increase in the cost of supporting a larger household." But since household income is controlled, household size should exert a strictly positive effect on optimal life insurance coverage.
11. Goldsmith (1983) also obtained a negative wealth effect using a homeownership indicator as a proxy for wealth, but Gandolfi and Miners (1996) found a significant positive effect of homeownership. In general, however, homeownership appears to be a poor, if not counterproductive proxy; not only does this measure exclude financial wealth, but the purchase of a home converts liquid assets to illiquid assets and initially *reduces* total wealth through closing costs and mortgage points.
12. During the period in question, between 41 percent and 60 percent of ordinary life insurance purchases were term policies.

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A MODEL OF MARKET STRUCTURE AND INCENTIVES TO IMPROVE PRODUCT QUALITY

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An important problem in industrial economics is whether changes in market structure lead to differences in firm incentives to improve product quality. Two types of models have emerged to address this issue. The work of Lee and Wilde (1980), Dasgupta and Stiglitz (1980), and Delbono and Denicolo (1991) approach the problem in a game theoretic framework by assuming that participants are awarded an exogenous prize. The incentives to play via R&D effort are linked to market structure.

More recent work by Moorthy (1985), Goel (1990), Caminal and Vives (1996), Wauthy (1996) and Anonymous (1997) have improved on a shortcoming of the earlier studies by explicitly linking firm quality efforts to consumer preferences or demand behavior. A problem with these later models is that in focusing on the relationship between quality incentives and demand structure, the authors in general make the assumption that the firm incurs no direct cost from improving quality other than marginal production costs. This assumption neglects a substantial production economics literature, for instance, Lundvall and Juran (1974), Chandra and Ullmann (1986), Modress and Ansari (1987), and Merino (1990), that focuses on the cost of impoving product quality.

The present paper develops a model of the relationship between a firm's incentive to improve product quality and its market structure in an attempt to bridge the above literature. The model assumes an arbitrary number of firms and is general enough to cover competitive, monopoly, Cournot, and Stackleberg behavior. Quality costs are also directly introduced into the firm's decisions. The model is a generalization of the models of Caminal and Vives (1996), Wauthy (1996), and Anonymous (1997), since quality costs are explicitly considered. The model extends the production economics literature by explicitly introducing market structure and demand into the quality decision.

The basic conclusion of the model is that a firm's choice of an optimal level of quality will consist of weighing the incremental cost of an improvement in product quality against the revenues generated by this improvement. A firm's ability to generate this revenue will be conditioned by both market share and the response of rival firms. The model shows that as the number of firms in the industry increases, potential revenue from quality improvement decreases. More interestingly, the model demonstrates that there is no difference in the ability to generate revenues from quality improvements for Stackleberg leader firms and Stackleberg followers. This latter result helps explain why in some industries such as brewing quality initiatives have come from the competitive fringe rather than the dominant firm. In the computer chip industry, however, an argument can be made that quality improvements have often come from the dominant firm.

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THE MODEL

Consider a situation in which there are N firms producing a product X. The price of X depends on market demand via the inverse demand function $P = P(X)$. Demand for the i^{th} firm's output will be a function of market price and the quality, q_i, of its product[1]. That is, $X_i = X(P, q_i)$.

Total costs for the i^{th} firm are given by:

$$(1) \qquad Tc_i = C_M X_i + C(q_i)X_i,$$

where C_M is the unit manufacturing cost and $C(q_i)$ is the unit cost of quality. Separating costs into manufacturing and quality costs and expressing costs on a per unit basis as in (1), has been done for simplicity[2]. However, a specification similar to this has been used by Leffler (1982), and is consistent with discussions of costs in the production economics literature [See Lundvall and Juran (1974), Chandra and Ullmann (1986), Modress and Ansari (1987), and Merino (1990)]. As a further simplification the model assumes identical cost functions for all firms. This allows the analysis to concentrate on the response of a representative firm.

The choice of optimal output and quality for the i^{th} firm occurs as follows. First, each firm chooses its profit-maximizing output. This results in a market price for the industry, which can be expressed as a mark-up over costs for the representative firm. Given this price, the i^{th} firm will shift its demand to generate additional output, revenue and profit by changing product quality.

Expressing profits for the i^{th} firm as ($\Pi_i = PX_i - TC_i$, and differentiating with respect to output results in the first-order condition

$$(2) \qquad \frac{\partial \Pi_i}{\partial X_i} = P + \frac{dP}{dX}\frac{\partial X}{\partial X_i}X_i - C_M - C(q_i) = 0.$$

Equation (2) reduces to the familiar marginal revenue equals marginal cost condition.

The change in market price with respect to the output change of the i^{th} firm will depend on price's responsiveness to market output and on how market output responds to firm i's output change. The change in market output will depend on both firm i's output change and on the response of firm i's rivals to its output change:

$$(3) \qquad \frac{\partial P}{\partial X_i} = \frac{dP}{dX}\frac{\partial X}{\partial X_i} = \frac{dP}{dX}\frac{\partial (X_i + \ldots + X_N)}{\partial X_i} = \frac{dP}{dX}\left(1+\sum \frac{\partial X_j}{\partial X_i}\right),$$

where the sum in the last expression in (3) is over all $j \neq i$. Substituting (3) into (2) results in the condition

$$(4) \qquad P\left[1+\frac{dP}{dX}\frac{X}{P}\left(1+\sum\frac{\partial X_j}{\partial X_i}\right)\frac{X_i}{X}\right] - C_M - C(q_i) = 0.$$

Using the definitions $\eta = -(dX/dP)X/P$ and $S_i = X_i/X$, we can solve (4) for P:

$$(5) \qquad P = \frac{\eta}{\eta - S_i\left(1+\sum\frac{\partial X_j}{\partial X_i}\right)}(C_M + C(q_i)).$$

Equation (5) states that the i^{th} firm views market price as proportional to costs, with the margin over costs conditioned by the price elasticity of market demand η, firm i's market share, S_i, and the total response of its rivals $\Sigma\partial X_j/\partial X_i$. The price defined in (5) will enter into the i^{th} firm's quality decision.

Expressing profits of firm i as $\Pi_i = (P - C_M - C(q_i))X_i$ and differentiating with respect to quality results in the first-order condition

$$(6) \qquad P - C_M - C(q_i))\frac{dX_i}{dq_i} = X_i\frac{dC(q_i)}{dq_i} \ .$$

Equation (6) states that optimal quality for the i^{th} firm must be adjusted such that the incremental net unit profit generated by the demand shift due to the increase in quality equals total incremental quality costs[3].

Using (5) in (6) results in

$$(7) \qquad X_i\frac{dC(q_i)}{dq_i} = P\frac{dX_i}{dq_i}\frac{S_i}{\eta}\left(1+\sum\frac{\partial X_j}{\partial X_i}\right) \ .$$

The left-hand side of (7) represents the incremental cost to quality improvement weighted by output. It is a measure of the costs incurred by devoting more resources to product quality. The right-hand side of (7) represents the revenue the profit-maximizing firm generates through quality improvement. According to (7), this revenue and the resulting net profit depends on market elasticity of demand, firm market share, and the response of its rivals. This equation is interesting in that it shows the incentives to increase product quality under alternative market structures.

In a perfectly competitive market, $S_i = 0$. This makes the right-hand side of (7) equal to zero. The equation suggests that there is no incentive for competitive firms to improve product quality. This would be expected in a market where the output of all producers are perfect substitutes, and the typical firm has no margin of price over marginal cost to cover the cost of quality improvement.

For the monopoly, $S_i = 1$ and $dX_j/dX_i = 0$ for all j. In this situation the incentive to increase product quality depends exclusively on market demand, and varies inversely with price elasticity η. This result is intuitive. Since the price mark-up and profit margin of high elasticity products tends to be small, there is a smaller incentive to improve the quality of such products.

For the firm operating in Cournot equilibrium rival responses are zero, so that $1 + \Sigma\partial X_j/\partial X_i = 1$. Additionally, the market share of all firms are equal, that is, $S_i = 1/N$ for all i. Proof of this latter condition is provided in Appendix A[4]. Given the two restrictions, equation (7) reduces to

$$(8) \qquad X_i\frac{dC(q_i)}{dq_i} = P\frac{dX_i}{dq_i}\frac{1}{N\eta} \ .$$

In a Cournot industry, all firms take price as given and choose output holding rival output constant. In this situation an increase in the number of firms decreases market price and the net margin of price over costs. Equation (8) suggests that as the number of firms in the industry increases, the resulting profit that can be generated by an increase in quality for the Cournot firm decreases, as does the incentive to improve product quality. This result is similar to Delbono and Denicolo (1991), who demonstrate that the incentive to innovate in a Cournot oligopoly decreases with the number of firms.

For the Stackleberg model we must consider conditions confronting both the leader and follower firm[5]. For the dominant firm, $1 + \Sigma\partial X_j/\partial X_i = 1/N$, while the market share is $S_D = N/(2N - 1)$. Condition (7) for the Stackelberg leader is

$$(9) \qquad X_i\frac{dC(q_i)}{dq_i} = P\frac{dX_i}{dq_i}\frac{1}{(2N-1)\eta} \ .$$

For each of the N - 1 follower firms, rival responses are zero, as in the Cournot solution. The market share for the typical follower is $S_F = 1/(2N - 1)$. Condition (7) for the Stackleberg follower also reduces to (9).

Equation (9) shows that the ability to generate revenues through product quality improvements is the same for both leader and follower firms in a Stackleberg industry. The dominant firm has a larger market share and potentially more sales revenue to cover the cost of quality improvements. However, this revenue advantage is offset by the fact that sales increases generated by the dominant firm's quality improvements will be countered by the reactions of the follower firms. The follower firms are in the opposite situation. None of the follower firms confronts rival firm response to its quality improvements. However, the revenue to finance quality improvements is smaller for each follower firm.

The Stackleberg model then suggests that once a firm has established itself as a *quantity leader* there is no necessary advantage to this leadership position in terms of its ability to generate revenues via product quality improvements. We believe that this result is consistent with casual empirical observation. In some industries, such as brewing, the argument can be made that major quality improvements have in many instances been initiated by the firms at the competitive fringe. We are referring to the introduction of new products by micro-breweries. In other industries, such as computer chips, quality improvements have been introduced by the dominant firms. An example is the continual upgrading of micro processors by Intel Corporation.

Equation (9) also shows that, similar to the Cournot firm, the resulting profit that can be generated by an increase in quality and the incentive to improve product quality decreases for the Stackleberg firm as the number of firms in the industry increases. This is again a result of the reduction in the net margin of price over costs as the number of firms in the industry increases. This is similar to a finding by Goel (1990), who shows that research and development efforts of Stackleberg leader firms tend to decrease as the competitive fringe expands.

Finally, for the condition $N \geq 2$, $1/N > 1/(2N-1)$ holds. A comparison of (8) and (9) shows that the ability to generate revenues through quality improvements is greater for the typical Cournot firm than for the Stackleberg firm. This suggests that in industries in which there is no dominant firm we might expect to see more quality improvements over time than in industries in which there is a definite leader-follower relationship. Casual observation suggests that this seems to be the situation with the automobile industry since the advent of global competition in the 1980s.

CONCLUSIONS

The present paper has been an attempt to develop a general model of the relationship between market structure and a firm's incentive to improve product quality. Profit maximization depends on a firm's choice of optimal output and optimal product quality given its mark-up of price over costs. The condition for optimal quality requires equating the marginal cost of improving quality with the marginal revenue of the quality improvement. This revenue depends on the elasticity of market demand, the firm's market share and the response of the firm's rivals. Special cases of the model reduced to the competitive, the monopoly, the Cournot, and the Stackleberg situations.

In constructing the model, we have made some obvious simplifications. First, the model is completely static. We have also assumed that in determining optimal quality the firm took market demand elasticity, its initial market share, and the responses of its rivals as given. In reality, each of these should be influenced by product quality. However, we believe these simplifications have been justified in constructing a more general model of the relationship between market structure and a firm's quality decisions than currently exists. Moreover, we hope that the predictions of the model are consistent with actual quality decisions made by firms under alternative market structures.

APPENDIX A: A Linear N-Firm Cournot Model

Assume there are N firms and that the inverse market demand is given by $P = a - b\Sigma X_i = a - b(X_1 + X_2 + \ldots + X_N)$. Suppose that for all i firms the total cost is $TC_i = TC$. Total profits of the i^{th} firm can be expressed as

(A1) $\Pi_i = [a - b(X_1 + \ldots + X_N)]X_i - TC.$

In Cournot equilibrium, each firm takes all other firms' output as given. This results in the following first order condition for the i^{th} firm

(A2) $$\frac{\partial \Pi_i}{\partial X_i} = a - bX_1 - bX_2 \ldots - 2bX_i \ldots - bX_N - MC = 0.$$

All n first-order conditions result in the following system of equations

(A3)
$$\begin{aligned} 2X_1 + X_2 \ldots\ldots\ldots + X_N &= \frac{a - MC}{b} \\ X_1 + 2X_2 \ldots\ldots\ldots + X_N &= \frac{a - MC}{b} \\ \ldots\ldots\ldots\ldots\ldots\ldots & \\ X_1 + X_2 \ldots\ldots\ldots + 2X_N &= \frac{a - MC}{b} \end{aligned}$$

This system reduces to

(A4)
$$\begin{aligned} 2X_1 + X_2 \ldots\ldots\ldots + X_N &= \frac{a - MC}{b} \\ -X_1 + X_2 &= 0 \\ \ldots\ldots\ldots\ldots\ldots\ldots & \\ -X_{N-1} + X_N &= 0 \end{aligned}$$

The last N-1 of these conditions imply that $X_1 = X_2 \ldots = X_N$. Substituting this condition into (A4) and solving for X_1 obtains

(A5) $$X_1 = X_2 \ldots = X_N = \frac{1}{N+1}\left(\frac{a - MC}{b}\right).$$

By (A5) the total market output for the N Cournot firms is

(A6) $$X = \Sigma X_i = \frac{N}{N+1}\left(\frac{a - MC}{b}\right),$$

with the share of the i^{th} firm being

(A7) $$S_i = \frac{X_i}{X} = \frac{1}{N}.$$

APPENDIX B: A Linear N-Firm Stackelberg Model

Assume there are N firms and that the inverse market demand is given by $P = a - b\Sigma X_i = a - b(X_1 + X_2 + \ldots + X_N)$. Suppose that for all i firms the total cost is $TC_i = TC$. Total profits of the i^{th} firm can be expressed as

(B1) $\Pi_i = [a - b(X_1 + \ldots.. + X_N)]X_i - TC.$

In a Stackleberg model, one firm is the leader firm, which we arbitrarily choose as firm 1, and other firms act as followers. We begin with behavior of the follower firms.

The typical follower firm acts as a Cournot firm, taking the output of all other firms including the leader as given. This results in the following first-order condition for firms 2 through N

(B2) $$\frac{\partial \Pi_i}{\partial X_i} = a - bX_1 - bX_2 \ldots.. - 2bX_i \ldots - bX_N - MC = 0.$$

All N-1 first-order conditions result in the following system of equations

(B3) $$\begin{aligned} &X_1 + 2X_2 \ldots\ldots\ldots + X_N = \frac{a-MC}{b} \\ &\ldots\ldots\ldots\ldots\ldots\ldots\ldots\ldots \\ &X_1 + X_2 \ldots\ldots\ldots\ldots + 2X_N = \frac{a-MC}{b} \end{aligned}$$

This system reduces to

(B4) $$\begin{aligned} &X_1 + 2X_2 \ldots\ldots\ldots + X_N = \frac{a-MC}{b} \\ &-X_2 + X_3 = 0 \\ &\ldots\ldots\ldots\ldots\ldots\ldots\ldots\ldots \\ &-X_{N-1} + X_N = 0 \end{aligned}$$

The last N-2 of these conditions imply that $X_2 = X_3 \ldots = X_N$. Substituting this condition into (A4) and solving for X_2 obtains

(B5) $$X_2 = \frac{1}{N}\left(\frac{a-MC}{b} - X_1\right).$$

The leader firm chooses output after taking the reactions of the followers into consideration. Expressing the leader firm's profit as $\Pi_1 = PX_1 - TC$, results in the following familiar marginal revenue equals marginal cost condition

(B6) $$P + \frac{\partial P}{\partial X_1}X_1 = MC.$$

The change in market price with respect to the leader firm's output can be evaluated by expressing this partial as the product of the slope of the demand curve and the change in market output with respect to the leader firm's output

$$\text{(B7)}\qquad \frac{\partial P}{\partial X_1}=\frac{dP}{dX}\frac{\partial X}{\partial X_1}=-b\frac{\partial[X_1+X_2...+X_N]}{\partial X_1}.$$

The change in market output with respect to the leader firm's output can be evaluated using reaction function (B5) and the equality of the N-1 follower firm's outputs

$$\text{(B8)}\qquad \frac{\partial[X_1+X_2...+X_N]}{\partial X_1}=\frac{\partial X_1}{\partial X_1}-\frac{(N-1)}{N}\frac{\partial X_1}{\partial X_1}=\frac{1}{N}.$$

Substituting (B8) into (B6) and using the definition of the inverse demand curve results in

$$\text{(B9)}\qquad a-b(X_1+X_2+....+X_N)-\frac{bX_1}{N}=MC.$$

Using (B5) in (B9) obtains the leader firm's output

$$\text{(B10)}\qquad X_1=\frac{1}{2}\left(\frac{a-MC}{b}\right),$$

while substituting (B10) into (B5) results in the typical follower firm's output

$$\text{(B11)}\qquad X_2=X_3=....=X_N=\frac{1}{2N}\left(\frac{a-MC}{b}\right).$$

From (B10) and (B11), it follows that total market output for the Stackleberg industry is

$$\text{(B12)}\qquad X=X_1+(N-1)X_2=\frac{2N-1}{2N}\left(\frac{a-MC}{b}\right).$$

The leader and typical follower market shares are therefore

$$\text{(B13)}\qquad S_1=\frac{X_1}{X}=\frac{N}{2N-1}$$

and

$$\text{(B14)}\qquad S_2=S_3=....=S_N=\frac{1}{2N-1}.$$

ENDNOTES

1. Note that in the present set-up market demand is not a function of product quality, while firm level demand is. This assumption has been made for simplicity, and makes sense under the following situation 1) each firm level's quality q_i is measured as a deviation relative to average product quality for the market; 2) there are relatively few substitutes for the product at the market level, so that overall product quality has little impact on market demand.
2. We also assume that unit manufacturing costs C_M are constant and that $C(q_i)$ is increasing in q_i. The assumption of constant marginal and unit manufacturing costs have been made for simplicity, while the assumption about quality costs is consistent with standard production economics theory. These restrictions assure that second order conditions for profit maximization are met provided that the marginal revenue of the representative firm is decreasing in output.
3. Equation (6) is similar to the condition for optimal quality derived by Dorfman and Steiner (1954) in their seminal work on optimal advertising and quality.
4. The proof in Appendix A is provided for the case of linear demand. Proof for the case of non-linear demand is similar.
5. The market shares for the dominant and follower firms in the Stackleberg model are derived for the case of linear demand in Appendix B. Proofs for the case of non-linear demand are similar.

PORTFOLIO BEHAVIOR OF LARGE U.S. CREDIT UNIONS UNDER UNCERTAINTY: A THEORETICAL EXPOSITION

Matiur Rahman* **and Douglas W. McNiel****

I. INTRODUCTION

To date, few attempts have been made to analyze the portfolio behavior of U.S. credit unions under uncertainty. Most previous studies have focused on output (loan) and portfolio management behavior in a nonstochastic environment [e.g., Cook and D'Antonio [1984]; Flannery [1974]; Smith [1984]; Smith, Cargill, and Meyer [1981]; Taylor [1971, 1977, 1979]; Walker and Chandler [1977,1978]]. The net income of a credit union is inherently stochastic for a variety of reasons including random variations in operating costs, regulatory changes and fluctuations in deposits, although credit unions can borrow from the central bank to partially offset a deposit outflow. Furthermore, they are likely to face stiffer competition from banks in the near future if a recent Supreme Court ruling that would limit most federal credit unions to new members of a single company or community stands [Felsenthal and Murray, 1998]. This ruling would affect almost 3,600 federally chartered credit unions that hold $132 billion in deposits from more than 32 million members. While the U.S. Congress and Senate may pass legislation to override the ruling, it adds to an uncertain business environment for U.S. credit unions. Therefore, analysis of the portfolio behavior of credit unions must take uncertainty into account. Smith [1988] uses a somewhat stylized model to incorporate uncertainty into the establishment of interest rates for loans and deposits. This paper employs the intermediation approach [Sealey and Lindley, 1977] to examine the portfolio behavior of credit unions within the analytical framework developed in Markowitz [1959] and Tobin [1958].

Credit unions in U.S. are retail-oriented mutual self-help organizations dating back to 1934. They can be either state or federally chartered. Workers from a similar environment or with the same employer pool their savings to provide loans to members for purchases of consumer durables or other consumption expenditures at interest rates lower than those charged by banks. Credit unions can offer relatively lower loan rates because they are tax-exempt. Although limited to retail deposit and lending activities, credit unions in some instances have become aggressive competitors of commercial banks. Currently, they have 70 million customers and their customer base is similar to that of commercial banks. Credit unions make up less than 6 percent of the consumer financial-services market but their assets have grown exponentially in recent years. Credit unions are now the fastest growing deposit-taking institutions in the U.S. The existence of credit unions forces banks to hold down their prices on some services.

Given the self-help and cooperative nature of credit unions, a question may arise as to what their objective is. Their primary focus is on the maximization of benefits of borrowing and lending members. Hence, they may or may not choose to maximize profits and/or minimize costs of production. Depositors and borrowers are also the owners. So the loan rate charged to the borrower, though generally revenue for the credit union, is also a cost to the borrower who is presumably an owner. Similarly, the rate paid to the depositor is a cost to the credit union but at the same time income to the depositor who is an owner of

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the entity. Under these circumstances, the loan and deposit rates may differ from those based on profit maximization and/or cost minimization considerations. Smith [1984] argues that, under general conditions, the objective of a credit union depends on the inherited balance sheet and whether it is dominated by borrowers or savers.

Under the Depository Institutions Deregulation and Monetary Control Act (DIDMCA), passed in 1980 and the Garn-St. Germain Depository Institutions Act (1982), the intensity of competition among banks and large credit unions has increased, and their degree of specialization has diminished. Moreover, in a world of unstable interest rates and severe competition from foreign banks and other financial institutions, the loans and deposits of financial institutions are becoming more uncertain [Bundt and Keating, 1988; Benston, 1983; Garcia et al., 1983]. Since credit unions cannot raise funds in the equity market, they may not be able to build adequate capital reserves to act as a cushion against uncertainty. While uncertainty is a problem for financial institutions in general, it is likely to be an especially severe problem for credit unions.

The primary purpose of this paper is to provide a mathematical explanation for the portfolio behavior of large U.S. credit unions in a situation of uncertainty. The study of credit unions under uncertainty is important because (i) this is an interesting and under-researched topic, (ii) large credit unions are gaining market share, (iii) their portfolio composition is undergoing changes due to deregulation and stiffer competition from banks and other depository institutions, and (iv) most of the research has focused on risk factors but has not addressed the uncertainty factor. The rest of the paper is structured as follows. Section 2 develops the model. Section 3 derives the comparative statics. Section 4 analyzes the effects, if any, of an increase in riskiness on the optimum combination of large credit unions' direct loans and securities investment. Finally, section 5 offers conclusions and remarks.

II. THE MODEL

There are ambiguities about the appropriateness of a general credit union objective function because of size differences of credit unions and their possible borrower/saver dominance. In order to maximize the monetary benefits allocated to a particular member group, dominated credit unions would restrict their size relative to neutral credit unions which attempt to maximize the benefits of all members [Taylor, 1971; Flannery, 1974]. In a classification based on total assets, neutral credit unions are found by Patin and McNiel [1991] to be the largest group, saver-dominated credit unions the second largest group and borrower-dominated credit unions the smallest group. Larger credit unions place restrictions on share ownership far less frequently than their smaller counterparts. There are several possible reasons for this, of which one is a desire to expand deposits. Thus, it seems likely that as a credit union increases in size its behavior approaches that of a profit-seeking firm [Flannery, 1974]. Recognizing that large and small credit unions might pursue different objectives, this paper focuses on large U.S. credit unions which in recent years have tended to behave increasingly like commercial banks. We thus use the expected utility approach where expected utility is presumed to be positively related to net operating income. Our model employs several simplifying assumptions. The only assets are reserves, loans and government securities. The level of deposits is assumed to vary randomly because of unforeseen events and unpredictable withdrawals by the member depositors.

Large credit unions seek to generate revenue through lending as well as investing in securities. The interest rate on loans (r_l) is assumed to be inversely related to the volume of loans outstanding (L). More specifically, $r_l = r_l(L)$; $r_l' < 0$ and $r_l'' > 0$. A credit union's total cost consists of interest payments on deposits and overhead expenses. The level of deposits (D) is determined by the deposit rate (r_d) which is exogenously given. Deposits are an increasing function of r_d, i.e., $D = D(r_d)$; $D' > 0$. Given these assumptions, a large credit union's net operating income (Y) function is specified as follows:

(1) $$Y = r_l L + r_s S - r_d D\beta - F$$

where r_S = rate of return on securities, S = amount of investment in securities, β = the random part of total deposits varying between 0 and 1 (i.e., $0 < \beta < 1$) and F = fixed cost. We assume that r_s is exogenoubly determined since even a large credit union is likely to account for a very small portion of total securities holdings in the US.

The expected utility function is given by

$$(2) \qquad \bar{U}(Y) = \int_0^\infty U(Y)f(\beta)d\beta; \; \bar{U}' > 0, \; \bar{U}'' < 0$$

The utility function is assumed to be continuous and at least twice differentiable. The credit union seeks to find the combination of L and S that maximizes $\bar{U}(Y)$ subject to the following balance sheet constraint:

$$(3) \qquad L+S = (1 - K)\, D + \bar{R}$$

where K is the reserve ratio and $\bar{R}$ is the capital reserve. To simplify the analysis, $\bar{R}$ is assumed to be exogenous.

The Lagrangian is:

$$(4) \qquad \mathscr{L} = \bar{U}(Y) + \lambda\,[\bar{R} + (1\text{-}K)\, D - L - S]$$

where λ is the Lagrange multiplier. Partially differentiating equation (4) with respect to S, L, and λ yields the following first-order conditions:

$$(5) \qquad \partial\mathscr{L} / \partial S = \bar{U}'(Y)\, \partial Y / \partial S - \lambda = \bar{U}' r_s - \lambda \equiv 0$$

$$(6) \qquad \partial\mathscr{L} / \partial L = \bar{U}'(y)\, \partial Y / \partial L - \lambda = \bar{U}'(r_l + Lr'_l) - \lambda \equiv 0$$

and

$$(7) \qquad \partial\mathscr{L} / \partial\lambda = \bar{R} + (1 - K)\, D - L - S \equiv 0$$

Solving equations (5) through (7), the optimal values of L, S, and λ are derived on a space of exogenous or pre-determined variables r_l, r_S and r_d, K and $\bar{R}$. In other words,

$\bar{S} = f(r_l,r_s,r_d,K,\bar{R})$, $\bar{L} = g(r_l,r_s,r_d,K,\bar{R})$ and $\bar{\lambda} = h(r_l,r_s,r_d,K,\bar{R})$.

However, the variables of our primary interest in this paper are S and L.

III. COMPARATIVE STATICS

This section of the paper examines the effects of changes in selected exogenous variables on the endogenous variables. Differentiating the first-order conditions further with respect to S, L, and λ yields the following system of equations, as presented in matrix form:

$$(8) \qquad \begin{bmatrix} a_{11} & a_{12} & a_{13} \\ a_{21} & a_{22} & a_{23} \\ a_{31} & a_{32} & a_{33} \end{bmatrix} \begin{bmatrix} \partial S \\ \partial L \\ \partial\lambda \end{bmatrix} = \begin{bmatrix} A \\ B \\ C \end{bmatrix}$$

where

$$a_{11} = \bar{U}''r_s^2,\ a_{12} = \bar{U}''r_s(r_l + r_l'L),\ a_{13} = -1$$

$$a_{21} = \bar{U}''(r_l + r_l'L)r_s,\ a_{22} = \bar{U}''(r_l + r_l'L)^2 + \bar{U}'(2r_l + r_l''L),\ a_{23} = -1$$

$$a_{31} = -1,\ a_{32} = -1, \text{ and } a_{33} = 0.$$

and the exogenous terms are $A = -U'dr_s$, $B = -U'dr_l$ and $C = -\bar{R} - (1-K)D$.

The determinant of the coefficient matrix is obtained as follows:

$$|H| = -2\bar{U}''r_s(r_l + r_l'L) + \bar{U}''r_s^2 + \bar{U}''(r_l + r_l'L)^2 + \bar{U}'(2r_l' + r_l''L) > 0.$$

Combining equations (5) and (6) yields $r_s = r_l + Lr_l'$. Using this relation, $|H| = (2r_l + r_l''L)U'$ which is positive if $2r_l > -r_l''L$, or equivalently, $2/L < -r_l''/r_l'$. It seems quite reasonable, provided $r_l'' > 0$ and the loan portfolio (L) is large.

Next, using Cramer's rule, we obtain

9(i) $$\frac{\partial S}{\partial r_l} = -U'/|H| < 0, \text{ and } \frac{\partial L}{\partial r_l} = U'/|H| > 0$$

9(ii) $$\frac{\partial S}{\partial r_s} = U'/|H| > 0, \text{ and } \frac{\partial L}{\partial r_s} = -U'/|H| < 0$$

9(iii) $$\frac{\partial S}{\partial k} = -D < 0, \text{ and } \frac{\partial L}{\partial K} = 0$$

and

9(iv) $$\frac{\partial S}{\partial \bar{R}} = -1, \text{ and } \frac{\partial L}{\partial \bar{R}} = 0$$

To offer case-by-case descriptions of the results, 9(i) implies that an increase in the loan rate will reduce a large credit union's investment in securities while raising the volume of loans to the credit union members. Equation 9(ii) shows that an increase in the rate of return on securities will increase securities investment and lower the volume of lending. According to 9(iii), an increase in the reserve ratio reduces the holding of securities and has no effect on the volume of loans. And, equation 9(iv) suggests that an increase in the capital reserve causes a dollar-for-dollar reduction in the investment in securities, but will not reduce the volume of direct lending.

IV. EFFECTS OF AN INCREASE IN RISKINESS

To show an increase in riskiness pertaining to equation (1), a new random variable β^* is defined as by $\beta^* = \beta\theta + \pi$, where Θ is a multiplicative shift parameter and π is an additive shift parameter. An increase in Θ will cause the mean and variance of β to rise. In order to have a mean-preserving increase in risk about β, it is necessary to reduce π simultaneously so that $dE[\beta\theta + \pi] = \bar{\beta}d\theta + d\pi = 0$ or $d\pi / d\theta = -\bar{\beta}$, where E is a linear expectation operator. As a result, the necessary conditions for maximization of expected utility subject to a balance sheet constraint can be rewritten as follows:

(10) $\mathcal{L}_s(\theta) \equiv 0$

(11) $\mathcal{L}_L(\theta) \equiv 0$

and

(12) $\mathcal{L}_\lambda(\theta) \equiv 0$

where $\mathcal{L}_s = \partial\mathcal{L}/\partial s$, $\mathcal{L}_L = \partial\mathcal{L}/\partial L$ and $\mathcal{L}_\lambda = \partial L/\partial\lambda$. Now, differentiating equations (10) through (12) with respect to Θ,

$$(13) \quad \begin{bmatrix} \\ H \\ \\ \end{bmatrix} \begin{bmatrix} \partial S & / & \partial\theta \\ \partial L & / & \partial\theta \\ \partial\lambda & / & \partial\theta \end{bmatrix} = \begin{bmatrix} E \\ F \\ G \end{bmatrix}$$

Elements of the coefficient matrix are same as before. But the constant terms in the right-hand side matrix are different. They are:

$$E = \overline{U}''r_s[r_d-(l-k)](\beta-\overline{\beta})D,$$

$$F = \overline{U}''(r_l+r_l'L)(\beta-\overline{\beta})[r_d-(1-k)]D \text{ and } G = 0.$$

The determinant and its sign remain the same. Using Cramer's rule

$$\partial S/\partial\theta = 1/|H|\overline{U}''(r_s-r_l-r_l'L)(\beta-\overline{\beta})[r_d-(1-k)]D$$

$$\partial L/\partial\theta = 1/|H|\overline{U}''(r_l-r_l'-r_s)(\beta-\overline{\beta})[r_d-(1-k)]D.$$

Again, the substitution of r_S for $r_l + r_l'L$ implies that $\partial S/\partial\theta = \dfrac{\partial L}{\partial\theta} = 0$.

It thus proves that an increase in riskiness does not necessarily alter a large credit union's optimum combination of securities and direct loans. A large credit union is believed to exhibit neutral behavior by offering equal benefits to both borrowers and savers unlike small credit unions, as stated earlier. As a result, despite an increase in riskiness, they might prefer to retain the earlier optimal combination of securities investment and direct loans. This finding seems to be counterintuitive from the viewpoint of standard microeconomic theory which suggests that a rise in risk is likely to alter the preference of a risk-averse economic agent in favor of relatively low-risk assets. Specifically, when deposits become more uncertain, the volume of direct loans is expected to fall. But a credit union may be able to use its capital reserve as a cushion to meet deposit uncertainty. It is also perhaps plausible for a large credit union because of its neutral behavior. Conjecturally, a neutral credit union will seek to preserve its borrowing and lending behavior unless the increase in risk is massive with the potential to become long-lasting.

V. CONCLUSIONS AND REMARKS

Large credit unions in an environment of uncertainty will change the allocations of their assets between securities and direct loans depending on the direction and magnitude of a change in the loan rate relative to the rate of return on investment in securities. The impact of an increase in the reserve ratio and capital reserve buildup will be absorbed completely by the securities investment and have no effects on the volume of direct loans. Additionally, an increase in riskiness does not seem to alter the portfolio behavior of large credit unions in the model presented in the paper.

In conclusion, this model may also be applicable to other depository institutions since the differences between large credit unions and commercial banks are diminishing due to the wave of deregulation that triggered innovations, desegmentation of markets for financial services and an intensely more competitive market environment.

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ON THE RELATIONSHIP BETWEEN BUDGET DEFICITS, INFLATION, AND INTEREST RATES: A MULTIVARIATE ANALYSIS

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I. INTRODUCTION

Whether federal government budget deficits cause high interest rates (short-term and long-term) has become one of the most hotly debated and contentious issues in monetary theory and public finance. Using regression analysis, Hoelscher (1983), Evans (1985), and many others find that the government budget deficit does not cause high interest rates. On the other hand, Makin (1983), Deleuw and Holloway (1985), Hoelscher (1986), Zahid, (1988), Cebula (1988, 1990, 1991, 1993 and 1997), Cebula and Rhodd (1993) and Dua (1993), employing a similar methodology, find that the federal budget deficit has a significant positive impact on interest rates. Although these regression studies have merit, they are weak because they attempt to equate correlation with causation.

In subsequent studies using Granger bivariate causality, Canto and Rapp (1982), McMillin (1986) and Darrat (1990) find that the budget deficit does not Granger-cause interest rates, while Miller and Russek (1991 and 1996) and Vamvoukas (1997) find the opposite. While these recent studies have made significant contributions by employing causality analysis instead of a regression approach, one deficiency is that the lag lengths are chosen arbitrarily. As Lee (1997) argues, the practice of arbitrarily selecting the lag length is a potential source of model misspecification. More specifically, too short a lag length results in estimation bias, while too long a lag length causes a loss of degrees of freedom and estimation efficiency. To remedy this deficiency, Hsiao's version of the Granger-causality approach is adopted in this study. The salient feature of Hsiao's approach (Hsiao, 1981) is that it can estimate the appropriate lag length and at the same time determine the direction of causality.

The purpose of this paper therefore is to extend the current literature by applying the techniques of cointegration and Hsiao's version of Granger causality to reexamine the causality between the federal government budget deficit and interest rates (both short- and long-term rates) in a multivariate framework. Section II presents the methodology and model. Sections III, IV and V report the empirical results, conclusions, and policy implications, respectively.

II. METHODOLOGY AND MODEL

The Granger (1969; 1980) test is quite simple and straightforward. A variable x_t is said to Granger-cause y_t if prediction of the current value of y_t is improved by using past values of x_t. The Granger causality test method is chosen in this study over alternative techniques in light of the favorable Monte Carlo evidence reported by Guilkey and Salemi (1982), and Geweke, Meese and Dent (1983), particularly for small sample applied work. Moreover, the Granger testing procedure can be readily generalized from a bivariate to a multivariate system.

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In the literature, there are four models that characterize the relationship between budget deficits and interest rates. First, in the classical loanable funds/crowding-out effect hypothesis, an increase in the budget deficit will require an increase in government borrowing. The increase in the supply of government bond increases interest rates. Second, in the Keynesian liquidity preference or ISLM frameworks, an expansionary fiscal policy causes the demand for output, money and credit to increase and thus raises interest rates. Third, according to the Ricardian equivalence theorem, an increase in the budget deficit produces an equal increase in private saving, which in turn neutralizes the impact of government borrowing, leaving interest rates unaffected [Barro (1974), Aschauer (1985), Kormendi (1985)]. Finally, monetarists maintain that an increased money supply tends to cause higher inflation and in turn through the Fisher effect higher interest rates. Most recently, Vamvoukas (1997) suggests that the money supply should be included in studying the relationship between budget deficits and interest rates in Greece. Accordingly, interest rates are modeled as a function of the federal government budget deficit, prices, and the money supply. Of course, this list of possible relevant variables is by no means exhaustive and we certainly cannot claim that our expanded model is totally immune from any omission-of-variables biases. Note that all variables are expressed in logs. Thus, the model can be written by specifying a VAR/causality model as follows:

$$(1) \qquad (1-L)\begin{bmatrix}\log y_t \\ \log x_t\end{bmatrix} = \begin{bmatrix}\alpha_t \\ \alpha_2\end{bmatrix} + \sum_{i=1}^{q}(1-L)\begin{bmatrix}\beta_{11i} & \beta_{12i} & \beta_{13i} & \beta_{14i} \\ \beta_{21i} & \beta_{22i} & \beta_{23i} & \beta_{24i}\end{bmatrix}\begin{bmatrix}\log y_{t-i} \\ \log x_{t-i} \\ \log z_{t-i} \\ \log m_{t-i}\end{bmatrix} + \begin{bmatrix}v_{1t} \\ v_{2t}\end{bmatrix},$$

where y = nominal long-term interest rates (I) or nominal short-term interest rates (R), as defined below,

x = nominal federal government budget deficits (BD),

z = inflation measured by the GDP deflator (P),

m = money supply measured by M_1 (M),

L = the lag operator, (1-L) = d is the difference operator such that $(1\text{-}L)y_t = y_t - y_{t-1}$ represents the first difference,

v_{it} = the white noise disturbance terms, and α, β and q represent the constant term, coefficients and maximum lag length, respectively.

This study uses annual data (see Endnote 1) over the period 1946-95 for the U.S. Following earlier studies (Cebula, 1988, 1990, 1991, 1993 and 1997), the nominal long-term interest rate is measured by the yield on Moody's Aaa rated corporate bonds while the nominal short-term interest rate is measured by the yield on three-month Treasury-bills. The data corresponding to the money supply, prices, and the federal budget deficit, as defined above, as well as the long-term and short-term interest rates are obtained from the *Economic Report of the President* (various years).

In testing for causality between budget deficits and interest rates, a multivariate, rather than a bivariate model is employed in this study. Bivariate casualty tests that have been used in previous studies for the U.S., for example, Canto and Rapp (1982), and McMillin (1986), and Miller and Russek (1991) among others, have fallen out of favor in macroeconomics. For instance, Granger (1969), Sims (1980), Lutkepohl (1982) and Serletis (1988) have all argued that Granger casualty tests are severely affected by the biases due to the omission of relevant variables.

In this paper we apply the multivariate approach used by Vamvoukas (1997) in his study of Greece and Cheng (in press) in his study of Japan with some modifications to reexamine the same issue for the United States. Unlike Vamvoukas's approach, the sequence in which each variable enters the estimating equation is determined by the specific gravity criterion (SGC) proposed by Caines, Keng, and Sethi (1981), as discussed below.

III. UNIT ROOT AND COINTEGRATION TESTS

Hsiao's approach to causality requires the series of all variables to be stationary. The results of the Phillips-Perron (PP) tests reveal that while the series of prices is I(0), the series of interest rates (short-term and long-term), the budget deficit, and the money supply are each I(1). The PP tests, however, indicate that the budget deficit, interest rates (short-term and long-term), and the money supply series each become I(0) after first differencing.

Subsequently, we perform the cointegration tests to check whether the standard Granger test is appropriate. Cointegration is the statistical approach which tests for the existence of long-run equilibrium relationships among non-stationary variables which are integrated to the same order. Technically, according to Engle and Granger (1987), if x_t, y_t, z_t, and m_t are each I(1), one would expect that a linear combination of x_t, y_t, z_t, and m_t would be a random walk. Yet, the four variables may have the property that a particular combination of them, $e_t = y_t - bx_t - cz_t - dm_t$, is I(0). If such a property holds, then we say that x_t, y_t, z_t, and m_t are cointegrated. Thus, since price (z_t) is I(0), in testing for the cointegration, we need only to consider the three I(1) variables, x_t, y_t, and m_t.

TABLE 1

RESULTS OF THE PHILLIPS-PERRON (PP) UNIT ROOT TESTS BEFORE AND AFTER DIFFERENCING THE DATA

Variables	Phillips-Perron (PP) Test	
A. Levels	**PP Value**	**Critical Value**
Budget Deficits (BD)	-2.3171	-3.13
Prices (P)	-6.7485*	-3.13
Long-Term Interest Rates (I)	-1.1183	-3.13
Short-Term Interest Rates (R)	-2.1320	-3.13
Money (M_1)	-0.06525	-3.13
B. First-Differences (D)		
DB	-5.3687*	-3.13
DP	-16.6100*	-3.13
DI	-4.8754*	-3.13
DR	-5.3477*	-3.13
DM	-6.6403*	-3.13

Notes: *denote stationary [(see Fuller (1976, p. 373)]

The Engle-Granger two-step cointegration tests for the two three I(1) variables sets ((i, BD, and M) and (r, BD, and M)) are performed first. The Engle-Granger cointegration test suffers from the normality problem—that is, which variable should be used as a dependent variable. To overcome this normality problem, we follow Miller (1981) by choosing the conditioning (left-hand-side) variable that maximizes the adjusted R-squared. The Engle-Granger tests (Table 2; panel A) show that neither the first set of the three I(1) variables (I, BD and M) nor the second set of the three I(1) variables (R, BD and M) is cointegrated.

TABLE 2
COINTEGRATION TEST STATISTICS

A. The Engle-Granger Test

Conditioning Variable	The PP Value	C.V.
CI (I, BD, M)	-2.8712*	-3.84
CI (R, BD, M)	-3.3260*	-3.84

B. The Johansen Test for I, BD and M

	Maximal Eigenvalue Test			Trace Test		
Null H	Alternative H	Eigenvalue	C.V. 99%	Alternative H	LR Ratio	C.V. 99%
r=0	r=1	22.16	26.41	r≥1	24.19	40.20
r≤1	r=2	3.02	19.83	r≥2	3.03	24.99
r≤2	r=3	0.01	12.74	r≥3	0.01	12.74

C. The Johansen Test for R, BD and M

	Maximal Eigenvalue Test			Trace Test		
Null H	Alternative H	Eigenvalue	C.V. 99%	Alternative H	LR Ratio	C.V. 99%
r=0	r=1	26.14	22.47	r≥1	38.68	40.20
r≤1	r=2	12.67	14.92	r≥2	16.21	24.99
r≤2	r=3	1.29	12.74	r≥3	1.29	12.74

Notes: C. V.=critical value; H=hypotheses and PP=Phillips-Perron.

*(Panel A) denotes that the linear combination of the three variables is I(1).

It is worth noting that the Engle-Granger test is bivariate in design. In contrast, the Johansen cointegration test [Johansen and Juselius (1990)] is a more powerful cointegration test, particularly when a multivariate model is used. Moreover, the Johansen cointegration test is robust to various departures from normality in that it allows any of the three variables in the model to be used as the dependent variable while maintaining the same cointegration results. The Johansen cointegration tests (the trace test and the eigenvalue test, as shown in panels B and C of Table 2) also reveal that neither set A of the three variables (I, BD and M) from the long-term interest rate model, nor set B of the three variables (R, BD and M) from the short-term interest rate model are cointegrated.

Thus, evidence obtained from the two cointegration tests indicates that the integrated variables do not have a co-movement tendency over the long run. Therefore, we conclude that interest rates (short-term and long-term), the budget deficit, and the money supply are not cointegrated and therefore the simple Granger causality test (Granger, 1988) is appropriate for both models.

IV. HSIAO'S VERSION OF THE GRANGER MULTIVARIATE CAUSALITY TESTS

While Granger's method is widely used in applied research, its application is restricted to models with identical lag lengths. Such assumptions are not generally valid for many macroeconomic time series. As noted before, an ad hoc lag selection approach has been used in previous studies on the budget deficit and long-term interest rates [e.g., Canto and Rapp (1982) and Miller and Russek (1991 and 1996)]. The problem is that there is no theoretical justification for assuming that two or more related variables must have identical predetermined lag lengths. To remedy this deficiency, Hsiao's approach [Hsiao (1981)] is employed in this study. For details of Hsiao's approach, see Hsiao (1981) and Cheng (1997).

Hsiao's method works well for a bivariate model. For a multivariate equation model, however, it does not ensure that the results of the autoregressive (AR) equation will remain the same when the order in which regressors are introduced is changed. In this study, as indicated, the specific gravity criterion (SGC) proposed by Caines, Keng, and Sethi (1981) is used to determine the sequence in which the regressors are entered at each stage.

Finally, after transforming the original data, we proceed to perform the causality tests. As stated in the appendix, if $FPE(m^*,n^*) < FPE(m^*)$, then budget deficits Granger-cause interest rates and vice versa. In Table 3A, the results indicate that for the long-term interest rate equation, budget deficits are entered into the equation first and since 0.8209764E-02 < 0.8407318E-02, we conclude that budget deficits Granger-cause long-term interest rates. Next, the money supply (M) is entered into the equation and since 0.84737786E-02 > 0.8209764E-02, we infer that the money supply does not Granger-cause long-term interest rates and this variable is dropped from the equation. Finally, prices are entered into the equation and since 0.8544817E-02 > 0.8209764E-02, we conclude that prices do not Granger-cause long-term interest rate and this variable will also be dropped from the equation. In sum, in the long-term interest-rate equation, only budget deficits Granger-cause long-term interest rates.

Conversely, for the budget-deficit equation, as indicated in Table 3A, interest rates are entered into the equation first and we conclude that interest rates Granger-cause the budget deficit, since 853.0833 < 934.8442. Subsequently, the price variable is entered into the equation and we conclude that the price variable does not Granger-cause budget deficit since 901.9052 > 853.0833 and this variable is dropped from the equation. Finally, the money supply is entered into the equation and since 902.7080 > 853.0833, we therefore conclude that the money supply does not Granger-cause the budget deficit and this variable is dropped from the equation. In sum, long-term interest rates are the only variable in the budget-deficit equation to Granger-cause the budget deficit in the United States.

Subsequently, the F-tests (Table 3B) are performed to estimate the magnitude of the coefficients and the results fully corroborate those of the causality tests. The budget deficit is found to cause long-term interest rates at the 10 percent level of significance whereas long-term interest rates are found to cause the budget deficit at the 1 percent level of significance.

Using the same method, we perform tests for the short-term interest rate model. The results, as shown in Table 4A, indicate that for the short interest rate equation, the condition is satisfied, which implies that prices Granger-cause short-term interest rates. When the money supply (M) is entered into the equation, the condition is not met. We thus conclude that the money supply does not Granger-cause interest rates and this variable is dropped from the equation. Finally, budget deficits are entered into the equation and we conclude that the budget deficit also does not Granger-cause short-term interest rates and this variable is dropped from the equation. In sum, in the short-term interest-rate equation, only prices

TABLE 3A

RESULTS OF HSIAO'S VERSION OF GRANGER CAUSALITY TESTS FOR THE LONG-TERM INTEREST RATE MODEL (I)

Controlled Variable	First Manipulated Variable	Second Manipulated Variable	FPE	Causality Inferences
I(j=1)			0.8407318E-02	
I(j=1)	BD(k=3)		0.8209764E-02	BD⇒I
I(j=1)	BD(k=3)	M(n=2)	0.8437786E-02	M≠>I
I(j=1)	BD(k=3)	P(n=1)	0.8544817E-02	P≠>I
BD(j=1)			934.8442	
BD(j=1)	I(k=2)		853.0833	I⇒BD
BD(j=1)	I(k=2)	P(n=1)	901.9052	P≠>BD
BD(j=1)	I(k=2)	M(n=1)	902.7080	M≠>BD

Notes: BD=Budget deficits; I=short-term interest rate, and P= prices.

The figures in parentheses behind the variables are the optimal lag lengths for the independent variable, including the lagged dependent variable.

TABLE 3B

THE T- AND F-TESTS FOR THE LONG-TERM INTEREST RATE MODEL (I)

(1) Interest Rate Equation:

$$I_t = \alpha_0 + \sum_{j=1}^{1} \alpha_i (1-L) I_{t-i} + \sum_{k=1}^{3} \beta_j (1-L) BD_{t-j}$$

(1.32009) (2.52333)
0.26 0.07***

(3) Budget Deficit Equation:

$$BD_t = \alpha_0 + \sum_{j=1}^{1} \alpha_i (1-L) BD_{t-i} + \sum_{k=1}^{2} \beta_j (1-L) I_{t-j} + v_{4t} \ .$$

(0.06518) (4.73204)
0.40 0.01*

The statistics of the joint significance of the coefficients are reported in parentheses and the corresponding p-values (or critical values) are reported below the coefficients.

* denotes significant at the 1% level.

** denotes significant at the 5% level.

*** denotes significant at the 10% level.

Granger-cause short-term interest rates. By the same token, for the budget-deficit equation, as indicated in Table 4A, short-term interest rates are entered into the equation first and we conclude short-term interest rates Granger-cause the budget deficit. Subsequently, prices are entered into the equation and we conclude that prices also do not Granger-cause the budget deficit and this variable is dropped from the equation. Finally, the money supply is added to the equation and we conclude that the money supply does not Granger-cause the budget deficit. In sum, in the budget-deficit equation, short-term interest rates are the only variable that is found to Granger-cause the budget deficit in the United States. The F-test (Table 4B) again is performed and prices are found to cause short-term interest rates at the 1 percent level. However, short-term interest rates are found to cause the budget deficit in the weak sense. According to Darrat and Mukherjee (1995), when a variable satisfies the final prediction error (FPE) criterion and if the parameters of the variable as a group fails to achieve statistical significance in a joint F-test, the variable is said to Granger-cause the left-hand variable only in the weak sense.

Our finding of the bidirectional causality between long-term interest rates and the US federal budget deficits is only partially supportive of those past studies (Hoelscher, 1986; DeLeuw and Holloway, 1985; Cebula, 1988, 1990, 1991 and 1997) that find that long-term interest rates Granger-cause budget deficits without feedback. The finding of this study confirms the loanable funds theory of interest and the Keynesian ISLM theoretical framework that budget deficits affect long-term interest rates. However, our results do not agree with several earlier studies, Hoelscher (1983), Kormendi (1983), Aschauer (1985), Evans (1985), and Darrat (1990) among others, which find that budget deficits do not cause interest rates (long-term/short-term) in the Granger sense. As indicated, the main weakness of these earlier studies cited above is that they attempt to equate correlation with causation.

Some researchers (e.g., Deleuw and Holloway, 1985 and Cebula, 1988 and 1991) have argued that it is essential to distinguish explicitly between the cyclical deficit and the structural deficit. Yet, Hoelscher (1983) maintains that these problems are unlikely to be severe with annual data. In this paper, annual data are used in this study and therefore, this should not be a problem.

The evidence obtained from the budget-deficits/short-term interest rates causality tests are revealing and interesting. It is found that the inflation rate (not the budget deficit) affects short-term interest rates, which basically supports the Fisher effect theory. A close inspection of the data reveals that the interest rates on three-month Treasury bills have usually moved along with the inflation rates. Further, the findings that short-term interest rates Granger-cause the budget deficit are consistent with those of previous studies (e.g., Zahid, 1988; Cebula, 1988 and 1991, and Miller and Russek, 1991 and 1996).

Interestingly, Zahid finds that budget deficits negatively cause nominal short-term interest rates. Our findings, however, do not support the Ricardian Equivalence Theorem. As Evans (1985) argues, most economists are hesitant to accept Barro's argument in part because they have misgivings about people's ability to foresee the future tax rate suggested by a large current budget deficit.

Surprisingly, we find that the money supply (M_1) does not Granger-cause interest rates (long-term or short-term). This finding is completely at odds with the Keynesian IS-LM framework and Friedman's modern quantity theory of money. We choose M_1 rather than M_2 or M_3, because data for the latter were not available prior to 1959. If either M_2 or M_3 were employed, the results could be different.

V. CONCLUSIONS

This paper seeks to advance and extend the current literature on the relationship between budget deficits and interest rates.

The main contribution of this study is methodological. Due to the practice of arbitrarily selecting the lag length, past studies have suffered from potential model misspecifications. That is, too short a lag or too long a lag may lead to results that are either biased or inefficient. Hsiao's approach, which was employed in this study, can estimate the appropriate lag length and at the same time determine the direction of causality.

TABLE 4A

RESULTS OF HSIAO'S VERSION OF THE GRANGER CAUSALITY TESTS FOR THE SHORT-TERM INTEREST RATE MODEL (R)

Controlled Variable	First Manipulated Variable	Second Manipulated Variable	FPE	Causality Inferences
R(j=2)			0.6991693E-01	
R(j=2)	P(k=5)		0.3612274E-01	P⇒R
R(j=2)	P(k=5)	M(n=1)	0.4023634E-01	M≠>R
R(j=2)	P(k=5)	BD(n=1)	0.4042208E-01	BD≠>R
BD(j=1)			934.8442	
BD(j=1)	R(k=1)		880.2010	R⇒BD
BD(j=1)	R(k=1)	P(n=1)	921.2053	P≠>BD
BD(j=1)	R(k=1)	M(n=3)	921.8268	M≠>BD

Notes: BD=Budget deficits; R=long-term interest rate, and P= prices.

The figures in parentheses behind the variables are the optimal lag lengths.

TABLE 4B

THE F-TESTS FOR THE LONG-TERM INTEREST RATE MODEL (R)

(1) Short-term Interest Rate Equation:

$$R_t = \alpha_0 + \sum_{j=1}^{1}\alpha_i(1-L)R_{t-i} + \sum_{k=1}^{5}\beta_j(1-L)P_{t-j}$$

(1.38308) (11.09676)
0.25 0.00*

(3) Budget Deficit Equation:

$$BD_t = \alpha_0 + \sum_{j=1}^{1}\alpha_i(1-L)BD_{t-1} + \sum_{k=1}^{1}\beta_j(1-L)R_{t-j} + v_{4t} \ .$$

(0.83388) (0.42515)
0.37 0.52

The statistics of the joint significance of the coefficients are reported in parentheses and the corresponding p-values are reported below the coefficients.

* denotes significant at the 1% level.

** denotes significant at the 5% level.

*** denotes significant at the 10% level.

The results of this study reveal a bidirectional causal linkage between budget deficits and nominal long-term interest rates. In other words, nominal long-term interest rates and the budget deficit are linked by a feedback causal mechanism. This confirms the classical loanable funds theory of interest and the crowding-out hypothesis that a large budget deficit tends to cause high long-term interest rates. The policy implication is that if the U.S. government wants to keep long term interest rates at a low level in order to maintain a steady economic growth rate, it should reduce its budget deficit to the minimum or eliminate it altogether. Furthermore, decreased interest rates tend to lower the government interest payments, and further reduce the budget deficit. Thus, the budget deficit and long-term interest rates affect and influence each other.

The second main finding of this study is that the inflation rate Granger-causes short-term interest rates without feedback, whereas short-run interest rates Granger-cause budget deficits but not vice versa. This supports the Fisher effect theory which states that, if the inflation rate soars, (short-term) interest rates rise. It implies that if the U.S. government can curb inflation, it may be able to reduce short-term interest rates. Thus, by lowering the inflation rate, the government may not only reduce short-term interest rates, but may also reduce its budget deficit.

APPENDIX

The procedure adopted to implement Hsiao's version of the Granger-causality tests in this paper is presented as follows:

(i). Equation (1) in the text can be broken down into two equations: an interest rate equation and a budget deficit equation. The interest rate equation in turn can be broken into four causality equations: univariate equation, bivariate equation, trivariate equation and four-variate equation (all variables are expressed in logs):

$$\text{(A1)} \quad (1-L)y_t = \alpha_0 + \sum_{j=1}^{M}\alpha_j\,(1-L)y_{t-j} + v_{1t}\,,$$

$$\text{(A2)} \quad (1-L)y_t = \alpha_0 + \sum_{j=1}^{M}\alpha_j\,(1-L)y_{t-j} + \sum_{k=1}^{N}\beta_k(1-L)x_{t-k} + v_{1t}\,,$$

$$\text{(A3)} \quad (1-L)y_t = \alpha_0 + \sum_{j=1}^{M}\alpha_j\,(1-L)y_{t-j} + \sum_{k=1}^{N}\beta_k(1-L)x_{t-k} + \sum_{n=1}^{P}\gamma_n(1-L)z_{t-n} + v_{3t}\,,$$

$$\text{(A4)} \quad (1-L)y_t = \alpha_0 + \sum_{j=1}^{M}\alpha_j\,(1-L)y_{t-j} + \sum_{k=1}^{N}\beta_k(1-L)x_{t-k} + \sum_{n=1}^{P}\gamma_n(1-L)z_{t-n}$$
$$+ \sum_{s=1}^{Q}\eta_s\,(1-L)m_{t-s} + v_{4t}\,,$$

Using equation (1) above for illustration, in step one we treat the dependent variable, y_t, as a one-dimensional autoregressive (AR) process initially, and compute the sum of squared errors (SSE) and the corresponding FPE as defined by Akaike (1969) using equations (A1) and (A5) with the maximum order of lags varying from 1 to M.

$$\text{(A5)} \quad FPE(m) = (T + m + 1)/(T - m - 1) \bullet (SSE/T),$$

where T = total number of observations,
m = the order of lags varying from 1 to M, and
SSE = sum of squared errors.

We then select the order which yields the smallest FPE, m*. In step two, we focus on equation (A2). We then treat y_t as a controlled variable and compute SSE and FPE using equations (A2) and (A6) with the order of lags set at m*, and x_t as a manipulated variable by varying the order of lags of x_t from 1 to N and determine the order which yields the smallest FPE, n*. The corresponding two-dimensional FPE is

$$\text{(A6)} \quad FPE(m^*,n) = (T + m^* + n + 1)/(T - m^* - n - 1) \bullet (SSE(m^*,n))/T,$$

where n = the order of lags on x(t) varying from 1 to N, and
m* = the optimum number of lags computed from (A5).

If FPE(m^*,n^*) is less than FPE(m^*), we then conclude that the budget deficit (x_t) Granger-causes interest rates (y_t). Subsequently, by using the same procedure, FPE(m^*,n^*) and FPE(m^*, n^*, p^*) can be obtained and compared with each other. By repeating the same procedure for the budget deficit equations, causality from interest rates, the money supply and the inflation rate to the budget deficit can also be estimated.

ENDNOTE

The author would be willing to make copies of raw data available to other researchers upon request.

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WHY A CONSUMPTION TAX IS BETTER FOR CHINA

Joseph Cheng*

CONSUMPTION TAX VS INCOME TAX

In the past decade, tax revenue in China has grown at a much slower rate than GDP, causing a tax revenue as a percent of GDP to fall from 34 percent in 1978 to 19 percent in 1991. Economic reforms tend to undermine revenue collection for two reasons: tax collections from the new private firms lag behind because of administrative difficulties; and state enterprise profits falls as increased competition erodes profit. The following table reveals the various sources of tax revenue, underscoring the problem of state owned enterprises.

TABLE 1

SOURCES OF TAX REVENUE FOR 1995 (IN 100 MILLION YUAN)

Industrial and Commercial	4589.68
Tariff	291.83
Agriculture	278.09
Subsidies to Loss Enterprises	(327.77)
Other	1206.21
Total Revenue	6038.04

As seen in Table 1, the net amount of revenue from state owned enterprise is actually negative. With the rapid decline of state enterprises in China, which have provided profit revenue to the state treasury in the past, the Chinese government is now relying increasingly on income taxes on the profits of private firms and individual incomes. The Chinese individual income tax system is similar to the American System. In the U.S., however, there has been talk of overhauling the income tax system or even eliminating it altogether. Although the latter is highly unlikely, China may learn from our debate about the problems inherent in the income tax system.

Economists are well aware that both the consumption and income taxes distort individual work and leisure choices. In addition, income taxes, by creating a wedge between the private rate of return and the after-tax rate of return on investment, also distort the tradeoff between present consumption and future consumption. Other than a lump sum tax, all taxes discourage whatever activities they tax. An income tax covers all earnings, but a consumption tax does not cover the portion that is saved. Thus, savings is encouraged. In general, economists believe that an income tax system generates more inefficiency than does a consumption tax, because an income tax system imposes a greater burden on capital than a consumption tax system. This is especially important in light of the unemployment problem faced by

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China, since capital investment is needed to provide new jobs. In raising revenue, government policies should minimize the disincentives created by the tax as much as possible, especially in situations where there is severe unemployment.

Fullerton, Shoven, and Whalley (1980) estimated that the annual efficiency gain from switching from an income tax to a consumption tax might be about 2 or 3 percent of national income for the United States. While the efficiency gain for China might be more or less, the magnitude is undoubtedly significant.

From a practical perspective, the administration of the individual income tax has thus far been unsatisfactory in China. This may be due to the dominance of cash transactions in the economy and the limited non-cash flows in the financial system. The incentive for businesses to comply is equally lacking. A countervailing force against the underreporting of business income is that it depresses corporate stock prices, which are linked to earnings. But since most businesses in China today are not incorporated, this neutralizing factor may not be significant. It is estimated that 80 percent of individuals and private enterprises and 40 percent of stated owned enterprises may have underpaid their taxes in recent years (Li 1991). Because of severe compliance problems, tax collection is becoming more aggressive. In Pingtan, for example, tax collectors scan premises and estimate revenue and costs; business revenues are taxed at a flat rate of 8 percent and profit is taxed on a sliding scale. Still, tax revenue as a percent of GDP, which was 10.9 percent in 1996, is still far below that of the U.S. (19 percent) (Johnson 1997). Further, the income tax system in China is plagued by the arbitrary nature of the tax contract system whereby tax assessors have the authority to vary profit tax rates based on special circumstances. Consequently, profit taxes are in effect negotiable.

An appropriate tax system, especially for a populous nation such as China, should have the following features:

1. It should be fair and equitable.
2. It should be simple to administer and have high compliance rate.
3. It should minimize the adverse impact on employment.
4. It should promote conservation of vital natural resources.
5. It should strengthen government's ability to protect the environment.

The consumption tax will be discussed with respect to these features. Progressivity, which many economists associate with equity, is generally achieved by taxing the consumption of individuals with higher consumption levels at higher rates. Generally, a progressive consumption tax is assumed to be administered as an income tax with a deduction for savings (Pechman 1985). In this case, an individual would still have to file a consumption tax form, which is basically the same as an income tax form plus a savings deduction form. Because the income tax form remains intact under such a consumption tax system, the consumption tax is potentially plagued with the same compliance problem associated with the income tax system.

THE GRADUATED EXCISE TAX

Given this problem, I propose that a consumption tax be administered on the expenditure side rather than on the revenue side; this could be accomplished in China by modifying the product tax, which is a form of excise tax already in place. Excise taxes are often thought to be regressive in nature; but this need not be the case. Progressivity can be achieved by imposing higher tax rates on products and services with high income elasticities. For future reference, such tax system will be called the Graduated Excise Tax, or simply GET.

Specifically, goods and service can be classified into five rate categories, with luxury goods placed in the highest rate categories, and necessities in the lowest category. For example, foods, health service, and clothing should be placed at the lowest category or even be totally tax exempt. Every final product and service would be placed in its own appropriate rate category, depending on the degree of income elasticity. The classification could also take other factors such as the environment and resource conservation into account. For example, items generating pollution either in production or consumption could be placed in the higher rate categories.

As can be seen, the GET not only raises needed revenue, but also can be used to accomplish socially desirable goals. Given China's rapid industrial expansion, environmental degradation is becoming a serious problem. However, the development of environmental protection polices in China is still in its infancy. It is widely accepted that setting tax rates on products commensurate with the levels of pollution they cause is more efficient than outright bans, arbitrary limits, or having no regulation at all. With China's huge population, the conservation of certain resources such as oil, minerals, and lumber should be encouraged. Thus, higher taxes on these commodities might be warranted. The Resource Tax, introduced in 1984, is a profit tax on coal, oil, and gas enterprises. However, the levels of the tax rate are based on profit margin or profit-sales ratio, which has little relationship with the degree of scarcity of the resource in question. Under this system, earnings below 12 percent of sales are exempt from the resource tax. Earnings between 12 and 20 percent of sales are taxed at 50 percent. Earnings beyond the 20 percent sales level are taxed at a still higher rate. Since enterprises with higher profit margins tend to have lower operating costs, the effect of this tax is to favor less efficient producers, thus causing efficient enterprises to underproduce and inefficient ones to overproduce. The GET here does not suffer from this problem since it is resource-specific rather than enterprise-specific. Further, the GET differs from the Product Tax and a standard consumption tax in that it is imposed at the retail level rather than the manufacturing level or personal level. The advantage of imposing the tax on domestic retailers rather than domestic manufacturers is the automatic exemption of taxes for goods to be sold abroad, thus keeping exports competitive in the world market. In comparison to the VAT (Value Added Tax), a commodity to be processed by five manufactures before becoming a final product would be taxed five separate times under the VAT. But the GET is levied only once at the retail level, thus it is more administratively efficient. Further, in cases where products could create negative external effects when used in a particular manner, the tax should be levied at the final stage. For example, alcoholic beverages might create negative externalities, whereas alcohol used for medicinal purpose would not. Since the GET is assessed at the final stage of the product, usage could be distinguished and tax rates would be applied accordingly, but not so for the VAT. In sum, the GET can help achieve socially desirable goals such as conservation of precious resources and achieve progressivity with less distortions to economic decisions, as well as require lower administration cost than the current tax system in China. The compliance rate should be higher and accounting costs lower than under either the standard consumption tax or the income tax because forms for individuals under GET would virtually be eliminated.

There might also be an advantage to the GET that goes beyond economics. Taxes imposed on income or work might be contrary to the rather strong work and savings ethics inherent in the traditional Chinese culture. Psychologically, people may resent taxes imposed directly on the money they work hard for more than taxes levied on products and services. Since the GET is not levied directly on either work or savings, it would be better tailored to the customs and values of Chinese society than the income tax system.

CONCLUSION

I propose that the GET replace the current indirect tax and the individual personal income tax. It is more economically and administratively efficient to raise the bulk of the tax revenue from one or two broad systems than from dozens of different types of taxes which overlap in terms of functions and in incidence, as China is currently doing. With China's great industrial capacity and potential, the magnitude of the economic gains that could be realized by the adoption of an efficient tax system is tremendous.

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EMPLOYMENT AND THE OUTCOME OF ELECTIONS FOR GOVERNOR IN NEW YORK

Edward Renshaw*

Since the consumer price index was extended back to 1913 one can use multiple regressions and two addition economic variables to explain all of the outcomes of presidential elections in the U.S. since Wilson promised to keep us out of World War I. During this period there has always been a political turnover if the civilian unemployment rate increased in the last two years of a presidential term and the CPI inflation rate was over 4.5 percent or if the four year increase in civilian employment was less than 3.5 percent (He and others 1998).

Data pertaining to employment are also of value in helping to explain the outcome of elections for governor in New York. In the post World War II period no political party has been able to keep a New York governor in office if nonagricultural employment in New York in the election year was down one percent or more from its high in the preceding four years. Harriman in 1954, Rockefeller in 1958, Carey in 1975 and Pataki in 1994 may have all been the beneficiaries of major declines in employment opportunities in the state of New York. See Table 1.

The story is different for minor declines in employment. There have been five election years since 1939 (1950, 1970, 1978, 1982 and 1990) when nonagricultural employment in New York was down from .36 to .46 percent from its preceding four year high and in each of these cases the incumbent governor was able to get reelected.

There is not much doubt that states can adopt measures that will sometimes prevent or ameliorate minor declines in employment. It is doubtful, however whether the state of New York, with its requirement for a balanced budget, can do much to prevent or ameliorate major declines in employment opportunities which are national in scope.

The four political turnover cases involving employment declines of one percent or more in New York in an election year are associated with the recessionary trough years for the nation as a whole (1954 and 1958), a prolonged recession following the food and energy price shocks of 1973-74 and a very slow recovery from the employment recession of 1990-92 which was national in scope.

It should be noted that the behavior of employment in New York has varied a lot in relation to what has happened over time at the national level. It is not national recessions per se that have led to political turnovers in New York but their severity and "unfortunate timing" in relation to local elections. Since Dewey was first elected governor in the middle of World War II there have been ten national recessions but only four gubernatorial turnovers in New York.

The good news with regard to recessions that are national in scope is that they are occurring less frequently. At some point in the future, however, there are likely to be some more elections that are decided on the basis of non-economic issues.

The rosy outlook for the Pataki administration (as of July 1998) is related in part to an exuberant stock market that is not likely to last forever. Hope does seem to spring eternal before a presidential election,

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however. Since 1941 all of the negative financial returns for the S&P composite stock price index have occurred either in post election years or in the following year.

Readers and/or politicians who are not too happy with the Pataki administration, however, can perhaps take some comfort in the fact that models with a long history of explaining election outcomes sometimes break down and do not explain the future as well as the past. The Fair model (1978) and two models publicized by Trahan and Renshaw (1990) to represent a turnover contour associated with the theory of a political business cycle and the value of a consensus forecast also failed to explain the turnover that occurred during the presidential election year of 1992.

TABLE 1

EMPLOYEES IN NONAGRICULTURAL ESTABLISHMENTS IN NEW YORK AND THE OUTCOME OF ELECTIONS FOR GOVERNOR, 1939-94.

Year	Governor Elected	Party Affiliation	Total Employment (Thousands)	Percentage Change[a]
1938	Lehman	(D)	N.A.	N.A.
1941			4,735.3	
1942	Dewey	(R)*	4,997.8	5.54
1943			5,226.3	
1946	Dewey	(R)	5,324.8	1.88
1948P			5,596.1	
1950	Dewey	(R)	5,576.0	-.36
1953P			5,935.6	
1954T	Harriman	(D)*	5,828.3	-1.81**
1957P			6,179.0	
1958T	Rockefeller	(R)*	6,027.2	-2.46**
1960P		6,181.9		
1962	Rockefeller	(R)	6,261.3	1.28
1965			6,518.7	
1966	Rockefeller	(R)	6,709.5	2.93
1969P		7,182.0		
1970T	Rockefeller	(R)	7,156.4	-.36
1974D	Carey	(D)*	7,077.1	-1.11**
1978	Carey	(D)	7,044.5	-.46
1981P		7,287.3		
1982T	Cuomo	(D)	7,254.6	-.45
1985		7,751.3		
1986	Cuomo	(D)	7,907.9	2.02
1989			8,246.8	
1990P	Cuomo	(D)	8,212.4	-.42
1994	Pataki	(R)*	7,818.7	-4.79**
1997			8,012.8	
1998	?	?	?	?

[a]The percentage change is based on the highest employment level in the preceding four years.

D associated with the year 1974 identifies a year of intervening decline in economic activity in the U.S. without a peak or trough in economic activity as defined by NBER.

P identifies a year containing a recessionary peak in economic activity.

T identifies a year containing a recessionary trough in economic activity.

*identifies turnover years when the incumbent political party lost the election for governor in New York.

**identifies declines in employment from the preceding four year high in an election year amounting to one percent or more. All of these cases are associated with election turnovers.

Source of employment statistics: New York State Department of Labor, *Employment Review.*

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SEARCH AND EXPERIENCE GOODS: A NOTE ON INVENTORY TO SALES RATIOS

Arthur S. Leahy*

In "Information and Consumer Behavior," Nelson (1970) distinguishes between search goods, whose qualities can be determined by the consumer prior to purchase, and experience goods, whose qualities cannot be determined before purchase. This paper focuses on one aspect of search and experience goods that was examined by Nelson, differences in inventory to sales ratios between these two types of goods.

Nelson's study was based on the merchandise inventories held by stores for sale to retail customers. He expects a store to stock more brands for search goods than for experience goods for two reasons. First, Nelson expects the market to support more brands for search goods than for experience goods (Nelson, 1970, pp. 312-18). Second, he expects stores to find it more difficult to carry a smaller number of brands of search goods than of experience goods, because a consumer would want to examine several search goods prior to purchasing these goods. Nelson's results show that the ratio of inventory to sales is greater for search goods than for experience goods. To see if Nelson's results would hold true at a different level of the distribution chain, this paper examines the difference between search and experience goods using data on inventories held by wholesale distributors for sale to retailers.

Nelson's distinction between search and experience goods is based on consumer responses to these different product types. As you move back along the distribution channel away from the final consumer, it is possible that the effect of the method by which the consumer obtains product information may diminish. Thus, what matters to consumers may not matter to retailers. Furthermore, inventory theory predicts that the inventory to sales ratio for a brand should be smaller as the sales for the brand increase. Because a wholesale distributor would tend to have greater sales than a retailer and carry a narrower range of merchandise, they would tend to have lower inventory to sales ratios.[1] If this difference in inventory to sales ratios has a differential effect on search versus experience goods, there may be a difference in the results from those of Nelson from this source as well.

INVENTORY TO SALES RATIOS

The difference in the inventory to sales ratios between search and experience goods is related to the difference in the number of brands between the two categories. Nelson indicates that there is an inverse relationship between the number of brands a store is economically required to carry and the sales for a brand, holding total store sales constant. As noted above, inventory theory would predict that the inventory to sales ratio for a brand would be smaller when the sales for a brand is greater. Therefore, the ratio of inventory to sales of a store should increase as the number of brands it carries increases.

As mentioned previously, Nelson expects a store to stock more brands for search goods than for experience goods. In order to test this hypothesis, Nelson used data from the 1929 *Census of Retailing.*

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* Internal Revenue Service and Central Michigan University.

Nelson's results show that the ratio of inventory to sales is indeed greater for search goods than for experience goods. This relationship exists for both durable and nondurable goods.[2]

To see if Nelson's results would hold true for different data, a different time period, and for a different level of the distribution chain, data from the 1994 Compustat database were used to examine the ratio of inventory to sales for all durable and nondurable wholesale distributors having inventory in that year. Although a wholesale distributor would tend to have greater sales for the more limited number of brands they carry than a retailer, and therefore lower inventory to sales ratios,[3] the difference in the size of the market between the retailer and the wholesale distributor should translate upstream along the distribution channel such that it would still be the case that there would be more brands for search goods. Therefore, there should be no difference in the distribution of inventory to sales ratios between search and experience goods due to this difference in the level of the market. For example, using figures from Table 1, if the inventory to sales ratio for a distributor of search goods were .18 versus .11 for an experience goods distributor, it would be expected that the ratio of inventory to sales for a retailer of these search goods would be approximately 1.6 times as large as this ratio for a retailer of these experience goods.[4]

TABLE 1
MEAN INVENTORY TO SALES RATIOS
DURABLE AND NONDURABLE GOODS DISTRIBUTORS, 1994

CLASSIFICATION 1			
	Experience Goods	**Search Goods**	**Differences**
Arithmetic Mean	0.11	0.18	0.07
Geometric Mean	-2.22	-1.78	0.44
$t^* = 1.46$			
CLASSIFICATION 2			
	Experience Goods	**Search Goods**	**Differences**
Arithmetic Mean	0.11	0.18	0.07
Geometric Mean	-2.21	-1.78	0.43
$t^* = 1.43$			

*Test result of the hypothesis that the difference in the geometric means is equal to zero.

Distributors were classified into search and experience categories based on their Standard Industrial Classification (SIC) code following Nelson's (1970, p. 325) classification of experience and search goods. For example, distributors classified by Compustat as being within SIC code 5045, Computers and Computer Peripheral Equipment and Software, were classified as selling experience goods. On the other hand, distributors classified by Compustat as being within SIC code 5072, Hardware, were classified as selling search goods.[5]

Two different classification systems of experience and search goods are used. In classification 1, the jewelry, watches, and silverware category is included with the experience goods. In classification 2, this category is included with the search goods. The problem with this particular category has been discussed by Nelson (1970). Watches should be classified as an experience good while jewelry and silverware should be classified as search goods.

A comparison of the geometric mean ratios of inventory to sales for the two different categories of goods was then made.[6] A test of the hypothesis that the difference in the means is equal to zero against the alternative hypothesis that the mean of the search good category was greater than the mean of the experience good category gave the results shown in Table 1.

The results show that the ratio of inventory to sales is greater for search goods than for experience goods, regardless of which classification system is used. However, the difference in these ratios is not statistically significant. Errors in classification may be responsible for this result due to the use of the rather broad four-digit SIC codes to categorize distributors. As Nelson (1970) points out, random errors in classification would tend to bias the results toward zero. Therefore, any bias of this type would work against obtaining a significant relationship. Nevertheless, the results suggest that the distinction between search goods and experience goods may not hold as you move down along the distribution chain away from the end user.

CONCLUSION

In conclusion, the ratio of inventory to sales at the wholesale level has been found to be greater for search goods than for experience goods, although the difference in these ratios was not statistically significant. The results suggest that as you move back along the distribution channel away from the final consumer, the effect of the method by which the consumer obtains product information diminishes. Retailers don't have the consumer's interest in examining the product. This effect would only apply, however, to Nelson's second explanation for the tendency of stores to carry more brands for search goods than for experience goods, i.e., stores would find it more difficult to carry a smaller number of brands of search goods than of experience goods because a consumer would want to examine several search goods before purchasing them. In this case, the constraint of retail store size, which limits the number of brands of both types of goods they can carry, would not be as restrictive for a wholesale distributor. Presumably, wholesale distribution space would be significantly cheaper than retail space.

Some additional factors operate. As mentioned previously, inventory theory predicts that the inventory to sales ratio for a brand should be smaller as the sales for the brand increase. Because a wholesale distributor would tend to have greater sales for the more limited number of brands they carry than a retailer, they would tend to have lower inventory to sales ratios. If the effect of the difference in the distribution channel has a differential effect on search versus experience goods, there may be a bias introduced from this source as well. There is no reason to expect this to be the case, however. Furthermore, inventory costs increase at an increasing rate as customer service levels increase (Kotler, 1988, p. 580; Kotler and Armstrong, 1989, p. 371). In addition, inventory carrying costs increase in direct proportion to the average amount of inventory carried (Brigham, 1986, p. 660; Weston and Brigham, 1987, p. 423). It is estimated that it costs twenty percent or more of the value of inventories to carry them for one year (Moore, p. 567). If the combined effect of these latter two factors varies with the distribution channel, and if this results in a differential impact on search goods versus experience goods, this may lead to a difference in the results as well from those obtained by Nelson.

On the other hand, the other reason given by Nelson for a store to stock more brands of search goods than of experience goods, i.e., the market would support more brands for search goods, should not be affected by the difference between the retail and wholesale level data. Regardless of the difference in the size of the market between the retailer and the wholesale distributor, the distinction between search goods and experience goods should translate upstream along the distribution channel such that it would still be the case that there would be more brands for search goods.

Overall, the combination of these effects would make it harder to find a significant difference in the inventory to sales ratios between search and experience goods at the wholesale level, which may explain why the results obtained here are different from those found by Nelson.

ENDNOTES

1. In 1992, the average retail firm had sales of $881,103 versus average sales of $8,369,509 for wholesale firms. The average inventory to sales ratio was 1.53 during 1992-1995 for retail firms versus an average inventory to sales ratio of .07 for wholesale firms in 1992 (U.S. Bureau of the Census, 1996).
2. When jewelry stores were included with the experience goods category, the difference in the ratios was in the right direction but was not significant.
3. See note 1.
4. .18/.11 = 1.6.
5. The sample includes 150 companies classified as selling experience goods, 16 companies classified as selling search goods, and 1 company classified as selling ambiguous goods. It is representative of the entire range of products sold by these distributors.
6. The use of the geometric mean reduces the likelihood of bias resulting from a few large observations in the data. It is particularly well suited for measurement of the central tendency of ratios (Brumbaugh and Kellogg, 1946, p. 492). Throughout this paper, the geometric mean is expressed in logarithmic form.

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1. Dal Didia
2. Richard Dietz
3. Joseph Eisenhauer
4. Barbara Howard
5. Elia Kacapyr
6. William Kolberg
7. Laurence Malone
8. Michael Schinski
9. Wade Thomas
10. Robert Tokle

NEW YORK STATE ECONOMICS ASSOCIATION (NYSEA)

50th ANNUAL CONVENTION

FINAL PROGRAM

State University of New York • College at Oneonta
Morris Conference Center – Oneonta, New York

September 26-27, 1997

Friday, September 26

8:00 – 10:00 PM	**NYSEA Convention Opening Reception** (LeCafe, Morris Conference Center)
	Wine and Cheese Reception
	Introduction: Wade L. Thomas, President, NYSEA
	Welcome: Alan Donovan, President SUNY Oneonta
Speaker:	Marc Lieberman, Vassar College *"Thinking Like a Non-Economist: The First Day of Econ 100"*

Saturday, September 27

8:00 – 10:00 AM	**Convention Registration & Continental Breakfast** (Craven Lounge, Morris Conference Center) Pick up final program, receipt/register, location directions, name tags
8:00 – 2:00 PM	**Textbook Display/Exhibits** (Craven Lounge, Morris Conference Center)
8:30 AM	**Sessions Begin**
12:00 – 1:30 PM	**Luncheon** (Otsego Grill)
	Speaker: David Levine, Delegate, Quebec Government House *"The Quebec/New York Partnership at the Crossroads of the 21st Century"*
	Afternoon Refreshments (Craven Lounge, Morris Conference Center)
3:30 – 4:30 PM	**NYSEA Business Meeting** (Room 103)

SESSION

8:30 – 10:00 AM **What's New in Education** (Room 103)

Chair: Michael Gordon, SUNY Canton

Peter Bell, SUNY Purchase
"Integrating Race and Gender Into the Teaching of Economics"

Larry Malone, Hartwick College
"Testing Reasonableness in Rethinking the Economics Major"

Jeannette Mitchell, Rochester Institute of Technology
"Economic Instruction and Political Socialization: A Gender Analysis of Six Upstate New York High School Economics Classes"

Michael Gordon, SUNY Canton
"Virtual Economics—An Interactive Center for Economic Education"

SESSION

8:30 –10:00 AM **Macroeconomics** (Room 105)

Chair: David Ring, SUNY Oneonta

Rod D. Raehsler and Lynn A. Smith,
Clarion University of Pennsylvania
"Structural Changes in the Consumption Function and Recent Patterns in Economic Growth"

Susanne M. Polley, SUNY Cortland
"The Role of Fiscal Disparity in Policy Maker Forecasting"

Thomas Kopp, Siena College
"European Economic Union: Are the Markets Listening?"

Discussant:
Elia Kacapyr, Ithaca College

SESSION

8:30 –10:00 AM **Transportation and Economic Development** (Room 104)

Chair: Rick Fenner, Utica College

William P. O'Dea, SUNY Oneonta
"An Economic Evaluation of Automated Highways"

Donald F. Vitaliano and Kathleen Helfrich,
Rensselaer Polytechnic Institute
"Size Economies and Cost Efficiency in the Provision of Local Road Services in New York"

Stephen A. Kolenda, Hartwick College
"Economic Development of the Golden Quadrangle"

Discussant:
Joseph G. Eisenhauer, Canisius College

SESSION

10:15 – 11:45 AM **Making Economics Come Alive in the Elementary School Classroom** (Room 105)

Chair: John Clow, SUNY Oneonta

Panelists:
Area teachers who have completed the teacher education course and are implementing various economic education activities.

SESSION

10:15 – 11:45 AM **Regional Economics** (Room 103)

Chair: Stephen A. Kolenda, Hartwick College

Elia Kacapyr, Ithaca College
"Data Sources for County Economies"

Kent Klitgaard, Wells College
"Calculating Income Distribution in the Cayuga Basin"

Robert Jones, Skidmore College
"Sectoral Employment Cycles in Upstate New York"

Discussant:
David Ring, SUNY Oneonta

SESSION

10:15 – 11:45 AM **Microeconomics** (Room 104)

Chair: William P. O'Dea, SUNY Oneonta

Joseph G. Eisenhauer, Canisius College
"Relative Magnitudes of Risk Aversion and Prudence"

Rick Fenner, Utica College
"Do Researchers Make Better Teachers?"

James F. Booker, Alfred University
"Using the World Wide Web in the Introductory Microeconomics Course"

Discussants:
Barbara Howard, SUNY Geneseo
David F. Vitaliano, Rensselaer Polytechnic Institute

SESSION

1:45 – 3:15 PM **Special Topics and Student Papers** (Room 104)

Chair: Susanne M. Polley, SUNY Cortland

Alex Stricker, Syracuse University
"Comparing School Aid Programs Using Household Valuations of Education Outcomes"

Ernest Enke and Frank Duserick, Alfred University
"Reinventing the Future: A Start on Reclaiming the American Dream"

Cynthia Compton, Frank Duserick, Ernest Enke, Kellie L. Hawks, Rebecca A. Hellinger, and Alison Zielinski, Alfred University
"HMOs Versus PSOs as Managed Care Providers"

Joseph Ford, Iona College
"Was It Just Another Sermon?: The Economic Letter of the Bishops A Decade Later"

Discussant:
James F. Booker, Alfred University

SESSION

1:45 – 3:15 PM **Economic Systems** (Room 103)

Chair: George J. Neimanis, Niagara University

Aberraman Robana, Alfred University
"Privatization in Tunisia"

Andrew Pienkos, Cornell University
"Sense and Nonsense in the Privatization Debate: The Polish Case"

Bogdan Mieczkowski, Ithaca College
"An Experience with Media"

Discussant:
Alfred Lubell, SUNY Oneonta

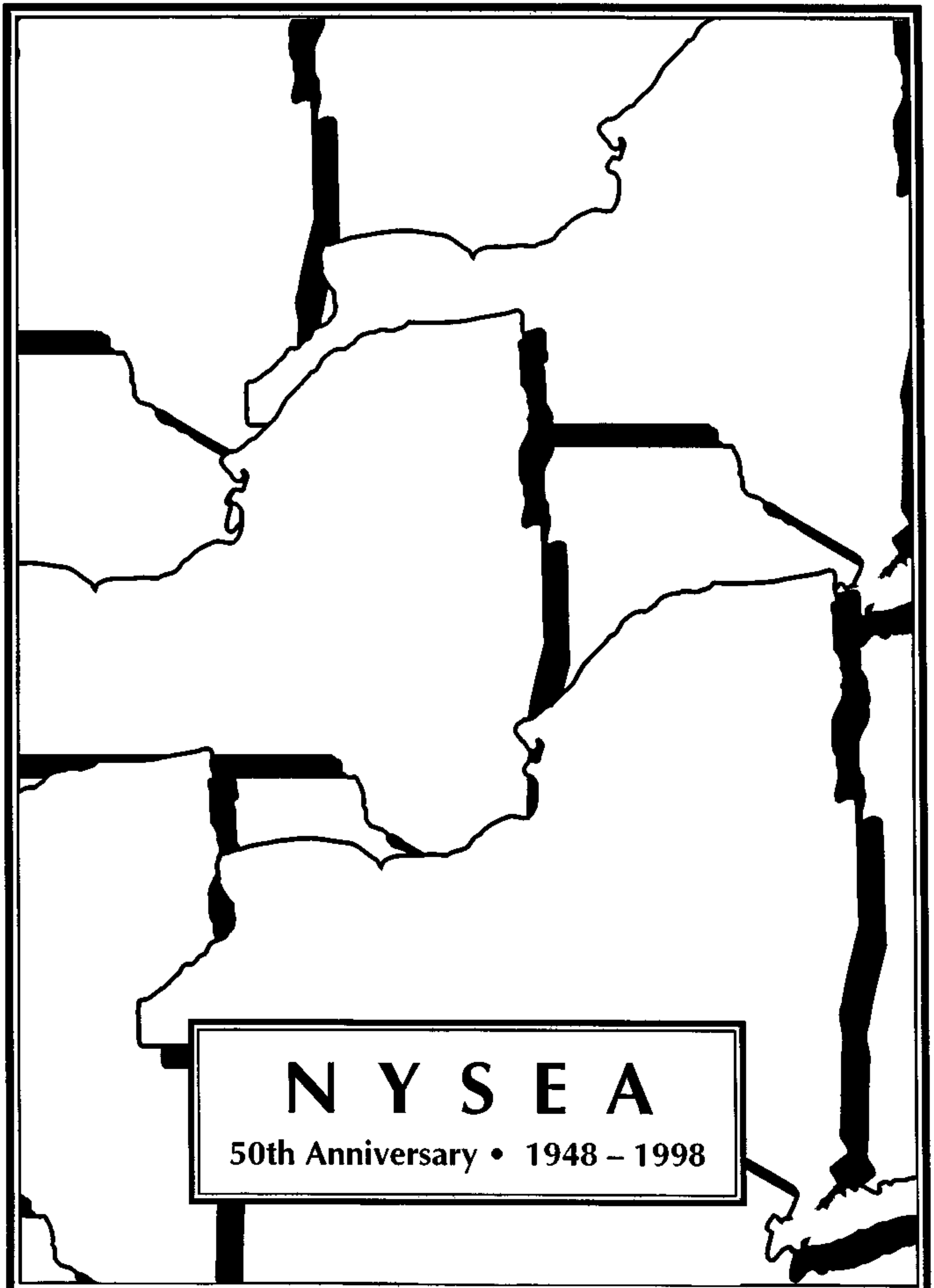
SESSION

1:45 – 3:15 PM **Quebec Economic Development: Issues and Policy** (Craven Lounge)

Chair and Moderator: Myrna Delson-Karan,
Quebec Government House

Presenters:
Paul Clermont, Director of Export Promotion,
Department of Industry, Commerce, Science & Technology,
Government of Quebec

Randall Capps, Western Kentucky University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