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The Price Elasticity of Demand for Whole Life Insurance

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ABSTRACT

In this study a real price index is created for whole life insurance sold in the United States from 1953 to 1979. New purchases of whole life insurance are shown to be negatively related to changes in this cost index, contrary to what has been widely accepted in the insurance literature, but consistent with economic theory. The existence of strong price elasticity of demand for whole life insurance does not ensure, however, that the insurance industry manifests a high degree of price competition.

FOR MANY YEARS THE insurance trade literature has been replete with assertions that life insurance consumers are insensitive to variations in price when making their insurance purchasing decisions. While many of these assertions had their foundation in the lore of sales agents, more recent allegations have been based upon the responses of life insurance consumers to surveys [29, 30] and the demonstration of wide price disparities among various companies offering similar products.

The predominant view in the insurance industry stands in contrast to a large number of theoretical studies appearing in the economic literature in which insurance cost invariably appears as a principal determinant in the consumer demand function (e.g., [12, 17, 35]). However, in economic time-series and cross-sectional statistical studies, the price of insurance has never appeared as a determinant of demand, to my knowledge. Indeed, this finding holds true even for empirical studies joined with theoretical analyses in the same article! The importance of price is given lip service, but fails to have been statistically measured.

It is the purpose of this paper to provide an initial examination into the price sensitivity of consumer demand for whole life insurance. The approach taken involves time-series analysis based on annual aggregate data over a sample of life insurance companies. The first section contains a brief review of some of the salient economic and insurance industry literature and a discussion of aspects of the problem to be investigated. This section is followed by a disquisition on real price indices. A cost index proposed elsewhere is modified and reestimated for the purpose of measuring the price elasticity of life insurance demand. In the

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third section, the statistical models used in testing are presented, and the nature of the insurance purchases data is discussed. The next section presents the results of the statistical estimations. Conclusions and directions for future research are provided in the final section.

I. Related Literature

Estimations of the price elasticity of life insurance demand do not appear to have been published previously in the literature. However, this has not precluded conjectures on the subject. In the insurance literature, life insurance has often been characterized as a "sold good, not a bought good" [22]. The meaning of this aphorism is that the initiative in a life insurance transaction in the ordinary (individual) life insurance market with which we are concerned here comes typically not from the buyer but from the seller [48]. Coupled with the finding that the complexity of the typical whole life contract renders it unamenable to easy price comparisons [2], it has been inferred that consumers are insensitive to variations in the price of life insurance.[1]

Other evidence has suggested that the life insurance consumer is not especially sensitive to price. Numerous studies have been conducted into the competitiveness of policy pricing among insurers; most have found wide ranges in the prices of ostensibly similar policies [2, 7–9, 26, 43].[2] This finding has been met with increased efforts at cost disclosure, in an attempt to provide sufficient information for insurance consumers that the discipline of market forces upon life insurers is promoted [10, 13, 15, 20, 22, 41, 47]. Even with these efforts at information dissemination, many consumers remain confused, and few make the price comparisons that characterize consumer shopping behavior for other goods and services of similar price [1, 29, 30, 39]. At the extreme, some have even claimed that the very existence of whole life insurance contracts is *prima facie* evidence of consumer irrationality! (The notion behind this assertion is that a rational person could be economically ahead by purchasing term insurance and investing the difference. Such a view does not take into account the many options that are part of the whole life policy [5, 44].)

The survey findings that extensive price searches are generally not conducted

[1] Examples of other references dealing with these and related issues are [14, 23, 34, 42].

[2] One author takes exception to the widespread finding of others of disperse life insurance prices [46], and gives an upper bound of less than 3.1 percent price dispersion for whole life insurance sold in America. His figures include an adjustment for varying firm costs, and thus do not reflect prices from a purely consumer-oriented point of view. Additional studies document some of the sources of variations in firm costs [25, 37]. Beresford [10] has found some evidence of high price dispersion for small policy sizes, and low price dispersion for large policy sizes; he hypothesizes that the pricing patterns do not indicate a lack of price competition, but rather may reflect imperfect discrimination by firms among heterogeneous consumers [41]. He further hypothesizes that the result of this discrimination may be that the life insurance market is segmented, with some companies selling low cost policies to low cost buyers (with lower lapse potential, lower mortality, higher policy size) and some selling high cost policies to high cost buyers. Using the cost index developed in this paper, which takes a consumer viewpoint, I have independently confirmed the finding of widespread price dispersion for whole life insurance products, at least in the policy size range of 10–25 thousand dollars.

prior to purchase are not necessarily indicative of consumer irrationality in their purchase behavior. Information economics theory relating to price search when the distribution of prices is known and unknown [11, 39, 40, 41, 45] supports the notion that nonextensive price searches may be optimal, especially when the price dispersions are low [46], when the comparisons are complex, and when there are sizable costs incurred by the search.[3] It should also be emphasized that lack of price search, *ipso facto*, does not necessarily imply a lack of price sensitivity in the consumer demand function for life insurance. A consumer's decision of whether to purchase and how much to purchase from a particular company could still be responsive to the perceived price, regardless of whether price comparisons have been made. It is consistent with consumer rationality both to adjust the quantity demanded to the perceived price and to conduct no price search, provided that the aforementioned factors that would limit the optimal search are present.[4]

Although virtually all theoretical treatises on life insurance demand have identified price as an important factor,[5] empirical research has heretofore omitted this factor in demand studies (e.g., [3, 16–19, 21, 27, 28, 33, 36]). This omission is not surprising in light of the fact that until recently an appropriate time-series price index did not exist.[6] These and other studies have attempted to link numerous variables to life insurance demand, including industry advertising expenses, size of the industry sales force, introduction of new products, disposable permanent and transitory income, consumer expectations, new household formations, marriages, births, net savings rates by the household sector, current yields on alternative assets, expected inflation, race, geographical location, nationality, and education. Many of these variables, which have some importance in cross-sectional studies (e.g., geographical location, nationality, education, race, introduction of new products), are unimportant in a time-series national study using aggregate data. Others (e.g., industry advertising expenses, size of industry sales force) seem to impact sales mostly over short time periods (monthly, quarterly). Finally, still others (e.g., births, marriages, household formations, expected inflation, interest rates) have been shown to exhibit ambiguous impact upon life insurance sales, depending upon the time period under study, and the lag specification of the variable. Only income consistently appears to have been linked positively to life insurance sales.

Of the aforementioned empirical studies, none has contained direct estimates of price elasticity of demand. However, by assuming that the supply price of

[3] Dozens of indices have been devised for ranking policy costs, and many generate conflicting rankings [7, 24, 46, 47]. The costs of search include the time required to acquire, absorb, and process information.

[4] The optimal number of price searches, which in the case of life insurance will also be a function of the size of the policy desired, is the subject of a future paper.

[5] Many of the important studies are cited in the reference section of [4].

[6] The new index designed for time-series analysis appeared in [6]. One earlier index that was estimated [32] did not take into account changes in mortality rates over time, although use of that index in conjunction with a mortality rate trend variable could perhaps remedy this omission in econometric testing of insurance demand. However, the index was designed with other purposes in mind, focusing on the cost of persisting with an already purchased policy during the tenth and twentieth policy years.

insurance followed a linear trend, Fortune [17] does include a simple time variable to control for price effects. In light of what is now known (i.e., the real price of insurance has not followed a linear trend, but has exhibited substantial oscillation [6, 32]), it would seem prudent to test directly for price sensitivity. The closest approximations in the literature to date are examinations of anticipated inflation's impact upon life insurance demand.[7] Admittedly, anticipated inflation is only one element affecting rationally perceived real insurance cost, and its effect may be offset by other factors such as changes in policy terms and provisions, mortality rates, and real interest rates.

II. Whole Life Price Indices

The considerations important in the construction of a real price index for life insurance have been dealt with at length in [6, 32], and related sources cited in those papers. The price index adopted in this paper is that of [6], chosen because it incorporates changing mortality experience into the rates and because its focus is on the rationally perceived cost of newly issued policies.

This index was modified slightly so that its construction would be more directly compatible with most of the theoretical and many of the empirical studies on life insurance demand, which relate real insurance in force to the real cost per unit of (real) insurance in force. The published index was presented as a ratio of the present value of expected premium cost, net of dividends and accumulations of cash values, per present-valued unit of indemnification benefits expected to be received, in excess of the actuarially fair cost. The transformed measure gives net cost per 1000 present-valued units of insurance expected to be in force over any arbitrary time horizon τ selected.[8] The exact formula used is presented in an appendix to this paper, but can be given in compact form as:

$$\mathrm{NPC}_\tau = C_\tau - B_\tau \tag{1}$$

where

- NPC denotes the net present cost per 1000 present-valued dollars of insurance expected to remain in force during the period $[0, \tau]$,
- C is the present value of expected costs per 1000 present-valued dollars of insurance expected to remain in force over the period $[0, \tau]$, and
- B is the present value of expected insurance benefits per 1000 present-valued dollars of insurance expected to remain in force over the period $[0, \tau]$.

[7] See, for example, [3].

[8] The present value of 1000 nominal dollars of insurance in force declines as the time horizon is increased. Thus, in order to have an amount of insurance in force that averages over time 1000 dollars in present value terms, a policy of larger nominal size must be purchased at the outset. This kind of adjustment to the price index is necessary if one is to avoid introducing "policy illusion" (i.e., the confusion arising from focusing on the policy rather than the underlying insurance) into the indices.

Probabilities of persistence were not incorporated into the price indices due to the lack of appropriate data. While data do exist relating to lapse experience, and are in fact used by actuaries in the pricing of policies, these data are unlikely to reflect rational consumer expectations at the time of policy purchase (see [6, pp. 154–55] for a detailed explanation).

The above formula entails the assumption that if the insured survives, he will surrender the policy at the end of year τ for its full surrender cash value, having contracted no policy loans during the interim. The price indices for each company in the sample were reestimated using this formula, for projected holding periods of 10 and 20 years. Alternative indices were computed using double-A-rated corporate and prime grade municipal bond 10-year yields-to-maturity for discounting expected future cash flows (see [6], footnote 20 for detail).[9] Aggregate price indices were then formed, where the price for a new policy from each company was weighted by the volume of new insurance in force sold by that company relative to the volume of new sales of companies comprising the entire sample. A second group of indices was also computed, where it was assumed that incremental cash values would be lent to the insured each year that the policy was projected to remain in force; the formula used for this calculation is an adaptation of Equations (5) and (6) of [6]. These indices were then compared with the corresponding indices that did not allow for policy loans, and the lower of the figures for each year was used in constructing "optimal" price indices. (Here the word "optimal" is used to denote that policy loans are only undertaken when they will lower the cost of insurance to the consumer.)

In this paper, the same sets of simplifying assumptions and abridged policy values used in [6] were utilized here, and readers are referred to that article for a more detailed discussion of the implications and limitations involved in their usage. Policy values (premiums, cash values, and dividends, if any) were obtained on the whole life policies of 22 stock and five mutual insurance companies. The 22 stock companies were selected at random from 66 companies for which comparable data have been continuously available since 1953. The five mutual companies chosen are among the six analyzed by Lin, who selected them based on their large size. Northwestern Mutual was excluded due to data interruptions that occurred subsequent to his study. New insurance purchases from these two samples each accounted for approximately one-fifth of the total new whole life insurance sold to Americans by all stock and mutual companies over the period under study. Published policy values for a male of age 35 were used in constructing the price indices. Although one might suppose from the principle of actuarial equivalence that prices are similar across all ages, substantial differences can be observed and these differences are not necessarily consistent across companies. Nonetheless, we found in separate tests that while there were price differences across ages, the changes in prices across all ages were highly correlated for each of the companies sampled. Thus, the likelihood of any severe distortions being introduced into the time pattern of the price indices by using the representative policy data of a 35-year-old male is low. Another compromise that was made in constructing the indices was necessitated by the lack of full dividend data for the mutual companies. Typically, dividend data for some of the later policy years were omitted from the published reports and had to be filled in with approxi-

[9] While the formula used in the creation of the indices can accommodate yields that differ by maturity, this preferred procedure was not implemented due to project constraints. Nonetheless, any distortion from using yields to maturity rather than forward rates is likely to be very minor relative to the fluctuations in cost produced by changes in the levels of various formula parameters.

mations. The technique used involved finding full dividend schedules and determining from them the typical patterns of accrual over the life of a policy, and then extrapolating from these patterns to create values for the missing periods. The indices estimated here retain all of the features of the former set that rendered it attractive for time-series study; indeed, the only two alterations of importance are the modification from an equally-weighted to a sales-weighted index, and the change in the numeraire from an expected cash inflow to an amount of insurance in force. These indices manifested correlations in the mid-0.80 range with those estimated previously in [6].

III. Empirical Model and Data

In testing for price elasticity in the demand for life insurance, potential problems of identification emerge. First is the often encountered identification problem of disentangling demand and supply curves. Fortunately, for the case under consideration, life insurance aggregate supply curves are generally regarded as being flat for reasons elucidated elsewhere [3, 17]. Accordingly, movements in the price of life insurance, indicated by changes in its real price index, are associated only with movements of the entire supply curve, and this potential source of identification problems disappears.

Another source of potential identification problems is the well-known fact that the whole life policies sold by the life insurance industry consist of components of pure insurance, savings, and various options [44]. Most empirical studies have featured attempts at separating the pure insurance component from the savings and other components when using aggregate data of insurance in force. However, the present study deals only with newly sold policies, none of which exhibit any surrender cash values during the first policy year (although policyholders who buy cash value policies obviously have future cash values in mind). As the focus of this study is on the demand for the whole life policy[10] as a unitary package, no further adjustments were undertaken to avoid the problem of identification.

The empirical model presented below was designed to conform to theoretical propositions that have been set forth elsewhere. While it would be inappropriate to belabor here the extensive underpinnings for the model that are supplied by the numerous studies (identified earlier in this paper), including three by this author, it shall suffice to indicate the major themes (in the form of partial derivatives) emerging from these theoretical studies (e.g., [17]). First, shifts in the demand curve for real insurance coverage are positively related to changes in human capital or permanent income under a wide range of circumstances. Second, except under extreme risk aversion, the real quantity of insurance in force demanded, whether provided by whole life, term, or some other insurance vehicle, is a decreasing function of its real price, Third, demand for life insurance is theoretically related to the amount of nonhuman (inheritable) wealth possessed by the insured; the sign of this relationship is ambiguous and dependent upon the shape of the absolute risk aversion function of the insurance consumer.

Other possible determinants of demand that have occasionally appeared in

[10] In a separate study the demand for term insurance is empirically examined [5].

time-series studies were not included. Some of these, such as mortality rate trends and interest rates were already included as factors in the real price series. Sociodemographic factors such as population growth, births, marriages, and household formations were not found to be statistically significant factors in preliminary testing, despite various imposed lag structures on these variables; thus, they were excluded from the final version of the model. Data pertaining to social security benefit levels and nonhuman wealth were also tested in preliminary versions of the regression model, but found to be statistically insignificant. The measure of nonhuman wealth used was similar to that used by Fortune and Beresford and suggested by Barro: net stocks of fixed, nonresidential capital plus net stocks of consumer durables, both in constant dollar terms. This wealth variable was so highly correlated with the real disposable personal income variable ($r = 0.996$) that it introduced collinearity into the regression model and rendered the estimated coefficients on the wealth and income variables unstable, while having only minor impact on the coefficient associated with the real price variable. As our interest was primarily upon the price sensitivity of insurance demand, this variable was excluded from further consideration.

The final version[11] of the model to be estimated was:

$$Q_t = \alpha_0 + \alpha_1 Y_t + \alpha_2 P_{t,WL} + \epsilon_t \tag{2}$$

In this model the dependent variable, Q_t, represents the real amount of new insurance in force written during year t. Data for this variable were taken from *Best's Life Insurance Reports* covering the period from 1953 through 1979. Unfortunately, these data relate to new sales of both whole life and endowment policies; disaggregated figures for these two types of insurance were not available. Nonetheless, the prices of these two forms of life insurance have been found to be highly correlated, and therefore it appears unlikely that any serious consequences will arise in using this data series. Further, the endowment policies comprise a trivial portion of the policies sold in this category (one percent in 1982 [31]). The sales data were collected for each of the companies that comprise the insurance price indices and separated into two groups: participating and nonparticipating policy sales. These data were deflated by the personal consumption expenditure deflator to render them in constant dollar terms.

The first independent variable, Y_t, represents the real amount of aggregate personal disposable income in the United States at time t, and is used as a proxy for human capital. The same personal consumption expenditures index was used to deflate this series, but because the income data are by nature flow rather than stock variables, the end-of-year values of the price index were not used to adjust the income figures; rather, the yearly average index numbers were used. A second version of this variable was constructed to proxy for permanent income. In this case, the commonly used technique (e.g., [17]) of taking a three-year moving average of the real income figures was used, $\bar{Y}_t$, where the midpoint of the three-year average centered on time t. It was reasoned that since the dependent variable

[11] Several functional forms of the relationships under study could have been used, as there is a lack of compelling theory that would specify a single preferred form. Several forms, including the popular loglinear form, were experimented with, but all provided essentially the same implications as the linear form presented here, although their fits were not as good.

is an end-of-year stock variable, the insurance consumer would have a very good idea of the amount of income to be received during the ensuing year; thus, the third year's actual income figure was used as a proxy for this expectation as part of the three-year moving average.

The second independent variable, $P_{t,WL}$, represents various measures of the real price of newly issued whole life insurance at time t, described at length earlier in this article. Separate sales-weighted indices were estimated for participating and nonparticipating forms of whole life insurance. ϵ_t is a disturbance term.

One aspect of the insurance price indices that influences the method of estimating the model is that the policy data published in *Best's Flitcraft Compend* and used in developing the insurance price indices are nonsynchronous with the insurance purchasing data. Policy values are revised at various times during the year, but the policy values used in the computation of the price indices were those in effect at the time of editing the reports for publication. While the figures on new insurance in force written during a year include policies purchased throughout the year (that have not lapsed) at potentially different prices, the insurance price index calculations used only those policy data effective at a particular point in time during the year. Close analysis of the nature of these data led us, on *a priori* grounds, to expect correlation in the error terms of up to a second-order moving average variety. This possibility should be accommodated in the method of estimating the model. A generalized least squares technique correcting for a second-order moving average process in the error terms was used for the regressions.[12]

IV. Test Results

In selecting which version of the price variables to test, there was little theoretical guidance available. While theory suggests that price is crucial to the demand for insurance, it does not specify clearly which of the price indices should be most relevant, other than that rational consumers should place more weight upon the ones that allow for policy loans. All told, eight prices indices were estimated for each kind of whole life insurance sold, where the expected holding periods varied from 10 to 20 years, municipal or corporate bond yields were used to discount expected cash flows, and where policy loans were or were not allowed. Each of these indices was used, in turn, with either a single-year income figure or the three-year moving average income figure. As was expected, since the price indices were correlated with each other to a high degree, it did not really make much difference which price variable was utilized in the regressions; the coefficients on all of them carried the same sign, although the preferred indices that allowed for policy loans generally exhibited higher statistical significance. Further, in terms of the regression statistics (estimated coefficients, t-statistics, coefficient of determination), it appeared to make almost no difference which of the two income variables was used. In this paper, only the results associated with the three-year income variable, the price indices using a 10-year expected holding

[12] See, for example, Pindyck and Rubinfeld [38, pp. 515–19].

Table I

Regression Results

Generalized Least Squares Technique, Second-Order Moving Average Process

(1) Policy Type	(2) α_0 (*t*-statistic)	(3) $\alpha_1(\bar{Y})$ (*t*-statistic)	(4) Beta (elasticity)	(5) α_2 (P_{WL}) (*t*-statistic)	(6) Beta (elasticity)	(7) *F*-statistic	(8) $\rho 1$ $\rho 2$	(9) CRSQ D.W.
Nonparticipating	3.90	0.0060	1.21	−0.0466	−0.49	57.8	−0.36	0.814
	(6.6)	(10.0)	(0.85)	(−4.1)	(−0.71)		−0.12	1.99
Nonparticipating*	5.88	0.0044	0.89	−0.0730	−0.50	4106.8	0.52	0.997
	(36.1)	(70.8)	(0.62)	(−28.1)	(−0.92)		0.48	1.77
Participating	1.08	0.0083	1.26	−0.0327	−0.32	72.9	−1.16	0.847
	(2.5)	(10.7)	(1.09)	(−2.8)	(−0.32)		−0.52	1.54
Participating*	2.12	0.0075	1.14	−0.0577	−0.26	91.7	−1.18	0.875
	(3.8)	(13.5)	(0.98)	(−3.9)	(−0.42)		−0.46	1.68

* denotes price index used in regression allowed for policy loans.

period, and the medium-term corporate bond yield in the discount function are reported.

The regression results presented in Table I suggest confirmation of the theoretical propositions listed earlier. All variables were measured statistically significant at well above the 95 percent level of confidence and exhibited the correct signs.

In reviewing the statistics relating to the income variable (see figures in parentheses in Column 4 of Table I), it is noted that the income elasticities for both nonparticipating and participating whole life demand decrease from a range of 0.85–1.09 to a range of 0.62–0.98 when the preferred price index is used as a co-regressor. Accompanying this modest decline in income elasticities were increases in the *t*-statistics (see figures in parentheses, Columns 3 and 5) on *both* the income and price variables of sevenfold for nonparticipating and approximately 30 percent for participating life insurance.

The variable of principal interest, heretofore untested in empirical studies, is the real price of whole life insurance. For both types of policies, prices were negatively related to new purchases (see Column 6), exhibiting elasticities ranging from −0.71 to −0.92 for nonparticipating and −0.32 to −0.42 for participating policies, depending upon the price index used in the regression. Of particular note is the finding of price elasticities for nonparticipating insurance that were more than double the magnitude of those for participating insurance. This rather curious result can be explained by further examining the nature of the price indices. It is recalled that with participating policies, the insured is in effect purchasing a partial hedge against a rise in interest rates, as life insurance companies tend to increase their dividends during times of higher profitability. The price indices were based on dividend scales effective at the beginning of a policy horizon, but in practice, dividends were more than doubled from those levels in recent years. The sales literature is replete with references to this practice, and it would be reasonable for consumers of participating policies to perceive prices as fluctuating less than indicated by movements in the indices. Accordingly, the elasticities are likely to be downward biased. Nonparticipating policies feature only contractually fixed values, so the price index is not subject to this bias. It is reassuring that the price indices relating to the latter exhibited greater statistical significance. A second, and more plausible, explanation for the higher price elasticities of the nonparticipating policies is that prices for such policies are much easier to compare because they have been shown to be less dispersed and because evaluation of dividends is not required.[13]

Income and price elasticities do not always give a good indication of the relative impact of income and price on purchases, as they are both tied to percentage changes in the underlying variables; where one variable has fluctuated in percentage terms substantially more than another, these elasticity measures are unable to provide impact comparisons. In Table I, beta statistics for both the income and price variables are reported; these statistics indicate each variable's influence upon the dependent variable in standardized units.[14] There it is shown (see Column 6) that a change of one standard deviation in the price variable was linked to a movement of from 26–32 percent (for participating) to 49–50 percent (for nonparticipating) of a standard deviation in new insurance in force written. The influence of income changes (Column 4) was associated with betas of 1.14 to 1.26 for participating and 0.89 to 1.21 for nonparticipating insurance, depending once again upon the co-regressor used. Thus, in standardized units, changes in real disposable personal income had roughly the same impact upon new purchases of insurance, whether through participating or nonparticipating policies.

The coefficients of determination, corrected for degrees of freedom, indicate that reasonably good fits were obtained, especially when considering the fact that all equations were estimated in deflated terms and with only two exogenous explanatory variables. There were some items, however, that did give rise to some concern. First, it was noted that the estimated coefficients on the income variable bounced when alternating among the co-regressor price indices. This is an indication of possible collinearity between the independent variables. An inspection for this found that the income variable showed much higher correlation with the price indices that did not allow for policy loans; accordingly, more confidence can be placed in the stability of estimations generated from the preferred price indices. Second, the moving average correction factors did not possess logical signs or size, except in the case of the nonparticipating insurance sales regressed on the preferred index. When the other price index was used as a regressor, the moving average correction factors had negative signs, but the signs turned positive when a better regressor was used. It is suspected that in the case of participating policies, the negative correction factors were a reflection of the looser fit of the regressions, and that if price indices were available that possessed a degree of credibility similar to that for nonparticipating insurance, this situation also might have been reversed.

[13] I am indebted to the referree for this point.

[14] The beta coefficients are computed by multiplying the estimated regression coefficients by the ratio of the standard deviations of their associated independent and dependent variables.

V. Epilogue

Based upon the statistical evidence presented above, we cannot reject the hypothesis that consumers are indeed responsive to price in the amount of insurance they buy. This finding contrasts with what has been a long standing tradition in the insurance trade literature.

Before this verdict can be deemed incontrovertible, it would seem prudent to undertake further studies of a more micro variety. Insurance sales and price data could be examined in disaggregated form (by company). In such studies, it would be extremely important to take into account other factors such as the sociodemographic data and new product introductions that could be crucial to the observed patterns of purchases. For example, migration tendencies could obscure the effects of other variables for a geographically concentrated insurance carrier, or a new product introduction could severely distort the data. In the study conducted here, where company operations were dispersed throughout the United States (especially for the nonparticipating company sample), these considerations were not binding. Nonetheless, the results of this study should be viewed with caution, as they depend upon annual and highly aggregated data. This dependence presents a significant problem because the market for life insurance may have changed substantially over the more than two decades needed to obtain enough data points to estimate econometric relationships. For example, over the period analyzed in this study the structure of the life insurance market underwent dramatic changes. Life insurance price comparisons methods were not well known during the first 10 years of the sample period and were not widely used until the last five. In addition, the pace of product innovation began to accelerate in the early 1970s. The present study relegates such items to the disturbance term. However, even in a national study such as this, the recent introduction of universal life insurance policies had such a spectacular impact and was so widespread that it would need to have been taken into account, were it not for the fact that our study period ended before its introduction.

Finally, it should be reemphasized that the strong negative price elasticity uncovered in this paper should not be taken as an indication of price competition in the insurance industry, as nothing in this study would suggest that consumers actively seek out the lowest cost insurance carriers. Indeed, price and sales data were disaggregated by company to check explicitly for this phenomenon. Company sales data (deflated) were regressed against industry sales data and the difference between company price and industry average price in a pooled cross-sectional time-series regression. No evidence was found of consumers shifting from high to low cost insurance carriers. Obviously, this finding is only preliminary and other market-related variables would need to be taken into account to confirm it.[15]

[15] In a study conducted contemporaneously with this exercise, Beresford [10] has taken a micro focus precisely along the lines suggested above. He did encounter evidence of consumers shifting from high to low cost insurance carriers over time, when controlling for a large number of market-related variables. The price index used by Beresford, however, was the "Interest-Adjusted Cost Index"

Appendix

In this appendix, a detailed rendition of Equation (1) is provided, along with some brief explanation. For a more complete explanation of the formula, the reader is referred to [6]. The net present cost of insurance coverage when no policy loans are undertaken, per 1000 present-valued units of insurance expected to be in force over time period $[0, \tau]$, is given by:

$$NPC_\tau = \frac{\tau i}{1 - \exp(-i\tau)} [P_\tau - D_\tau - TD_\tau - DB_\tau - CV_\tau - TDD_\tau] \quad \text{(A1)}$$

where

$$P_\tau = \sum_{t=1}^{\tau} \frac{s_{t-1}p_t}{\frac{1}{(1 + i_0)} \prod_{j=0}^{t-1} (1 + i_j)} \quad \text{(A2)}$$

$$D_\tau = \sum_{t=1}^{\tau} \frac{s_{t-1}d_t}{\prod_{j=1}^{t} (1 + i_j)} \quad \text{(A3)}$$

$$TD_\tau = \sum_{t=1}^{\tau} \frac{s_{t-1}q_t td_t}{\prod_{j=1}^{t} (1 - i_j)} \quad \text{(A4)}$$

$$DB_\tau = \sum_{t=1}^{\tau} \frac{s_{t-1}q_t 1000}{\frac{\sqrt{1 + i_t}}{(1 + i_0)} \prod_{j=0}^{t-1} (1 + i_j)} \quad \text{(A5)}$$

$$CV_\tau = \frac{s_\tau cv_\tau}{\prod_{t=1}^{\tau} (1 + i_t)} \quad \text{(A6)}$$

$$TDD_\tau = \frac{s_\tau td_\tau}{\prod_{t=1}^{\tau} (1 + i_t)} \quad \text{(A7)}$$

and where

- i_t is the appropriate discount rate in year t for after-tax cash flows occurring during the year, adjusted only for default risk;
- p_t is the premium payable at beginning of year t, per 1000 dollars of insurance in force;
- cv_t is the cash value at the end of year t, per 1000 dollars of insurance in force;
- τ is the year of intended policy surrender;
- d_t is the dividend payable in year t, per 1000 dollars of insurance in force;

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published in many states, rather than the index put forth here. While admitting that the Interest-Adjusted Cost Index exhibits many shortcomings when compared with other models, he opted for the index because it was readily available and it is also the one that consumers are most likely to encounter in sales presentations. Unfortunately, no price elasticities are reported by Beresford, but a statistically significant shifting pattern was detected in the study. Thus, his results can be viewed as complementary to those presented here.

td_t is the terminal dividend payable at the end of year t (or during year t in the event of policy surrender (or death));
s_t is the probability of survival through the end of year t; and
q_t is the probability of death in year t given survival through year $t - 1$.

In a second version of this formula which allows for policy loans, Equations (A5) and (A6) were replaced by Equations (5) and (6) of [6]. For those years where this version of the formula resulted in lower cost figures for the consumer, they were substituted in the "optimal" price index. The first term in the product of (A1) is an adjustment factor for rendering the declining present value of insurance in force over time into an average present value of 1000 dollars. The expression represented by (A2) is the present value of premiums that are expected to be paid up through year τ. These premiums are paid contingent upon survival of the insured. (A3) gives the present value of expected dividends to be received as long as the policy remains in force. (A4) and (A7) represent the present value of the expected terminal dividend, which is available in some policies that have been in force for 10 years or more and received either upon death of the insured or policy surrender, respectively. In (A5), the present value of expected insurance indemnification is given, while (A6) gives the present value of the cash value to be received upon policy surrender, adjusted for the likelihood of survival through year τ.

In aggregating the individual company indices, a final adjustment was made to weight each index by the amount of new insurance in force written during the policy year, relative to that written in the other companies that comprised the sample. Separate mortality tables were used for each of the years from 1953 through 1979 in an attempt to capture evolving mortality patterns over time. Thus, new vectors of s_t and q_t were employed for each year under study.

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