Does the Internet Make Markets More Competitive? Evidence from the Life Insurance Industry

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The Internet may significantly reduce search costs by enabling price comparisons on-line. This paper provides empirical evidence on how Internet comparison shopping sites affected the prices of life insurance in the 1990s. With micro data on individual insurance policies and with individual and policy characteristics controlled for, hedonic-type regressions show that increases in Internet use significantly reduced the price of term life insurance. Further evidence shows that prices did not fall with rising Internet usage in the period before the sites began, nor for insurance types that were not covered on the sites. The results suggest that the growth of the Internet has reduced term life prices by 8–15 percent. The results also show that the initial introduction of the Internet search sites is initially associated with an *increase* in price dispersion within demographic groups, but as use spreads, the dispersion falls.

We would like to thank Eric Anderson, Judy Chevalier, Mark Duggan, James Garven, Robert Hartwig, Thomas Hubbard, Ken Isenberg, Kent Jamison, John Johnson, Peter Klenow, Olivia Mitchell, Jim Poterba, Todd Sinai, Alan Sorensen, Mark Warshawsky, Alwyn Young, and numerous seminar participants for helpful comments and Jeffrey Butler, Andrew Lee, and Soojin Yim for excellent research assistance. We would also like to thank Ken Isenberg and LIMRA International for assistance with data from the LIMRA Buyer Studies and Forrester for access to the Technographics database. For research support we are grateful to the Kennedy School Dean's Faculty Research Fund and the TIAA-CREF Institute (Brown) and the National Science Foundation (grant SES-9984567), Alfred P. Sloan Foundation, and Centel/Robert Reuss Foundation (Goolsbee).

[*Journal of Political Economy*, 2002, vol. 110, no. 3]

I. Introduction

The last five years have witnessed an explosion in the growth of electronic commerce and Internet marketplaces as alternatives or supplements to traditional retail markets (McQuivey et al. 1998). Consumers can now go on-line and comparison shop between hundreds of vendors with much less effort than in the physical world. The traditional economic view suggests that, as a result, the Internet should reduce search costs for consumers and thereby reduce prices and make markets more competitive.

Despite this presumption of increased competition, however, existing empirical work about the Internet has not been as supportive of the theory as one might expect. Although data availability has limited analysis of the sector (existing work has mainly entailed comparing prices on- and off-line for a specific category such as books), the results from this literature have not conformed to the traditional view of falling search costs. These studies have generally found large dispersion of prices on-line and prices either modestly lower or actually higher than their off-line counterparts.[1] To the extent that there is a conventional wisdom in such a new area, it is that the Internet may have fostered product differentiation and price discrimination more than price competition.[2]

Because of the data constraint, however, little is known about the impact of the Internet on off-line prices. Instead, most papers take off-line prices as exogenous. In this paper, we shall present the first empirical evidence on the impact of Internet competition on prices and dispersion off-line. In this sense, our results are similar to those of the existing empirical literature on search.[3] By combining Internet and life insurance industry data sets over time, we are able to document how important the Internet—and the reduction in search costs that it creates—can be for market competition.

We examine term life insurance, a somewhat homogeneous product

[1] Work by Lee (1998) on cars and Bailey (1998) on books, compact discs, and software suggests that prices were actually higher on-line than in retail stores. More recent work by Brynjolfsson and Smith (2000) on books and compact discs and by Clay, Krishnan, and Wolff (2001) on books has found that prices are the same or lower on-line but on-line price dispersion is quite high, perhaps greater than in retail stores.

[2] See the work of Bakos (1997, 1998) or the survey of Smith, Bailey, and Brynjolfsson (2000). Although addressing a different question, the results of Goolsbee (2000, 2001) suggest that on-line buying is sensitive to local retail price variation and to local sales tax rates. Contemporaneous work by Ellison and Ellison (2001) and Smith and Brynjolfsson (2001) analyzes consumer demand using data from shopbots for books and for computer memory. Chevalier and Goolsbee (2001) also look at the book industry.

[3] This includes the work on the impact of price advertising on pricing behavior such as Benham (1972), Kwoka (1984), or Milyo and Waldfogel (1999). It also includes other work exploring the sources of price dispersion such as Pratt, Wise, and Zeckhauser (1979), Dahlby and West (1986), Van Hoomissen (1988), or Sorensen (2000).

with high search costs and relatively high markups, for several reasons. First, in the mid 1990s, a group of Internet price comparison sites began that dramatically lowered the cost of comparing the prices (i.e., premiums) of similar term life policies across companies. This can have an important impact in an industry in which high customer search costs create the potential for market power among existing merchants. Second, life insurance as an industry is quite important in its own right. It is one of the most widely held financial products in the United States, and the face value of life insurance policies sold in 1999 exceeded $2.5 trillion. Premia typically amount to several percent of gross domestic product annually (see Cawley and Philipson 1999; American Council of Life Insurance 2000). If the Internet reduces prices in this market, the potential welfare implications are enormous. Third, there has been a very serious price decline in the cost of term life insurance in the 1990s that is not well understood and has taken place concurrently with the rise of the Internet (see the description in Dugas [1999]). We shall try to examine the ways in which the two are related.

To analyze the relationship, we take individual policy-level microdata from LIMRA International on the prices of insurance policies as well as various owner and policy characteristics and match them to microdata on the growth of Internet usage and on-line insurance research from Forrester by the same owner characteristics. In essence, we fit hedonic regressions for the price of life insurance on characteristics of the policies and the individuals and then include a measure of how likely the individual is to have used the Internet over time or to have researched insurance on-line.

The results indicate that once the on-line insurance sites began, the faster a group adopted the Internet, the faster prices of term life insurance fell for that group. The total impact of the rise of Internet use from 1995 to 1997 reduced term life prices by 8–15 percent. This implies an increase in consumer surplus of about $115–$215 million annually for these specific policies. The results are robust in that rising Internet use did not have any effect on prices during the period before the insurance web sites existed, nor did it affect the prices of types of life insurance that were not covered by the web sites (i.e., whole life policies). Neither can the results be explained by changes in mortality across groups. Interestingly, the data also show that the Internet-induced reduction in search costs actually increased price dispersion on introduction. As it became more widespread, price dispersion fell.

The paper proceeds as follows: In Section II, we discuss the life insurance industry and the role of the Internet comparison sites. In Section III, we discuss the theory of search when customers have different search costs. In Section IV, we discuss our data sources and the basic specification. In Section V, we present the basic results. In Section VI,

we consider alternative explanations of the results. In Section VII, we examine price dispersion within groups. In Section VIII, we present conclusions.

II. The Life Insurance Industry and the Internet

A. *Overview of Life Insurance*

The market for life insurance is the largest private individual insurance market in the world. In 1999, nearly 59 million life insurance policies were purchased in the United States, with a face value of over $2.5 trillion, bringing the total number of policies in force to 366 million, with a total face value of $15.5 trillion (American Council of Life Insurance 2000).

Life insurance can play a number of important roles in the portfolios of most households. The primary function of life insurance is to protect a primary earner's dependents against potentially catastrophic financial losses in the event of the death of the insured. As such, over half of all life insurance policies are purchased by individuals between the ages of 25 and 44 (LIMRA International's Buyer Study, various years). Other possible reasons for owning life insurance include opportunities for tax-advantaged savings or the provision of liquidity to estates subject to U.S. estate tax laws (Brown 1999; Holtz-Eakin, Phillips, and Rosen 1999).

There are many types of policies available. One distinction is made between individual, group, and credit life insurance. Individual life insurance policies are sold directly to individuals and are underwritten separately for each purchaser. Group policies are often provided by employers or unions and are underwritten for the group as a whole. Credit life insurance is designed to guarantee payment of a mortgage or other loan in the event of the insured's death. Of 59 million policies sold in 1999, 20 percent were individual life policies, but these policies account for 56 percent of the face value of coverage. The reason is that group and, especially, credit life policies tend to be small.

Within individual life insurance policies, there are two basic types: term and whole. Of the total amount of individual coverage purchased in 1999, 57.6 percent was for term insurance. Term life policies provide life insurance coverage for a specified period of time, such as one year or five years. When the term period ends, these policies provide no additional benefit to the insured. As such, term life policies are pure insurance over the period of the contract and are relatively homogeneous. Whole life policies are not term dependent (hence they are also known as permanent life or cash value policies). Instead they provide insurance over the "whole of one's life" (Graves 1994). In addition, these policies typically include a savings component that builds up a

cash value over time. Policy owners can borrow against this cash value, and the accumulation in the cash value account is generally tax-deferred. If at any point the individual cancels the policy, the owner is entitled to receive the full cash value, minus a surrender fee and any outstanding policy loans. For these reasons, whole life policies have higher premiums per thousand dollars of coverage than term policies. The fact that they are less homogeneous also means that the traditional issues of search theory are perhaps not as relevant for them.

B. *Life Insurance and the Internet*

By 1996, there were a number of insurance-oriented web sites that provided consumers with access to on-line quotes for insurance products. The customer would, essentially, answer the medical questionnaire on-line including age, gender, personal medical history, and the like and then enter the amount of coverage sought. The sites would then report numerous companies that would offer such a policy and would give a price quote from each. A simple example for a 30-year-old nonsmoker with no medical problems searching at www.quickquote.com is shown in figure 1.

Importantly, in almost all cases, the individual does not buy the product on-line directly from these sites. Indeed, most industry analyses have emphasized the conservative nature of the off-line insurance business and sellers' reluctance to conduct commerce on-line (see Temkin et al. 1998; Klauber 2000). With these search services, a connection to the off-line seller remains. Consumers must still take a blood test, for example, to qualify for various policies. The sites are almost strictly a comparison/referral device. But with the creation of these sites, the costs of comparing prices for a given set of risk factors, age, gender, and other characteristics became extremely low. Users can get dozens of quotes in a matter of seconds that would previously have taken a great deal of searching. These Internet search sites essentially provided an information source between the consumer and the life insurance company that was formerly available only to brokers (see Garven 2000).[4] We do not have information on the total number of users of these types of sites in our sample, but the data in Forrester's Technographics 2000 database and in Clemmer et al. (2000) indicate that by 1999, more than 5 million households had researched life insurance on-line.

Two important aspects of the Internet insurance sites help us to dis-

[4] Some of the major sites include Insweb, Insurance.com, Quickquote, Quotesmith, Insurerate, Answerfinancial.com, and many others. They are reviewed periodically by www.gomez.com. Quotesmith began as a phone-in comparison service (with somewhat limited appeal). In late 1995, it became the first to provide quotes on-line. Insweb was the first large site, and it began at the end of 1995. Most of the others began soon after.

Quotes...

Quotes for Male Non-Smoker Preferred 5 Year Term Life Insurance (MNP5)

Annual Premiums	5 Year Total	Company	Plan
○ $450.00	$2250.00	SECURITY CONNECTICUT LIFE INSURANCE COMPANY	TERMSMART-5
○ $530.00	$2650.00	ZURICH KEMPER LIFE	SUPER-T 5
○ $565.00	$2825.00	CNA LIFE INSURANCE COMPANY	405G
○ $590.00	$2950.00	WESTERN-SOUTHERN LIFE ASSURANCE COMPANY	E-TERM 5 PREFERRED PLUS
○ $640.00	$3200.00	UNITED OF OMAHA	PRIORITY VALUE TERM 5
○ $700.00	$3500.00	WESTERN-SOUTHERN LIFE ASSURANCE COMPANY	E-TERM 5
○ $750.00	$3750.00	THE MIDLAND LIFE INSURANCE CO.	ALTERMATIVE FG

Compare These Plans | Lowest Possible Quote | All Possible Quotes

Apply for Insurance

FIG. 1.—Sample page from Quickquote.com

tinguish the Internet/search cost hypothesis from alternative explanations of the price declines. First, the comparison sites have focused almost exclusively on term life insurance. This is the more commodity-like product and is therefore easy to compare. Whole life policies are more differentiated, and the sites did not provide comparison quotes for them. Second, the comparison web sites mainly did not start until 1996, whereas Internet usage had already increased significantly for many groups prior to that time. Growth in Internet usage before 1996

should not affect competition in term life insurance, only growth after the comparison sites came on-line.

III. Literature and Theory: Search Costs, Pricing, and the Internet

Our approach is to think of the Internet as reducing search costs and analyze its impact empirically. In that sense, it is in the spirit of the empirical search models mentioned above. Since the original work on search theory of Stigler (1961), there have been numerous models analyzing the impact of search costs and differential information on the distribution of market prices.[5] The exposition most relevant for our empirical work is that of Stahl (1989).

The Stahl model begins with most customers (specifically a share, $1 - \mu$, of them) paying a search cost for each price quote/store visit they receive. A small fraction of the customers (μ), however, have no search cost. Customers search sequentially, and in the Nash equilibrium, the stores choose prices from a distribution rather than have a pure strategy. The customers with positive search costs have an endogenously determined reservation price and stop searching whenever they find a price below that. The customers with zero search costs get price quotes from all the firms and buy from the lowest-priced one.

While, in reality, Internet comparison sites do not reduce search costs to zero, they do dramatically reduce the costs associated with getting price quotes. We shall therefore interpret the share using the Internet to research insurance as a proxy for μ, the share of the population that is fully informed, and examine what happens to price and dispersion as μ changes.[6]

There are three basic results stated in the Stahl model that have direct predictions for our empirical work (and, in essence, summarize key findings of the search literature).

First, and most simply, when there are asymmetric search costs across customers (i.e., some have zero search costs and others do not), firms will draw equilibrium prices randomly from an equilibrium distribution rather than all of them charging a single market price. We should expect to see price dispersion in equilibrium.

Second, as the share of customers with complete information (μ) increases, the price distribution shifts downward monotonically. In other

[5] See the work of Diamond (1971), Salop and Stiglitz (1977), Varian (1980), Burdett and Judd (1983), Carlson and McAfee (1983), or Stahl (1989). Recent work by Baye and Morgan (2001) and Hall and Rust (2001) provides important analyses of the potential role of Internet shopbots and middlemen.

[6] The important issue in the Stahl model is not that search costs actually be zero but rather that there be a group of customers who get price quotes from all the firms and are fully informed. This is not a bad approximation of what the comparison sites enable.

words, as the share of consumers with no search costs increases, average prices should fall.

Third, when $\mu = 0$, the price distribution is degenerate at the monopoly price. When $\mu = 1$, the distribution is degenerate at the competitive price. As μ increases from zero to one, the distribution moves continuously from one to the other. This is important because it implies that the relationship between search costs and price dispersion is *not* monotonic. Increasing the share with no search costs will *increase* price dispersion for small enough starting levels of asymmetric information across consumers. If μ is large enough to begin with, then increasing μ will reduce dispersion. The case with large initial μ is the one assumed by most empirical work on search. Since we shall be observing the initial entry of the insurance web sites, however, this may correspond with a starting μ close to zero. As the share using the Internet to compare prices on-line rises from zero, price dispersion should first rise and then (eventually) fall.[7]

A final issue, one not raised in the Stahl model but important for our empirical work, is the distinction between the offer price distribution and the transaction price distribution. Most of the theoretical literature concerns what happens to the distribution of offered prices. Our data will pertain to transactions rather than offerings. This distinction does not undermine our analysis but instead raises two potential ways in which prices can shift downward. Customers facing the same offer distribution can shift to lower-priced firms or firms can lower their offer prices (or both). The distinction is important because in one model, only the Internet customers will benefit, whereas in the other, even non–Internet users see price declines. As we describe below, our data will not allow us to directly disentangle the relative importance of these two effects, but the magnitudes will suggest that prices must be falling for more than just the Internet users themselves.[8]

[7] Sorensen (1999) has suggested, in a slightly different model, that the maximum dispersion of list prices occurs at very low levels of search costs and that for plausible ranges, reducing search costs reduces dispersion. In our data, we shall have transaction prices (i.e., quantity weighted) rather than list prices, which is likely to influence this result. We simulated the Stahl model using a linear demand curve and the basic cost structure given in Stahl's numerical example and computed the expected difference between the highest and lowest prices and the variance. We found that the dispersion was increasing with μ up to about .1 in this case.

[8] This is similar to asking why the firms do not just price-discriminate using the Internet by offering different prices for the same good on-line. This did not take place during the period of our sample. Industry participants believe that the reason is that firms must file their rates with state regulatory agencies and justify them on some grounds other than price sensitivity as well as because of the administrative costs of filing new policy types and the inherent channel conflicts that arise from the agents in the normal sales channel (personal communication with Robert C. Bland, chairman and founder of Quotesmith.com [2001]). Very recently, firms have begun experimenting with products sold strictly over the Internet, so price discrimination may become more relevant in the future.

IV. Data on Prices of Insurance and Internet Usage

A. *Data on Life Insurance*

LIMRA International conducts annual surveys of purchases of individual life insurance contracts in the United States. Each year, LIMRA uses a sample of approximately 30,000 policies issued by an average of 46 participating companies per year, collecting detailed information on the policy characteristics and prices as well as some demographic information on the insured individuals including age, state of residence, occupation, and income. For purposes of this study, we have combined data from six Buyer Studies covering the period 1992–97. The LIMRA data are the most comprehensive in the industry and are widely used for empirical work on life insurance.[9] They do not include company identifiers, however, so we cannot include firm dummies.

We shall be primarily concerned with the prices of term life policies and how they respond over time as their buyers begin using the Internet. To keep the product as homogeneous as possible for our pricing regressions, we restrict the sample to level term policies owned by the premium payer, insuring the life of only one person, for people aged 20–75, and without any other riders (e.g., consumer price index [CPI] cost of living adjustment). We also look only at terms of five years or less (about 70 percent of term insurance). We do this because during the late 1990s, state insurance regulators were discussing changes to reserve requirements for policies with long-term premium guarantees (now known as "regulation triple X"). This regulatory action may have affected prices of longer-term policies in a way that is difficult to adequately control for.[10]

Several individuals lack some of the requisite demographic or policy information, so we must drop them.[11] Even with these various restrictions, we still have almost 11,000 person-year observations and about one-third of the total term life insurance in the sample. Summary statistics for the insurance variables are listed in table 1. In the raw data, without any controls, the median age-state-year group had a standard deviation of log prices of about .55.

[9] More details on these data can be found in LIMRA (1999).

[10] Note that we did find similar results when longer-term policies were included in the sample.

[11] The two problem variables in terms of missing values were the occupation code (and, of course, the occupation dummies), which was missing for about 38 percent of the observations, and the marital status variable, which was missing for another 10 percent (nonoverlapping). We repeated the results leaving out occupation and marital status to increase the sample and found the same answers.

TABLE 1
SUMMARY STATISTICS, 1992–97

Type	Term
Premium/($1,000 of face)	3.62
	(4.91)
Real amount of policy (in thousands of 1990 dollars)	132.97
	(136.41)
Length of policy	2.27
	(1.86)
Nonsmoker	.77
	(.42)
Male	.67
	(.47)
Policy is rated	.08
	(.27)
R^{length}	.88
	(.10)
Participating policy	.88
	(.32)
% on-line	.17
	(.14)
Observations	10,812

SOURCE.—Authors' calculations using data from LIMRA International and Forrester.
NOTE.—Standard errors are in parentheses.

B. *Data on Internet Use*

It would have been easiest to estimate the impact of the Internet on prices if the LIMRA data had asked the individuals directly whether they had checked insurance sites on-line. Lacking such information, we instead create a measure of the probability of Internet usage for each individual in each year on the basis of the person's observable characteristics. To compute this measure, we turn to the Technographics 1999 survey of Forrester, the leading market research company on the information economy.

Forrester conducted a nationally representative survey of almost 100,000 people in late 1998 that gathered information on their computer ownership, Internet use, on-line buying behavior, and the like, as well as demographic and geographic information on the individuals.[12] One of the questions Forrester asks of those with on-line access is how long they have been on-line. Another is whether they have ever researched various products on-line, and one of the products they report on is insurance. Importantly for our purposes, the Forrester survey collects age, state, occupation, and income information that we can match to the LIMRA data. Occupation and income are harder to match than

[12] More details on the Forrester data can be found in Bernhoff, Morrisette, and Clemmer (1998) and Goolsbee and Klenow (in press).

age and state because the occupation codes do not match precisely across the two data sets and because the Forrester income pertains to the family whereas the LIMRA income pertains to the individual.

We compute for each age-state-year, age-occupation-year, occupation-state-year, and age-income-year the share of people in that group that had on-line access in December of that year.[13] The retrospective data on on-line usage go back to 1994. For 1993 and 1992, we scale each group's 1994 on-line usage by overall rate of growth of domain names as tabulated by the Internet Software Consortium's Internet Domain Survey: Number of Internet Hosts (http://www.isc.org/ds/host-count-history.html; accessed May 25, 2000). In the few regressions in which we use the early information, this adjustment had little impact on the results since on-line usage rates were extremely low in those two early years. The overall share of people with on-line access rose from 2.6 percent in 1992 to 5.1 percent in 1993, 8.9 percent in 1994, 15.7 percent in 1995, 26.7 percent in 1996, and 38.8 percent in 1997. Of key importance for our regressions is the considerable variation in both the levels and growth patterns of on-line usage between groups. Not all groups grew at the same rate over time.

Because we are concerned with the use of the Internet for comparing insurance prices, including a measure of Internet usage in a price regression is equivalent to assuming that the use of insurance sites is proportional to use of the Internet (i.e., some constant fraction of Internet users go to insurance shopping sites). Since the insurance sites largely did not begin until 1996, our basic measure of Internet use for the group will be zero until 1996 and then equal to the share of people on-line after that. We shall also show results that compare the impact of Internet usage in the earlier years on insurance prices to check whether rising Internet use is spuriously correlated with prices.

C. *Specification*

Over the last half of the 1990s, life insurance consumers witnessed a large decline in the price of term life insurance. When no controls are taken account of, the average annual premium paid per $1,000 for a renewable one-year term policy was $3.20 in 1993 and by 1997 had fallen more than 20 percent to $2.50.

The actuarially fair pricing of a one-year term policy that pays out a face value of F on the last day of the year (with no other costs assumed) will depend on the probability of dying during the period, q_a for an

[13] The reason we do not compute the age-occupation-state-income-year groups is that we do not have enough data to estimate such a model. To be fully general would mean including dummies for more than 500,000 potential cells, and we have about 11,000 observations.

individual of type a, and on the interest rate r according to $P = q_a F/(1 + r)$. Higher expected mortality rates (high q) and lower interest rates (r) raise the marginal cost and thus the premium. This approach is consistent with the typical regulatory approach of setting reserve requirements strictly on the basis of interest rates and mortality (Graves 1994). Extending this formula to multiple-year policies is straightforward.

Our regressions will attempt to explain the price paid for term policies. The dependent variable is the log of the annual premium per $1,000 of face value of insurance. The Internet variables described above are our main variables of interest. As for our controls, we do not have a direct calculation of the survival probability for the individual, so we include standard variables to proxy for it including age dummies, a nonsmoking dummy, a gender dummy, marital status dummies, and a dummy for whether the policy is "rated," meaning that the individual belongs to a special risk class because of some personal behavior such as being an amateur pilot. We also include state dummies and occupation dummies to account for differences in health or demographic characteristics across groups that are correlated with life expectancy as well as dummies for whether the policy was purchased from an own agent and whether it was a participating policy.[14]

In addition to these variables, we want to allow for economies or diseconomies of scale in the costs of policies of different sizes and lengths, as discussed in Cawley and Philipson (1999). Therefore, we include policy length dummies and several terms for the value of the policy in real dollars (these are the log of the real amount, the real amount, and the real amount squared as well as dummies equal to one if the reported value was censored at the maximum value in the year). In practice, though significant, these nonlinearities had little effect on our results. We tried various functional forms and found the same answers. We use the monthly CPI as the deflator and the inverse of one plus the Baa bond rate for the interest rate term (raised to the length of the policy for term lengths more than one year). We also include year dummies. The coefficient on the year dummies gives us a price index in log terms for the cost of identical term life insurance over the period.

[14] Participating policies are typically issued by mutual life insurers. They allow the policy owner to participate in the company's surplus via distribution of a policy owner dividend.

TABLE 2
BASIC SPECIFICATION

Type	Term (1)	Whole (2)
D93	.0609	−.0597
	(.0133)	(.0134)
D94	−.0146	−.0533
	(.0124)	(.0111)
D95	−.0677	−.0671
	(.0128)	(.0119)
D96	−.1874	−.0111
	(.0133)	(.0148)
D97	−.2702	−.0031
	(.0131)	(.0145)
Nonsmoker	−.4596	−.1573
	(.0098)	(.0079)
Male	.1867	.1035
	(.0095)	(.0095)
Rated	.6140	.3365
	(.0201)	(.0141)
R^{length}	1.4532	.8046
	(.3612)	(.0494)
Participating	−.0001	−.0312
	(.0103)	(.0092)
Own agent	.0955	.4629
	(.0249)	(.0173)
Others	Amount, length, marital status	Amount, interest rate, marital status
Dummies	Age, state, occupation	Age, state, occupation
R^2	.837	.764
Observations	10,812	29,917

NOTE.—The dependent variable is the log of the annual premium per $1,000 of face value of insurance. Variables are defined in the text. In addition to the coefficients listed, both regressions include the log of the real face value, the real face value, and the real face value squared; dummies if the face amount was censored at the maximum reported value; and dummies for marital status as well as age, state, and occupation, as indicated at the bottom of the column. Col. 1 concerns term life policies, and the regression also includes dummies for policy length. Col. 2 concerns whole life policies, and the regression also includes policy length as defined in the text and the interest rate term itself as well as the interest rate term to the power of the policy length. Standard errors are in parentheses.

V. Basic Results

A. *An Overview of Price Trends for Term and Whole Life Policies*

The results from this regression are listed in column 1 of table 2. The explanatory power of the regression is high with an R^2 of .837. These variables explain a large fraction of the variance in policy prices. The coefficients on the explanatory variables are fully in line with expectations. Policies for men cost about 20 percent more than identical policies for women, and those for smokers cost 45 percent more than for nonsmokers. When interest rates rise (lowering the inverse interest rate term), prices are reduced. Most important, the results show a dramatic decline in prices of term life insurance, especially toward the end of the sample. Relative to real prices in 1992, prices for identical policies

TABLE 3
RESULTS BY CATEGORY: TERM PRICES

	State		Occupation		Age	
Sample	Calif., Wash., Va. (1)	Ala., La., Ky., Ark. (2)	High Skill (3)	Low Skill (4)	<30 (5)	>45 (6)
D93	.0801	.1580	.0439	.0697	.0857	.1239
	(.0395)	(.0555)	(.0229)	(.0372)	(.0264)	(.0321)
D94	.0262	−.0399	−.0221	−.0195	−.0426	.0379
	(.0377)	(.0454)	(.0215)	(.0359)	(.0257)	(.0325)
D95	−.0605	−.0788	−.1029	−.0171	−.1127	−.0095
	(.0354)	(.0589)	(.0220)	(.0331)	(.0290)	(.0338)
D96	−.1932	−.0920	−.2030	−.1484	−.2530	−.0996
	(.0377)	(.0503)	(.0227)	(.0384)	(.0276)	(.0328)
D97	−.3203	−.1254	−.3311	−.2293	−.3496	−.1411
	(.0411)	(.0526)	(.0218)	(.0413)	(.0260)	(.0344)
R^2	.839	.828	.811	.866	.741	.820
Observations	1,451	623	3,347	1,297	2,248	205

NOTE.—The dependent variable is the log of the annual premium per $1,000 of face value of insurance. All the regressions concern term life policies. The sample is restricted to the group listed at the top of the column. Variables are defined in the text. In addition to the coefficients listed, all the regressions include 20 variables and age, state, and occupation dummies. They are the same as those in table 1. Standard errors are in parentheses.

were about 1 percent lower in 1994 but almost 19 percent lower in 1996 and 27 percent lower in 1997.

Thus prices seemed to fall most at the time the Internet insurance comparison sites came on-line. Whole life prices make an interesting comparison since the insurance sites did not cover such policies. Column 2 of the table repeats the specification of column 1 for the price of whole rather than term policies.[15] Interestingly, at the start of the sample, the price changes for whole and term policies were very similar: term life prices in 1995 were 6.8 percent below 1992 levels, and whole life prices were 6.7 percent below. In 1996 and 1997, however, prices dropped dramatically for term policies, whereas whole life policies remained constant or even rose slightly.

B. *Overview of Price Trends for Term Life across Selected Demographic Groups*

Next, in table 3, we repeat the term life hedonic regressions but compare price changes among groups for which Internet usage grew at different rates to get suggestive evidence as to whether there is any apparent role for the Internet (all groups had close to zero Internet usage in 1992).

[15] Since whole life policies do not have a limited duration, there is no way to limit the length of the policies to five years or less. We estimate the policy length as being 80 minus age for women and 72 minus age for men. Given the longer time frame of these policies, we use the five-year bond rate rather than the one-year rate and include the interest rate on its own in the regressions, though this did not matter for the results.

Column 1 gives the results for policies in California, Washington, and Virginia—the states with the fastest Internet penetration by the end of the sample (more than 40 percent in 1997). Column 2 looks at policies in Alabama, Louisiana, Kentucky, and Arkansas—the states with the slowest penetration by the end of the sample (about 25 percent in 1997). The results show that prices for identical policies in states with rapid Internet adoption fell significantly faster at the end of the sample (1997 prices were 32 percent below 1992 levels) than they did in states with slow Internet adoption (1997 prices were about 13 percent below 1992 levels).

The same thing is true in columns 3 and 4, which compare policies for people in high-skill occupation codes (professionals, students, and military), which had average Internet use of about 49 percent by 1997, to policies for people in low-skill occupation codes (operatives, service workers, and farmers), which had Internet usage of 22 percent by 1997. In columns 5 and 6, we see that the price declines were also significantly larger for people under age 30 (Internet use of 46 percent by 1997) than for people over age 45 (Internet use of 34 percent by 1997).

These regressions suggest a correlation between Internet use and price declines. In our attempt to attach a causal relationship between the two, however, we need more detailed data on changes in Internet usage and we need to confront potential alternative explanations. We address these issues in the sections below.

C. *Basic Results*

In table 4, we add the probability of Internet usage (calculated from the Forrester data described above) to the price regressions. We compute the Internet usage in each year share for age-state, age-occupation groups, age-income, and occupation-state groups. The standard errors are corrected for the fact that the Internet usage variable varies only by group-year and not by individual-year. In every case, the coefficients are negative and significant, suggesting that prices for identical term life policies for people in a given group fell more during those periods in which the group had faster adoption of the Internet.

Note that because there are age, occupation, state, and year dummies in the regression, these results cannot be explained by level differences in price or life expectancy across groups or time periods. People aged 25–30 may have lower life insurance prices than people aged 45–50 because of health differences, life style choices, and many other reasons. These reasons may be correlated with Internet usage, but this will not appear as a positive coefficient on Internet usage in our regression. It will be absorbed in the age dummies.

The magnitudes of the coefficients indicate that increasing the share

TABLE 4
LOG REAL PRICE AS A FUNCTION OF INTERNET USAGE

Type	Age × State (1)	Age × Occupation (2)	Age × Income (3)	Occupation × State (4)
% use Internet	−.5109	−.2269	−.3454	−.1819
	(.1189)	(.0955)	(.1078)	(.0860)
D93	.0606	.0600	.0118	.0605
	(.0143)	(.0110)	(.0170)	(.0133)
D94	−.0100	−.0142	−.0301	−.0142
	(.0149)	(.0160)	(.0135)	(.0121)
D95	−.0681	−.0669	−.0394	−.0672
	(.0130)	(.0146)	(.0151)	(.0129)
D96	−.0515	−.1240	−.0955	−.1409
	(.0333)	(.0289)	(.0341)	(.0269)
D97	−.0663	−.1757	−.1401	−.2005
	(.0499)	(.0403)	(.0454)	(.0379)
R^2	.838	.837	.829	.838
Observations	10,812	10,812	8,676	10,806

NOTE.—The dependent variable is the log of the premium per $1,000 of face value of insurance. All the regressions concern term life policies. The variable % use Internet is the share of the group that had Internet access in the given year. Variables are defined in the text. In addition to the coefficients listed, all the regressions include 20 variables and age, state, and occupation dummies. They are the same as those in table 1. Standard errors are in parentheses.

of a demographic group that uses the Internet by 10 percentage points lowers prices for that group by about 1.5–4.5 percent, depending on the specification. Because of the potential measurement error in the occupation and income variables mentioned above, we shall concentrate our results below on the age-state variation, but the findings were very similar in almost every case, no matter which one we used.

In addition, the Internet usage variable seems to explain a large part of the total decline in prices over this period. In the baseline results without Internet use, as previously listed in column 1 of table 2, prices fell about 27 percent over the sample. In these specifications, once we control for the role of Internet usage, the year dummies are significantly less important. The total decline is only 6 percent and is not significant in the age-state regression, meaning that the growth in Internet usage can explain about three-quarters of the total declines in term life prices. Even in the regressions in which the Internet variable is measured with error (i.e., include occupation or income), the Internet still appears to explain between one-quarter and one-half of the total decline.

As described above, the implicit assumption in these results is that a constant fraction of all Internet users check insurance sites on-line, and this fraction does not vary across groups. Even with that assumption, unless the fraction is literally one, the coefficient will be modified by some unknown scaling factor. To loosen these restrictions, we turn to the question in the Forrester data about whether the individual with on-line access has ever researched insurance on-line. We compute the

share of each group that has done so (as of 1998) and multiply it by the share with on-line access in each year. This gives us a measure of the share of the group that both has on-line access and has researched insurance on-line. This puts a reasonable scale factor on the results and simultaneously allows for different groups to have differing likelihoods of researching insurance on-line.

One problem with this measure is that since only 10 percent of on-line users report researching insurance and the mean share of Internet users is only about 27 percent in 1996 and 38 percent in 1997, there are many smaller demographic groups that suffer from small-sample problems, so the composite measure may tend to add noise to the Internet variable (i.e., the true share doing on-line research is roughly 2.7 percent in 1996, but this will tend to show up as zero in the data for small demographic groups). This measurement error will tend to bias the coefficient toward zero.

The results from using this insurance measure with age-state-year as the explanatory variable are presented in column 1 of table 5. Despite the added noise from the small-sample problem, the coefficient is still negative and significant. Raising the share of the group using the Internet to research insurance on-line by 1 percent lowers prices by about 2.5 percent.[16] We shall use this insurance research variable in the remaining results (though, as in this case, we found the same general results using the straight Internet usage variable in all the specifications).

Given the observed impact of the Internet on term life prices, we can make a back of the envelope calculation as to the gain in consumer surplus from the price declines generated by growth of the on-line comparison sites. We do this by multiplying the change in the price generated from the increase in Internet usage over the period (8–15 percent in our specifications) by the total amount of term life that was sold in 1995 (the year prior to the introduction of these sites).

The total annualized new premiums of all individual life products sold in 1995 was $9.6 billion. According to LIMRA (2000), 15 percent of these premiums were for term policies, for a total of $1.44 billion of *new* term business. Our results indicated that the price declines resulting from the increase in on-line usage from 1995 to 1997 generated an annual increase in consumer surplus of about $115–$215 million, quite

[16] A coefficient that exceeds one in absolute value, as it does here, is consistent with a search externality in the sense of Salop and Stiglitz (1977); i.e., when a large share of the members of a group begin using the Internet to research insurance, this can reduce prices for everyone in the group, not just the Internet users. Because our data give the share of the entire group that researches insurance on-line rather than the share of the potential life insurance buyers in each group, however, we cannot be sure about the absolute magnitude of the coefficient, so we shall not pursue the externality point further. The magnitude, however, suggests that more must be going on than that informed customers are buying from low-priced firms.

TABLE 5
FURTHER CONTROLS

	Research (1)	Nonlinear (2)	Mortality (3)	Whole Life (4)	Interactions (5)	Early Years (6)	Pr(Rider) (7)
% research	−2.5360	−3.9660	−2.3760	.3876	−3.1570	−2.4360	.2230
	(.3701)	(1.0120)	(.3689)	(.3871)	(.4831)	(.3922)	(.4168)
(% research)^2		14.2280					
		(8.5030)					
% research (use from 1992–95)						.6110	
						(.7014)	
ln(mortality)			.1072				
			(.0392)				
D93	.0600	.0604	.0571	−.0597	.0409	.0587	−.0508
	(.0161)	(.0143)	(.0144)	(.0142)	(.0146)	(.0148)	(.0040)
D94	−.0143	−.0143	−.0186	−.0533	−.0122	−.0188	−.0318
	(.0128)	(.0128)	(.0129)	(.0117)	(.0417)	(.0136)	(.0028)
D95	−.0677	−.0677	−.0721	−.0671	−.0646	−.0765	−.0272
	(.0129)	(.0129)	(.0131)	(.0126)	(.0152)	(.0171)	(.0060)
D96	−.1164	−.0909	−.1156	−.0215	−.1007	−.1176	−.0690
	(.0161)	(.0228)	(.0159)	(.0181)	(.0200)	(.0162)	(.0120)
D97	−.1625	−.1351	−.1548	−.0177	−.1415	−.1650	−.0891
	(.0214)	(.0292)	(.0216)	(.0212)	(.0263)	(.0216)	(.0170)
R^2	.838	.839	.839	.764	.885	.838	.087
Observations	10,812	10,812	10,812	29,917	10,812	10,812	11,499

NOTE.—The dependent variable is the log of the annual premium per $1,000 of face value of insurance in all the columns except col. 7. Col. 4 concerns whole life policies, whereas all other columns concern term life policies. The variables are defined in the text. The dependent variable in col. 7 is the binary variable of whether the policy has a rider. Each regression also includes 20 variables and age, state, and occupation dummies. Standard errors are in parentheses. The % research variable in the years before the comparison sites existed is the on-line usage in the year times the share of the group having researched insurance on-line in 1998. The results were quite similar when just % on-line was used instead.

large for a service used by only a small number of people. This figure may understate the magnitude of the impact of the Internet because new term policies are dwarfed by renewals of term policies; renewals might also decline since policyholders have the choice of replacing an expensive existing policy with a low-priced alternative. Including renewals is probably the upper bound, though, since we would expect renewals to be much less price-sensitive than new policies because of switching costs. There were roughly $7 billion of term life renewal premiums in 1995 (American Council of Life Insurance 2000), so if the Internet caused a similar 8–15 percent reduction in these prices, that would put an upper bound of an additional $560 million to $1 billion in consumer surplus. It is important to emphasize again, though, that since we do not know the identities of the companies in our sample, we cannot refute the hypothesis that the Internet comparison sites caused people to choose policies from companies with lower quality along some dimension that we do not measure or in policy attributes that we do not observe. If true, the change in price would not represent a pure increase in consumer surplus.[17]

In column 2 we consider the possibility that the impact of the Internet is nonlinear. The initial introduction of the Internet may matter a lot for prices, but once usage is widespread, the markets may be competitive. When we include a square term in the regressions, there is some evidence of nonlinearity, but it is only borderline significant. For most of the range in the sample, the marginal effect of increasing the share of the group researching insurance on-line is fairly constant, so we shall just include the linear term in the results that follow. At the ninetieth percentile in the data (about 6 percent of the group having researched insurance on-line), for example, the marginal effect is still 85 percent of the marginal effect at zero Internet use. The projected declines in the marginal effect are mainly outside the observed values in the data. The impact of having a greater share of users on-line would be insignificantly different from zero when about 19 percent of the group researched insurance on-line (and the point estimate would be zero at 27 percent).

These basic specifications point to a correlation between the growth in Internet insurance site usage and declines in prices for term life insurance. In the next section we consider the viability of some alternative explanations for these findings.

[17] This caveat may not be as relevant in our sample since our evidence is based on short-term policies and the primary measure of quality here—the likelihood that the company will pay on the death of the insurance holder—is partially insured by state insurance guarantee funds.

VI. Alternative Explanations

A. *Changes in Mortality*

The most straightforward alternative explanation of the results is that changes in Internet use by a group are spuriously correlated with changes in mortality rates that directly reduce the cost of life insurance for the group. As a general matter, mortality improvements are important for insurance prices. Mortality has declined over most of the twentieth century, and, unsurprisingly, the price of term life insurance has, as well. Mortality improvement from 1992 to 1997, however, was gradual and cannot explain the sharp price declines witnessed at the end of the sample and especially pronounced for groups with rapidly growing Internet use.

As a specific test of the importance of mortality changes, in column 3 of table 5 we compute the log mortality rate for each age-state-year using population data from the Bureau of the Census and the number of deaths from the National Center for Health Statistics. We also tried including lags and leads of the mortality rate, but the results were identical. Note that since we already include state, age, year, and occupation dummies, we are identifying the impact of changes in mortality relative to the group mean on the prices of insurance. The coefficient on log mortality is positive and significant on prices, as expected, but the coefficient on the Internet term is not significantly different from the one in the previous regression.[18]

Another piece of evidence against the view of spurious correlation with life expectancy is the evidence on whole life prices. Changes to life expectancy should influence both term and whole life policies. Since the comparison sites did not cover whole life policies, however, we do not predict any reduction in search costs in that arena, so the Internet should have no effect on prices. The results, presented in column 4, show that rising shares of the group using the Internet to research insurance are not associated with lower whole life prices at all. The coefficient is 0.388 (and not significant) compared to the significant term life coefficient of −2.5.

[18] An alternative mortality-based explanation is to argue that the sample of life insurance buyers changed in 1996, with less healthy individuals becoming less likely to buy insurance. To explain our results, however, would require that the selection effect be stronger for groups with higher Internet use. To test for this, we ran a sample selection probit on data from the 1992 and 1998 Survey of Consumer Finances and found no evidence that the probability of owning term life insurance changed differentially by age, income, education, or occupation groups.

B. *Unobservable Differences across Groups*

Our results account for age, occupation, and state fixed effects. If there are distinct differences in the life expectancies of various interactions of those variables in a way that is correlated with Internet usage, this could bias our results. To deal with this issue, in column 5, we add age-occupation-state interaction dummies. When we do this, the number of dummy variables relating to these factors rises from 68 to 2,933. Now rather than just younger people having, on average, prices different from those of older people, high-skill different from low-skill, and so forth, we allow young, high-skill people in California to have prices different from those of young, high-skill people in Nevada and all the other permutations. Once we do this, we are identifying the impact of the Internet exclusively from the changes across time within a given group—whether prices fall more for 30–35-year-old service workers in Florida in those years in which their probability of using the Internet rose more.[19]

The results still show the same effect of the growth of Internet usage and, if anything, are larger than before. The coefficient is −3.15 versus −2.53 previously.[20] Note that the increase in the R^2 is modest despite the increase in the number of dummies. It rises from about .84 to .89.

C. *Spurious Correlation of the Growth of Internet Usage with Other Factors*

Fundamentally, any alternative explanation of the results we have found must be based on the idea that the growth in Internet use for a group is correlated with some other unobserved factor that is reducing prices for that group.

One way to check this general hypothesis is to estimate the effect of Internet usage on insurance prices during the period in which there were no on-line insurance sites (i.e., 1992–95). During this early period, there is no reason for rising Internet usage to be correlated with lower insurance prices unless it is spuriously correlated with some other factor. In column 6 of table 5 we add a variable that is equal to the share of the age-state-year with Internet access for 1992–95 interacted with the share having researched insurance on-line and then zero in 1996 and 1997 (in addition to our standard measure, which is zero from 1992 to 1995 and then positive in 1996 and 1997). The results show that prices fell significantly with the rising use of the Internet during the period

[19] We do not include the full set of possible dummies by age-state-occupation-income because the remaining cell size for all but the largest groups would be extremely small.

[20] Again, the results were very similar when on-line insurance usage by age-occupation, age-income, etc. are used or pure on-line usage rather than on-line insurance research. We do not report them here to save space.

in which the insurance sites existed and with approximately the same magnitude as before, but rising Internet usage had no significant effect on prices before the sites existed (and the point estimates are positive).[21]

D. *Unobservable Differences in Policy Attributes*

As we discussed above, the rise of the Internet sites might lead people to buy their policies from less reputable companies by focusing competition mainly on price, which is easy to observe. Similarly, it might lead people to get lower-cost policies that are worse on some other dimensions that we cannot observe (and perhaps the buyers themselves do not realize) such as renewability terms, exclusions, and so on. Here, the alternative explanation would not be that the Internet had no effect but rather that it changed competition to being about price at the expense of other factors.

In general, there is little we can do to test these issues given the nature of our data. One thing we can do, however, is look at the probability that individuals purchase riders on their policies (amendments to the original contract such as a CPI rider that lets the insurance value go up with the rate of inflation). These riders are not easily searchable on the Internet sites but are worth something to consumers, so this might be one form of observable quality change. If the rise of the Internet appears to reduce the probability of getting a rider, this might suggest changes in policy characteristics that are correlated with prices.[22]

In column 7 we report the results of a linear probability model of the probability that the buyer's term policy has a rider as a function of all the same explanatory variables. We see that raising the probability of using the Internet to research insurance has no significant effect on the probability of having a rider, and the magnitude is extremely small. At least on this one observable dimension, it does not appear that the

[21] A different type of evidence might be to compare price declines internationally, where Internet usage has been much smaller than in the United States. Unfortunately, we have only scant data on prices for other countries. Illustrative industry data presented by Cooper (2000), however, do give price information for a few specific policy types in the United Kingdom over our 1995–97 period. The results seem to confirm our expectation of only modest price declines. Recall that in our regressions, term life prices for identical policies fell by about 20 percent over these two years and by more for heavy Internet users such as the young. In the United Kingdom, prices for £100,000 five-year term policies for 30-year-old nonsmoking men fell only 3.2 percent and for women actually rose by 3.8 percent. For £300,000 policies, prices fell 6 percent for women and 9 percent for men. Prices for 45-year-olds actually fell a bit faster, close to 10 percent. These declines both seem smaller and have different relative magnitudes across groups relative to those in the United States.

[22] In our previous price regressions, we restricted the sample to policies without riders so as to make the products homogeneous. This explains why the sample is larger here (about 10 percent of policies have riders). Repeating those regressions but including any type of rider as a dummy variable indicates that the presence of a rider increases the price of the policy by about 17 percent on average.

TABLE 6
PRICE DISPERSION

	Standard Deviation	
	(1)	(2)
% research	3.871	3.477
	(.971)	(.981)
(% research)^2	−68.503	−50.555
	(29.503)	(30.017)
(% research)^3	307.002	187.001
	(203.984)	(205.370)
Constant	.264	...
	(.005)	
Dummies	None	Age, state, occupation
R^2	.028	.086
Observations	1,248	1,391

NOTE.—The dependent variable is the standard deviation of residuals from the price regression in col. 1 of table 2. Standard errors are in parentheses.

Internet has led to a change in the composition of policy types, such as toward lower-quality policies. Of course, worse policies on other dimensions that we cannot measure could explain them.

VII. Price Dispersion and the Internet

The results confirm that, consistent with the theory of search, as comparison shopping has become easier, average prices have fallen significantly. Much of the existing empirical literature about the Internet (and about search theory, too) has examined whether price dispersion falls when search costs are lowered. We have noted, however, that the theory does not have a monotonic prediction for price dispersion, especially when the starting share of fully informed consumers is low, as it is here. Further, our data are transactions, as opposed to list price data, so they will be weighted by volume. This will tend to accentuate the nonmonotonicity of the relationship at low levels of Internet use.

Using our regression results, we can examine the amount of price dispersion within observable groups and correlate it to the share of people using the Internet to research insurance (our proxy for having no search costs). To do this, we take the residuals from the price specification in column 1 of table 2 and compute the standard deviation within the age-state group for each year. This is the amount of price dispersion within a group that cannot be explained by the observable characteristics of the people or the policy types. The standard deviation in the residuals for the median age-state-year group is about .26.

In column 1 of table 6, we regress these measures of price dispersion on the on-line insurance use measure by age-state-year as well as the square and the cube of the measure to allow for nonlinearity (though

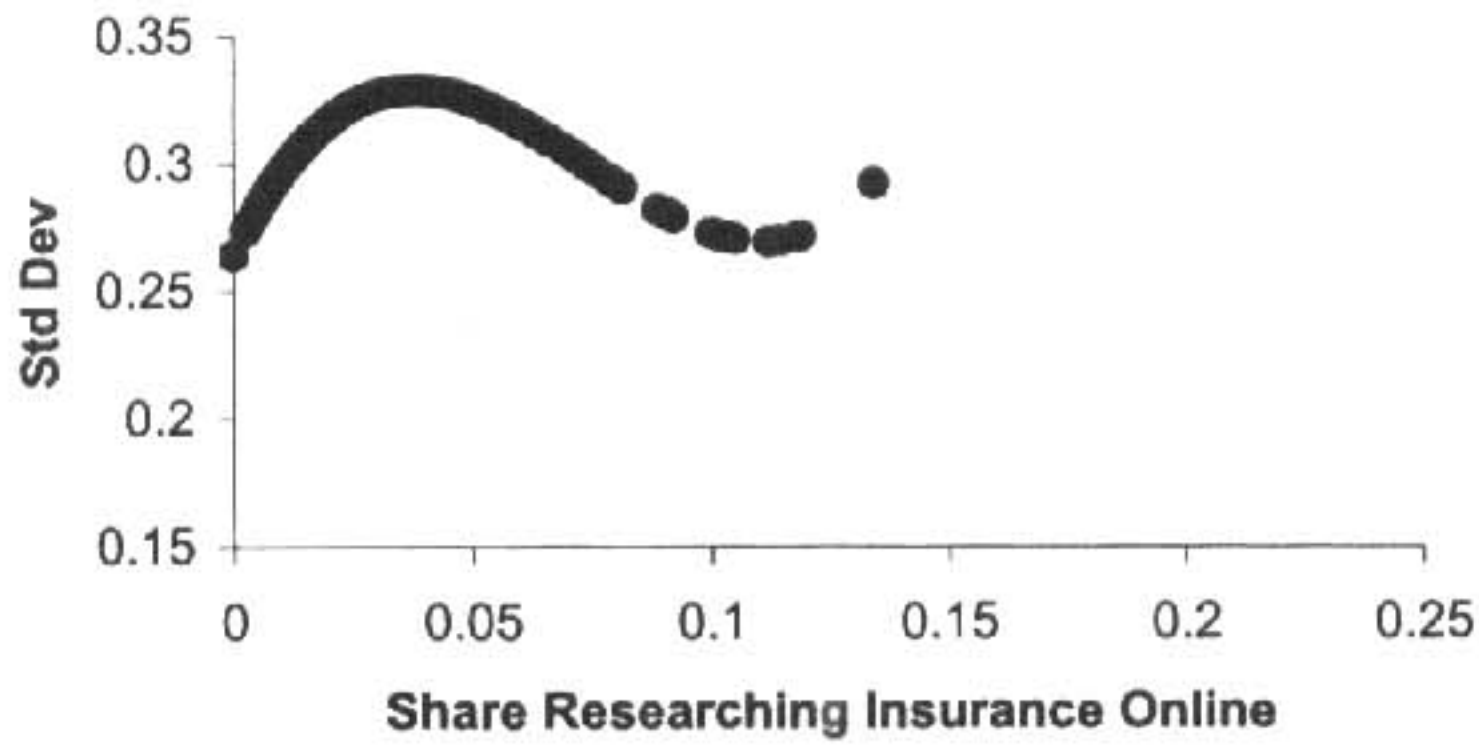

FIG. 2.—Predicted standard deviation of price residuals by age-state-year. Source: authors' calculations.

the standard errors are not corrected for the fact that the residuals are themselves estimated). In column 2, we also allow for age, state, and occupation dummies. In both regressions, the results show evidence of nonlinearity. We graph the predicted values as a function of the share in figure 2 for the specification without fixed effects to show the direction of the nonlinearity.[23]

The evidence indicates that price dispersion within groups is actually rising with the share of people researching insurance on-line for low shares and then falling with the share on-line once that share exceeds about 5 percent. Although this may seem counterintuitive, it is consistent with the theoretical predictions of the literature. When no one has access to full information, giving the information to a small number of people tends to increase the amount of price dispersion.

VIII. Conclusions and Future Directions

In this paper we have examined the market for term life insurance from 1992 to 1997 and documented that the growth of Internet price comparison sites appears to have made the market significantly more competitive. Controlling for policy characteristics and a variety of individual and group controls, we find that as the share of people in a group that use the Internet to research insurance on-line grows, the more their quality-adjusted prices fall. The data also show, consistent with the theory, that increasing the probability of using the Internet tends to raise

[23] The graph with the fixed effects looked the same. We also found the same nonlinear pattern using the interquartile range and the total range rather than the standard deviation. To save space, we do not report the results in table 6.

price dispersion initially and then reduce it as Internet usage continues to grow. The results seem somewhat robust: the growth of Internet use does not appear to reduce the price of whole life policies (which were not covered by the Internet insurance comparison sites), the growth of Internet use before 1996 (when insurance comparison sites did not exist) did not reduce prices, and the results are not affected by adding detailed controls for changes in group-specific mortality.

Overall growth of Internet usage can potentially explain a significant share of the large price declines of the 1990s. The rise of the Internet from 1995 to 1997 appears to have reduced term life prices by about 8–15 percent. Back of the envelope calculations suggest that, although seemingly a modest niche of Internet commerce, Internet comparison sites may have increased consumer surplus by $115–$215 million per year and perhaps by as much as $1 billion.

In this sense, our results show that, at least for some financial products, the ability of the Internet to reduce search costs can have a significant impact on market power. When it does so, it may lead to large consumer welfare gains, potentially at the expense of supplier profits. The implications for the market value of on-line and off-line companies could not be more important.

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