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Price and the health plan choices of retirees[☆]

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Abstract

This study analyzes health plan choices of retirees in an employer-sponsored health benefits program that resembles "premium support" models proposed for Medicare. In this program, out-of-pocket premiums depend on when an individual retired and his or her years of service as of that date. Since this price variation is exogenous to unobserved plan attributes and retiree characteristics, it is possible to obtain unbiased premium elasticity estimates. The results indicate a significantly negative effect of premiums. The implied elasticities are at the low end of the range found in previous studies on active employees.

JEL classification: I11; D12

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1. Introduction

In recent years, a number of proposals have been made to reorganize Medicare by placing a greater emphasis on market-based competition among health insurance plans.[1]

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[1] The most prominent examples are proposals that came out of the National Bipartisan Commission on The Future of Medicare. In 1999, the majority of the Commission's members supported a market-oriented reform

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In contrast to the current system of administrative pricing, payments to health plans in these “managed competition” or “premium support” models would be based on competitive bidding. The government’s contribution toward coverage would be set so as to expose beneficiaries to differences in premiums charged by competing plans. Individuals choosing a more costly plan would be required to pay the difference between the plan’s premium and the government contribution. Advocates of market-oriented reform argue that these financial incentives will cause beneficiaries to migrate to lower cost plans, which, in turn, will create a strong incentive for plans to control costs in order to compete on price.

To assess the potential impact of market-oriented reform proposals, it is necessary to understand the price sensitivity of Medicare beneficiaries. While a number of studies examine the effect of premiums on the choice among alternative health plans, this literature has largely focused on the behavior of non-elderly employees. An obvious concern is that the results from these studies may not generalize to the Medicare population. In particular, there are several reasons to suspect that older consumers may be more reluctant to switch health insurers and less sensitive to price than younger workers.

This study investigates the effect of premiums on the health insurance decisions of retirees in a setting that resembles prominent Medicare reform proposals. The analysis is based on 6 years of data from a large employer that offers its retirees several health insurance options. Like many employers, in the mid-1990s this company changed its policies concerning the financing of these insurance options. Whereas it had previously paid the full amount for each retiree’s insurance, the amount the firm now contributes toward a retiree’s coverage depends on when the person retired and her years of service at that point. Because otherwise similar retirees who either retired at different times or at the same time with different years of service face very different relative prices for the same menu of health plans, this policy creates an ideal natural experiment for estimating the effect of price on health plan choices. Changes over time in the employer contribution and plan premiums create additional price variation that is also plausibly exogenous to other health plan attributes and retiree characteristics that are likely to affect the demand for insurance.

Using these data, I estimate conditional logit models of the health plan choice decision. I find a negative and statistically significant effect of price on the probability a health plan is chosen by retirees. The results are robust to alternative assumptions concerning the definition of the choice set and to changes in the composition of the estimation sample.

Section 2 reviews the existing literature on the effect of price on health plan choice decisions. Sections 3 and 4 describe the data and econometric methods, respectively. Regression results are presented in the Section 5; Section 6 concludes.

proposal, though this proposal did not achieve the super-majority necessary for a formal recommendation. Subsequent legislative proposals include bills authored by the chairmen of the Commission, Senator John Breaux and Congressman Bill Thomas and more recent legislation proposed by Breaux and Senator Bill Frist (S. 357). These bills have important similarities to reform strategies advocated by academic health economists and other policy analysts. For example, see Enthoven (1988), Aaron and Reischauer (1995), Butler and Moffit (1995) and Dowd et al. (1996).

2. Previous research on health plan choice

A number of studies examine the effect of out-of-pocket premiums on health plan choice decisions, typically using data from a single employer and focusing on the decisions of employees rather than retirees. A key methodological issue for this literature is the source of variation in out-of-pocket premiums. When the variation comes entirely from differences across plans, correlation between premiums and other unmeasured plan characteristics becomes a possible source of bias. Barringer and Mitchell (1994) suggest that this type of omitted variable bias is a likely explanation for their counterintuitive finding of a positive effect of price in some models. Similarly, in their cross-sectional analysis of data from the Federal Employee Health Benefit Program, Wedig and Tai-Seale (2002) find that the estimated effect of premiums is very sensitive to the inclusion of other plan characteristics in the regression model.

The most convincing evidence on premium elasticities comes from studies that rely on within-plan differences in employer premium contributions for identification. Feldman et al. (1989) use data from 17 Twin Cities firms that offer an overlapping set of health plans to their employees. In this data set differences in employer contribution policies mean that different individuals face different prices for the same plans. Three other studies analyze the effect of price differences generated by changes in a single employer's contributions from one year to the next. Cutler and Reber (1998) examine the effects of a change in Harvard University's health plan contribution policy that changed out-of-pocket premiums for employees over a 2-year period. The employee response to the price change in the first year implies an "enrollee perspective" premium elasticity of −0.3. The response in the second year implies an elasticity of −0.6. Royalty and Solomon (1999) analyze 2 years of data from Stanford University. Their conditional logit results imply enrollee perspective elasticities in the same range; results using a fixed effect logit specification imply even stronger price effects. Strombom et al. (2002) estimate premium elasticities using data on employees of the University of California (UC). The range of their elasticity estimates is similar to those of the two earlier studies.

There are reasons to expect Medicare beneficiaries to be less sensitive to health insurance premiums than the active employees analyzed in these studies. Research on how the elderly perceive health insurance options suggests they place much more importance on factors such as quality of care, freedom of referral, and the burden of paperwork than on premiums (Harris, 1997; Harris and Keane, 1999). Thus, to the extent that Medicare beneficiaries view plan alternatives as being differentiated along these other dimensions, their choices will be less strongly influenced by differences in premiums. Experimental research finding that elderly consumers are more likely to treat health insurance premiums as a signal of quality (Uhrig and Short, 2002/2003) also points to a negative relationship between age and price sensitivity. An additional reason to expect older consumers to have a less elastic demand for health insurance comes from the fact that health insurance decisions are subject to persistence and "status quo bias" (Neipp and Zeckhauser, 1985; Samuelson and Zeckhauser, 1988). To the extent that older consumers face higher "switching costs," they will be less willing than younger individuals to change plans in response to a change in relative prices. In particular, Medicare beneficiaries with strong ties to certain providers may be reluctant to switch from one managed care plan to another if it means also having to change providers or to risk an interruption in treatment.

The results of several studies suggest the importance of switching costs in health insurance decisions. Strombom et al. (2002) estimate separate premium elasticities for 18 mutually exclusive groups of employees hypothesized to differ in terms of the cost of switching health plans. Consistent with the switching cost hypothesis, they find that price sensitivity declines with age and employment tenure and is lower for individuals with higher expected medical care utilization. Royalty and Solomon (1999) obtain qualitatively similar results when they test for differences in price sensitivity related to age, employment tenure and a different measure of health status. Wedig and Tai-Seale (2002) also find that new employees are substantially more price sensitive than incumbent employees, and Beaulieu (2002) finds some evidence that younger employees are more price sensitive than older ones.

While limited, direct evidence on the behavior of elderly consumers also suggests that they are less sensitive to health plan premiums than younger consumers. In an earlier study, I examine how retirees from the UC responded to changes in out-of-pocket premiums caused by a change in the University's premium contribution (Buchmueller, 2000). The results indicate that while changes in out-of-pocket premiums have a statistically significant effect on the decision to switch plans during open enrollment, UC retirees are less price sensitive than active employees who faced similar price changes. I also examine the effect of rising premium contributions on the percentage of retirees choosing fee-for-service Medigap coverage rather than an HMO. The structure of this part of the analysis resembles the work by Cutler and Reber (1998). The enrollee perspective elasticities for UC retirees range from −0.12 to −0.24, which is smaller than the range of Cutler and Reber's (1998) estimates.

Atherly et al. (2004) use data from the 1998 Medicare Current Beneficiary Survey (MCBS) to estimate premium elasticities for beneficiaries living in areas where more than one HMO was available through the Medicare + Choice program. Since they exclude from their analysis individuals with retiree health benefits through a former employer, their sample represents a different subset of the Medicare population than the UC retirees. Dowd et al. (2003) conduct a similar analysis using aggregate (county-level) data from 1999. The results from these two studies are quite similar to the results for UC retirees, with estimated premium elasticities of −0.14 (Atherly et al.) and −0.13 (Dowd et al.).

Each of these studies on the plan choices of Medicare beneficiaries has significant methodological limitations. Because in the UC data there was little price variation among HMOs, the choice that is modelled is the decision to join the single PPO option or to enroll in one of several HMOs. Consequently, the results may not apply well to the situation envisioned by Medicare reform proponents, where there is price competition among multiple managed care plans. This problem may be exacerbated by the fact that over half of UC retirees were already in HMOs before there was a price differences between the HMOs and the PPO option. It may be that many of the PPO enrollees who faced the price increase had a strong aversion to more tightly managed care and were therefore willing to pay substantially higher prices rather than switch to an HMO. As a result, the UC results may understate the average demand elasticity for the Medicare population.

Since the other two studies on Medicare beneficiaries are based on national data, it would seem that they are not subject to the same concerns about generalizability as ones using data from a single firm. However, because of numerous exclusions, the samples used are not representative of the entire Medicare population either. In addition, there are other potentially important methodological problems with each study. The most significant one is that because both use cross-sectional data all the variation in prices is across plans. Since it is not possible to fully control for plan benefits and other important plan characteristics that may be correlated with price, omitted variable bias is a concern.[2]

This problem may be exacerbated by substantial ambiguity about the actual plan choice of beneficiaries. It is common for HMOs that participate in Medicare to offer multiple plans that differ in terms of both the extent of coverage and the premiums charged to beneficiaries. Nearly 40% of HMO members in the MCBS data used by Atherly, Dowd and Feldman are enrolled in HMOs offering multiple options. It is not possible in either the MCBS or the aggregate data used by Dowd, Feldman and Coulam to distinguish enrollment in a less generous, lower cost option from enrollment in the more comprehensive and more costly plan offered by the same HMO. In these cases, both studies assign enrollment to the HMOs lower cost plan. It is not clear how this imputation affects the estimated price effects.[3]

These methodological problems do not arise in the data I use in this study. Because of the way the employer's premium contributions are set, there is substantial *within* plan variation in out-of-pocket premiums. In addition, since I use administrative enrollment data I know exactly what plans an individual was offered, the out-of-pocket premiums for each plan and which option was chosen. I turn now to the details of these data.

3. Data

3.1. The Sample

This analysis is based on administrative health plan data from an employer with roughly 2700 employees located in the Southwestern United States. The data are for the years 1997–2002, during which time the number of retirees eligible for health benefits grew from 924 to 1244. Many of these are individuals who retired before age 65. There are arguments for and against including these early retirees in the analysis. Since plan benefits do not change when a retiree becomes covered by Medicare, but the required premium contributions do, pooling retirees above and below age 65 provides additional price variation. Including pre-Medicare retirees also allows for larger sample sizes. On the other hand, to the extent that

[2] The expected direction of this bias is unclear. If more expensive plans offer richer benefits or contract with more highly regarded providers, the correlation between price and plan quality will cause the effect of price to be biased toward zero. However, rules governing HMOs in Medicare may lead to a bias in the other direction. During the period analyzed in these studies, plans in areas with high capitated payment rates tended to offer additional benefits, such as richer drug coverage, and charged zero premiums.

[3] In the study by Atherly, Dowd and Feldman, there is an additional measurement problem. While one of the options they model is Medigap coverage, they lack good data on the Medigap premiums faced by individuals in their sample.

younger retirees behave differently than older ones, the results may not generalize well to the Medicare population. To mitigate these concerns, I exclude retirees under age 60 from the analysis. To further reduce heterogeneity within the sample, I also exclude individuals who retired before January 1986. These inclusion criteria result in a maximum sample size of 3230 observations on 724 retirees.

3.2. The choice set

In each year from 1997 to 2002, retirees and active employees had four health insurance options to choose from: three health plans and a cash payment for declining coverage. Two of the health plans are HMOs, which I will refer to as HMO A and HMO B in order to maintain the anonymity of the employer providing the data.[4] The other plan on the menu is a PPO. For enrollees living in its service area, the PPO has a $250 per-person deductible, a 10% coinsurance rate for providers in the plan's network, and a 50% coinsurance rate for non-network providers. For retirees outside the service area who use non-network providers, the coinsurance rate is 20%.

While all employees and retirees have the option of declining coverage, the exact menu of health plans depends on where an individual lives.[5] Approximately two-thirds of retirees face the full choice set, 7% face the choice of one HMO and a PPO option, and roughly one-quarter choose between the PPO and the cash payment for declining coverage.

3.3. The price variable

What makes these data well suited for analyzing the effect of price on the choice of health plans is that there is substantial variation in premium contributions facing retirees sharing a common choice set. Essentially, the data set combines the type of cross-sectional price variation exploited by Feldman et al. (1989) with the intertemporal variation that is the basis for the estimates by Cutler and Reber (1998), Royalty and Solomon (1999) and Strombom et al. (2002).

The cross-sectional variation comes from the way that the employer's contribution depends on a retiree's prior work history. The rules for determining the employer contribution differ across three groups of retirees: (1) those retiring on or before January 1, 1988; (2) those retiring between January 2, 1988 and January 1, 1993; and (3) those retiring after January 1, 1993. For the first group, the employer contribution covers the full cost of all plans. For individuals retiring between 1988 and 1993, the employer contribution depends on coverage tier (i.e., single, two-party) and is set below the premium of the PPO, which is the most costly plan. As a result, premiums vary by plan, coverage tier and Medicare status. All pre-1993 retirees who decline insurance coverage receive $75 per month. For

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[4] Both plans are headquartered in the same city as the employer, have long histories there and are similar in other important respects. For example, they receive comparable scores on various quality "report card" measures. In 2001, 14% of all Medicare beneficiaries in the county where the employer and the two plans are located were enrolled in Plan A, and 12.4% were enrolled in Plan B.

[5] Because the data I use comes from the same system that generates Open Enrollment materials sent to employees and retirees, there is no ambiguity as to which plans are available to each individual.

Table 1
Summary of monthly premium contributions, 2002

	PPO	HMO-A	HMO-B	Decline
Retired before January 1988				
Single coverage, Medicare	0.00	0.00	0.00	−75.00
Single coverage, pre-Medicare	0.00	0.00	0.00	−75.00
Retiree + spouse, both Medicare	0.00	0.00	0.00	−75.00
Retiree + spouse, one Medicare	0.00	0.00	0.00	−75.00
Retiree + spouse, both pre-Medicare	0.00	0.00	0.00	−75.00
Retired January 1988 to January 1993				
Single coverage, Medicare	50.24	0.00	0.00	−75.00
Single coverage, pre-Medicare	65.49	42.00	38.25	−75.00
Retiree + spouse, both Medicare	144.81	0.00	0.00	−75.00
Retiree + spouse, one Medicare	160.06	42.00	38.25	−75.00
Retiree + spouse, both pre-Medicare	160.06	104.75	93.82	−75.00
Retired after 1993, 20 years of service				
Single coverage, Medicare	97.14	20.00	14.60	−60.00
Single coverage, pre-medicare	109.34	75.65	69.03	−60.00
Retiree + spouse, both medicare	229.13	40.00	29.20	−60.00
Retiree + spouse, one medicare	244.12	95.65	85.55	−60.00
Retiree + spouse, both pre-Medicare	256.32	172.11	153.28	−60.00
Retired after 1993, 10 years of service				
Single coverage, Medicare	190.93	60.00	43.80	−30.00
Single coverage, pre-Medicare	197.03	142.96	130.40	−30.00
Retiree + spouse, both Medicare	427.76	120.00	87.60	−30.00
Retiree + spouse, one Medicare	442.78	202.96	179.95	−30.00
Retiree + spouse, both pre-Medicare	448.85	306.84	273.21	−30.00

individuals who retired after January 1, 1993 the employer's contribution decreases by a fixed percentage for each year of service less than 25.[6] Thus, for this group, out-of-pocket premiums vary within as well as across plans.

To give a sense of how the rules of the program lead to differences in relative prices, Table 1 presents the 2002 retiree premium contributions for different retiree profiles. The data in the top panel show that all plans are free for individuals who retired before January 1988. As a result, these retirees represent a "control group" when considering the effect of premiums on health plan decisions. For individuals retiring between 1988 and 1993 (second panel), out-of-pocket prices depend on coverage tier and Medicare status. In 2002, the difference between the cost of the PPO and HMO B ranged from \$27.24 (\$65.49–\$38.25)

6 Post-1993 retirees with 25 or more years of service receive the same contribution, and therefore face the same prices, as retirees in the middle cohort. For post-1993 retirees with less than 25 years of service, the out-of-pocket premium for plan j is $P_j = F_j - C(1 - .04(25 - s))$, where F_j is the plan's full premium (i.e., what the health insurer charges the employer), C is the amount the employer contributes for the middle group of retirees and s is the individual's years of service at the time of retirement.

for single, pre-Medicare coverage to $144.81 for two-party Medicare coverage; the mean difference was $70 per month.[7] For Medicare beneficiaries in this cohort the 2002 prices for the two HMOs are identical, while for pre-Medicare retirees Plan A requires a slightly larger contribution.[8]

The last two panels of Table 1 summarize the situation of post-1993 retirees with 20 and 10 years of service. For a single Medicare-eligible retiree with 20 years of service, the monthly cost for the PPO is $82.54 more than the cost of the less expensive HMO. This amount is larger than the difference facing an otherwise identical individual who retired between 1988 and 1993 ($50.24), but smaller than the difference for someone who retired at the same time with only 10 years of experience ($147.13). For all post-1993 retirees HMO A is more expensive than HMO B; this price difference is larger for retirees with fewer years of service.

One potential concern regarding the variation induced by the company's premium contribution policy is that it may have affected the timing of some employees' retirement decisions. If such effects were large, it would raise questions about the exogeneity of the price variable. As it turns out, this is not an issue, since the change in the health insurance contribution policy was enacted retroactively. An examination of the timing of retirements since 1985 suggests that, if anything, the company chose the retirement date cut-offs to fall just before spikes in retirement in the fourth quarter of 1988 and the second quarter of 1993 (data not shown).

Additional variation comes from changes in the level of the employee contribution and change in total premiums for the plans.[9] For pre-1988 retirees premium contributions are constant at zero throughout the time period analyzed, whereas relative prices have changed over time for the other two cohorts. For most individuals in most years, the PPO option is more expensive than the HMOs. The average difference between the contribution required for the PPO and the price of the lowest cost HMO option declined from 1997 to 1999 and has increased thereafter. Premiums have evolved differently for the two HMOs. In some years they have the same price for all retirees, while in other years, like 2002, this is true for some retirees but not others.

3.4. Covariates

The administrative data includes information on age, gender, marital status[10] and whether or not the covered individual is a surviving spouse of a former employee of the company. As with most studies in this literature, other correlates of medical care utilization are not

[7] This mean is calculated using the premium contributions corresponding to each individual's chosen coverage tier.

[8] Differences in the way the PPO and the two HMOs are underwritten lead to differences in the relationship between premiums and coverage tier. For example, note that for the 1988–1993 retirement cohort there is no difference in the PPO premium charged for a husband and wife who are both under age 65 and a husband and wife where one person is on Medicare and the other is not. In contrast, these two types of couples face different HMO premiums.

[9] In the regression analysis, prices are normalized to 2002 dollars using the Consumer Price Index.

[10] Unlike data used in most prior studies, this is actual marital status, not simply whether or not the retiree has chosen to cover a dependent spouse.

Table 2
Retiree characteristics by retirement date

	By retirement date			
	Before January 1986	January 1988 to December 1987	January 1988 to December 1992	After January 1993
Age	77.46 (7.58)	69.77 (5.41)	68.36 (4.42)	64.31(3.00)
% Married	46.41	71.81	69.78	75.67
% Male	51.32	74.13	70.65	70.11
% Surviving spouse	35.58	16.47	9.89	4.24
% Remaining in state	90.99	88.86	86.04	88.33
% Living in non-metro county	25.75	29.58	38.92	43.37
ZIP-level median income, 65–74 year olds ($000)	33.13 (10.04)	33.43 (9.19)	31.43 (8.97)	31.47(8.40)
Number of observations	1099	862	1254	1114
Number of retirees	222	174	258	292

available. As a result, it is not possible to test for differences in price-sensitivity related to expected health care costs or to conduct other tests related to the problem of adverse risk selection. However, the lack of health information does not pose problems for estimating average elasticities for all retirees. Retiree income is also unobserved. As a proxy, I use data from the 2000 Census on the ZIP-code level median income for households with heads between the ages of 65 and 74. The data on ZIP code is also used to create an indicator variable for retirees living outside of metropolitan areas. This variable enters the regression models to account for the fact that HMO coverage will tend to be less attractive to individuals living in rural areas.

Table 2 summarizes the observed characteristics of the retirees in the sample, stratified by retirement date. The figures illustrate why I exclude individuals who retired before 1986. They are substantially older and different in other ways from the more recent retirement cohorts. Since pre-1988 retirees face very different prices than the later two cohorts, including these older retirees would raise a concern that the estimated effect of price would be biased by unobserved heterogeneity. Individuals who retired between 1986 and 1988 are also older than more recent retirees, though the difference is not large: their mean age is 69.7 years compared to 68.2 for individuals retiring between 1988 and 1993. They are also similar to the latter two cohorts in terms of the percent who are married and the percent who are male. Individuals who retired between 1986 and 1988 are less likely to live in metropolitan areas than members of the two later cohorts. Differences among the three cohorts in the ZIP code level income variable are not significant.

3.5. The distribution of plan enrollment

To give a preliminary sense of how price affects the health plan choices of these retirees, Table 3 reports the distribution of plan enrollment for 2002 broken down by retirement cohort. Overall, the PPO is the most popular plan, enrolling almost half of the retirees in the sample. HMO B has a higher enrollment share than HMO A (32% versus 13%); 9% of

Table 3
The distribution of health plan enrollment by retirement Cohort, 2002

	Full sample	By retirement date		
		January 1986 to January 1988	January 1988 to January 1993	After January 1993
PPO (%)	45.92	63.75	52.10	30.90
HMO	44.75	34.38	36.13	51.85
HMO A	13.12	11.25	9.66	17.01
HMO B	31.63	23.13	26.47	40.63
Waive coverage	9.33	1.88	11.76	11.46
Number of observations	686	160	238	288

Notes: Figures are for retirees observed in 2002. The sample sizes are not the same as in Table 2 because that table includes all retirees observed in any year.

the sample decline coverage. Differences across retiree cohorts suggest that price is a factor affecting health plan choice decisions. PPO enrollment is greatest for pre-1988 retirees, for whom such coverage is "free" and is lowest for post-1993 retirees, who face the highest prices for the PPO option. The inverse pattern is observed for the percentage of retirees choosing HMO coverage. As would be expected, waiving is least common for pre-1988 retirees, though there is no significant difference across the other two cohorts in the percent waiving coverage.[11]

4. Econometric specification

I estimate premium elasticities based on a conditional logit regression model[12] in which the expected utility that individual i receives from plan j is assumed to be a linear function of plan attributes and individual characteristics

$$V_{ij} = \beta P_{ij} + Z_{ij}\gamma + X_i\theta j + u_{ij} \tag{1}$$

The variable of primary interest is P_{ij}, the price that individual i must pay for option j. The vector Z represents other plan attributes. Since in these data the non-price attributes of each plan (e.g., benefits, provider panels, quality reputation) are the same for all individuals,

[11] This latter result is partly explained by the fact that larger share of the post-1993 retirees are under age 65. As will be shown below, retirees who are not yet eligible for Medicare are less likely to waive coverage.

[12] A well-known limitation of this model is that it is subject to the independence of irrelevant alternatives (IIA) condition, which implies strong restrictions on the predicted substitution patterns. An alternative model that is not subject to IIA is the "mixed logit" model (McFadden and Train, 2000; Train, 2003), which allows utilities to be correlated across alternatives. I estimated several versions of mixed logit models incorporating different distributional assumptions. While these alternative models performed slightly better than the standard conditional logit model in terms of log-likelihood the models did not differ in any meaningful way in terms of the implied elasticities.

Z consists of a set of plan dummies[13] plus a dummy variable that equals one for two-party coverage and zero for single coverage.

The vector X_i includes individual characteristics that are assumed to be related to preferences for the different alternatives. Age enters as a continuous variable along with an indicator variable that equals one for individuals who are under age 65 to account for the fact that a person's outside options change discretely when she qualifies for Medicare. Since a key source of price variation is across retirement cohorts, it is important to be sure that the results are not sensitive to the treatment of age. For this reason, I also estimated models where age was parameterized as a quadratic and with a set of categorical variables. Since the estimated price elasticities are not at all sensitive to the way age enters the model, I report the results from this more parsimonious and easily interpreted specification. The other control variables are the ZIP code level median income variable and indicator variables for marital status, coverage as a surviving spouse of a former employee, and residence in a non-metropolitan county. The error term u_{ij} captures unobserved factors affecting person i's assessment of option j.

There are two potentially important issues relating to the definition of the choice set. The first is that single and married retirees face a different set of options. Whereas singles simply choose from among the three plans and the option of waiving coverage, married retirees also have the decision of whether or not to cover their spouse. Most previous studies have treated the decision to cover dependents as exogenous to price and modelled the choice among plans conditional on that decision (Feldman et al., 1989; Barringer and Mitchell, 1994; Royalty and Solomon, 1999; Buchmueller, 2000; Strombom et al., 2002). This assumption is reasonable in cases where the employer contribution covers a large share of the cost of family coverage and, therefore, there is little financial disincentive to cover dependents. However, for many retirees in this data set the incremental cost of covering a dependent spouse can be substantial and approximately 16% of married retirees in the sample choose single coverage. Thus, treating the choice of coverage tier as exogenous ignores one margin where price matters. Auxiliary regressions of the decision of married retirees to cover a dependent spouse (conditional on choosing a plan) indicate that this decision depends significantly on the difference between the price of two-party and single coverage. For this reason, I treat single and two-party coverage from each plan as distinct choice alternatives. This means that for married retirees the choice set has up to seven alternatives (three plans times two coverage tiers plus the waive option).

The second specification issue pertains to the treatment of the option of waiving coverage. Previous studies have not modelled the decision to decline coverage, presumably because in the cases analyzed there is little financial incentive to decline coverage, making such behavior very rare. However, it is clear from Tables 1 and 3 that retirees in this data set both have an incentive to decline coverage and appear to respond to that incentive. Therefore, I include waiving as an option in the choice set. For the purpose of sensitivity testing and to allow for clear comparisons with previous work, I also estimate models on a restricted sample of individuals who chose to enroll in one of the health plans

[13] Chernew et al. (2002) interpret these dummies to represent consumer's average assessment of plan quality. It is likely that that they also pick up such things as provider panel size and the convenience of provider locations.

offered by their former employer. This is equivalent to treating the waive option and the three plans as being in two distinct "nests" and estimating the choice within the plan nest.

5. Results

5.1. Main results

Table 4 presents coefficient estimates from the conditional logit model. The first column is for the full sample, in which married and single retirees are pooled. In columns 2 and 3, the sample is stratified by marital status.

Since the option of waiving coverage is treated as the baseline alternative, the coefficients on the interactions of individual characteristics and the plan dummies are interpreted as the effect of a characteristic on the probability of choosing the particular plan relative to the probability of waiving coverage. Thus, the fact that all the coefficients for the interactions with the under 65 dummy variable are positive means that pre-Medicare retirees are more likely to accept coverage than those over age 65. Because for all three plans the benefits do not change when someone enrolls in Medicare, I interpret this result to reflect the high premiums and other barriers faced by the near elderly in the market for non-group health insurance. Conditional on Medicare coverage, in the married subsample the demand for PPO coverage increases with age, which is consistent with the findings of previous studies (Feldman et al., 1989; Barringer and Mitchell, 1994; Cutler and Reber, 1998; Royalty and Solomon, 1999; Buchmueller, 2000). These age effects, however, are imprecisely estimated; most of the coefficients do not achieve conventional levels of statistical significance.

The pattern of the coefficients on the marital status and gender interactions suggest that married men are more likely to take-up coverage than married women, while gender differences among single retirees are not statistically significant. This is consistent with previous research on the take-up behavior of active workers (Buchmueller, 1996/97). Retirees living outside of metropolitan areas are more likely to choose PPO coverage than to either decline coverage or enroll in an HMO.[14]

In all models, the estimated coefficient on the out-of-pocket premium is negative and statistically significant at the 0.01 level. The price coefficient is larger in absolute value for single retirees than for married retirees. Beyond this, the coefficients themselves are not useful for gauging the magnitude of the price effect or comparing effects across samples. Therefore, I calculate two quantities that give a sense of the magnitude of the price effect: "enrollee perspective" premium elasticities[15] and the predicted change in market share caused by a $5 increase in premiums. For the conditional logit model, the own-price elasticity

[14] Note that each HMO option only appears in an individual's choice set if the plan is truly an option i.e., if the person lives in the plan's service area.

[15] The enrollee perspective elasticity is based on the out-of-pocket prices faced by enrolees i.e., the total premium less the employer contribution. Evaluating the same price effect at the total premium charged by the health plan results in a much larger elasticity. I discuss this below.

Table 4
Conditional logit coefficients: retirees age 60 and older

	(1) All retirees	(2) Married	(3) Single
Premium	−0.0074 (0.0015)	−0.0065 (0.0015)	−0.0181 (0.0060)
Two party coverage	2.182 (0.170)	2.132 (0.169)	
Married × PPO	−1.558 (0.398)		
Married × HMO A	−2.358 (0.463)		
Married × HMO B	−1.766 (0.435)		
Age (in years) × PPO	0.054 (0.037)	0.132 (0.045)	−0.045 (0.046)
Age (in years) × HMO A	0.003 (0.051)	0.088 (0.058)	−0.010 (0.088)
Age (in years) × HMO B	−0.078 (0.047)	0.022 (0.056)	−0.220 (0.074)
Age < 65 (0,1) × PPO	0.392 (0.336)	0.678 (0.386)	0.443 (0.658)
Age < 65 (0,1) × HMO A	0.454 (0.439)	0.772 (0.512)	0.495 (0.830)
Age < 65 (0,1) × HMO B	0.407 (0.397)	1.084 (0.381)	−0.668 (0.794)
Male × PPO	1.054 (0.344)	1.084 (0.381)	0.995 (0.644)
Male × HMO A	0.791 (0.422)	1.071 (0.523)	0.356 (0.774)
Male × HMO B	0.684 (0.376)	0.736 (0.411)	0.672 (0.726)
Surviving spouse × PPO	1.071 (0.558)		1.123 (0.626)
Surviving spouse × HMO A	−0.506 (0.728)		−0.862 (0.808)
Surviving spouse × HMO B	−0.536 (0.676)		−0.670 (0.761)
ZIP code level average income × PPO	0.0160 (0.015)	0.023 (0.017)	0.015 (0.032)
ZIP code level average income × HMO A	−0.007 (0.021)	0.012 (0.023)	−0.063 (0.040)
ZIP code level average income × HMO B	0.000 (0.017)	0.015 (0.018)	−0.037 (0.034)
Non-metropolitan county × PPO	1.539 (0.318)	1.269 (0.374)	2.704 (0.805)
Non-metropolitan county × HMO A	0.152 (0.469)	0.026 (0.545)	0.859 (1.001)
Non-metropolitan county × HMO B	0.346 (0.388)	0.015 (0.434)	1.777 (1.036)
Number of observations	3206	2383	823
Log likelihood	−3430.64	−2734.14	−654.084

Notes: The choice set includes available health plans plus the option of waiving coverage, which is treated as the omitted option. For married retirees single and two-party coverage are treated as separate options. Huber-White standard errors (in parentheses) account for the fact that there are multiple observations per individual. All models include main effects for each plan (three variables) and interactions with year dummies (15 variables). The surviving spouse dummy is excluded from the married sample regressions because surviving spouses do not have the option of two-party coverage.

of demand is given by

$$\eta = \frac{\partial \ln \mathrm{Prob}_j}{\partial \ln P_j} = \beta P_j(1 - \mathrm{Prob}_j) \tag{2}$$

where β is the price coefficient, P_j is the price of plan option j and Prob_j is the probability that j is chosen. The fact that the elasticity depends on plan prices and market shares means that a given model will generate different elasticity estimates for different plans.[16] Therefore, it is necessary to be careful in making comparisons across studies as differences in elasticities may be driven not only by real differences in price sensitivity but also by differences in average prices facing consumers. The predicted loss of market share is less sensitive to the level of observed prices. It is calculated by predicting for each observation the probability of choosing a plan at given prices and the probability of choosing that plan if its premium is increased by $5 while the price of all other plans remained constant. The difference between these two predictions is calculated for each person and then averaged over the estimation sample.

The elasticity estimates and their standard errors are reported in the first column of Table 5. In the second column is the mean change in the probability of choosing each option brought about by the $5 price increase. The last column expresses this effect as a percent of the initial market share. For the full sample, the elasticities range from −0.14 to −0.37 depending on the plan. Evaluated at the overall sample means, the estimated enrollee perspective elasticity is −0.27 with a 95% confidence interval of [−0.38, −0.17]. Using the full premium rather than the out-of-pocket price actually faced by retiree's yields an "insurer perspective" elasticity of −3.3. The simulations indicate that a $5 price increase (holding constant the price of competing plans) reduces the probability that a plan is chosen by between 0.002 and 0.008. Relative to predicted probabilities using observed prices, this represents a loss of market share of between 1.3 and 3%.

The results for the married subsample are quite similar, which is not surprising given that roughly three-quarters of the full sample is married. The price effects are stronger for single retirees. For singles, a $5 price increase is predicted to reduce enrollment by an average of 1.5% points. Relative to the mean baseline market share among singles, this represents a decline of 4.5%. Combining the parameter estimates from the single sample regressions with the mean premium for single coverage yields an insurer perspective elasticity of −3.5.

5.2. Alternative specifications of the choice set

The stronger price effect for single retirees may be due to the fact that switching plans is more costly for two people rather than one. Alternatively, this result could simply be due to the fact that the choice among plans, which is the only decision facing single retirees, is more sensitive to price than the decision to cover a dependent spouse, which is incorporated into the price effect in the married sample. To test for this possibility, I estimate a set of models that for married retiree's conditions on the choice of coverage tier. That is, similar to previous studies, I treat a married individual's choice between single and two party coverage as exogenous and then model the choice among the different plans within each person's chosen coverage tier.

The key results from these regressions are reported in Table 6. The price coefficient, elasticities and simulated market share effects are quite close to the ones reported in

[16] I calculate the elasticities at the sample means for each plan and at the overall sample means for the estimation. An alternative approach is to calculate individual-level elasticities using the prices facing each observation and the predicted probability of choosing each plan and then take the average over the estimation sample. The two techniques yield similar results.

Table 5
Estimated price effects

	Elasticity	Effect of a $5 price increase	
		Change in market share	As a % of initial market share
All retirees			
PPO, single	−0.215 (0.042)	−0.004	−2.87
PPO, two-party	−0.361 (0.071)	−0.007	−1.87
HMO A, single	−0.161 (0.032)	−0.002	−3.35
HMO A, two-party	−0.375 (0.074)	−0.004	−3.11
HMO B, Single	−0.140 (0.028)	−0.003	−3.19
HMO B, two-party	−0.255 (0.050)	−0.008	−2.34
Average over all plans	−0.272 (0.054)	−0.004	−2.78
Married retirees			
PPO, single	−0.196 (0.045)	−0.002	−2.92
PPO, two-party	−0.317 (0.072)	−0.006	−1.64
HMO A, single	−0.147 (0.033)	−0.001	−3.12
HMO A, two-party	−0.329 (0.075)	−0.004	−2.74
HMO B, single	−0.132 (0.030)	−0.002	−3.01
HMO B, two-party	−0.224 (0.051)	−0.007	−2.06
Average over all plans	−0.264 (0.060)	−0.004	−2.54
Single retirees			
PPO	−0.305 (0.102)	−0.013	−3.89
HMO A	−0.341 (0.114)	−0.015	−6.77
HMO B	−0.268 (0.114)	−0.017	−6.05
Average over all plans	−0.325 (0.108)	−0.015	−5.29

Notes: Price effects are based on the conditional logit results reported in Table 4. Elasticities are calculated at the mean price and market share for each plan in each estimation sample. Standard errors for the elasticity estimates are presented in parentheses. The effect of a $5 price increase is the estimated loss of market share for a plan that raises its premium by $5 while other plan premiums remain constant. It is calculated for each observation and then averaged over the estimation sample.

Tables 4 and 5, indicating that for married retirees price has a similar effect on the choice among plans and the decision to take two-party rather than single coverage. This implies that the difference between married and single retirees represents a difference in behavior, rather than an artifact of the model.

As noted, another difference between the results presented in Tables 4 and 5 and those from previous studies is that the previous studies do not include the decision to waive coverage in the choice set. Therefore, I also estimate models on the subset of people who elect to take coverage. For the sake of brevity, I do not report these models in a separate table, but instead describe them here.

For married couples, conditioning on take-up does not change the results. The price coefficient is −0.0077, which is essentially identical to the estimate from the baseline model (−0.0074). The two models imply similar elasticities and the same loss of market share associated with a price increase. For single retirees, the estimated price effects are

Table 6
Plan choice conditional on coverage tier: estimated price effects

	Price coefficient	Elasticity	Effect of a $5 price increase	
			Change in market share	As a % of initial market share
All retirees [$N = 3206$]	−0.0068 (0.0015)			
PPO		−0.212 (0.047)	−0.005	−1.44
HMO A		−0.248 (0.055)	−0.005	−2.75
HMO B		−0.168 (0.037)	−0.007	−2.06
Average over all plans		−0.224 (0.050)	−0.006	−1.98
Married retirees [$N = 2383$]	−0.0054 (0.0015)			
PPO		−0.194 (0.055)	−0.004	−1.14
HMO A		−0.231 (0.066)	−0.004	−2.29
HMO B		−0.175 (0.050)	−0.006	−1.57
Average over all plans		−0.205 (0.058)	−0.004	−1.57

Notes: The choice set includes available health plans plus the option of waiving coverage. The decision of whether or not to cover a dependent spouse is treated as exogenous. Huber–White standard errors (in parentheses) account for the fact that there are multiple observations per individual. Independent variables are the same as in Table 4. Elasticities are calculated at the mean price and market share for each plan in each estimation sample. The effect of a $5 price increase is the estimated loss of market share for a plan that raises its premium by $5 while other plan premiums remain constant. It is calculated for each observation and then averaged over the estimation sample.

larger when the sample is limited to people taking up coverage. The elasticity (averaged over all plans) is −0.490, compared to the elasticity of −0.325 reported in Table 5. This pattern is consistent with other research showing that for active employees the decision to take-up coverage offered by an employer is less sensitive to price than the choice among a set of plans conditional on having some coverage.[17]

5.3. Additional sensitivity tests

To test the sensitivity of the results, I estimate the conditional logit model on a number of alternative samples. First, I cut the sample in different ways by retirement date. This is important given that price differences across the three retirement cohorts are an important source of identifying variation. First, I drop individuals who retired before January 1988. This leaves a sample in which nearly everyone faces a higher price for the PPO compared to the HMOs, though the size of that differential varies. Many, though not all, individuals in this sample face different prices for the two HMOs. I also estimate a set of regressions on a sample that excludes the post-1993 cohort of retirees. In this case, the main contrast is between pre-1988 retirees, for whom all plans are free, and later retirees who face a higher price for the PPO and, in some cases, differences in prices between the two HMOs. The results for these two sub-samples (available upon request) are quite similar to the results for the full sample.

[17] See, for example, Cutler (2003) and Gruber and Washington (2005).

Next, I consider the effects of limiting the sample to retirees over the age of 65. The results are robust to this change as well. A $5 increase in premiums is predicted to reduce a plan's enrollment by 2%. This is only slightly smaller than the 2.8% market share reduction predicted for the full sample. Since excluding pre-Medicare retirees eliminates a different source of price variation, this provides another check on the identification strategy.

Finally, to test for the stability of the price effects over time, I cut the data by year, estimating separate models for the periods 1997–1999 and 2000–2002. The estimated price coefficients for the two periods are essentially identical.

5.4. Comparisons with previous studies

These results can be put in perspective by comparing them to estimated price effects from other studies. The most direct comparisons are with the two studies on the health plan choices of retirees. In my earlier study on UC retirees, which modeled the demand for PPO coverage, the mean PPO market share was 51% and the mean out-of-pocket premium for that option was $64 in 2002 dollars. The corresponding figures for this data set are almost identical: 51% of retirees electing coverage chose the PPO option and the mean PPO premium was $62. These similarities in price and market share make it straightforward to compare elasticities. The main difference between the two studies is that in the current study there is much more price variation, including price differences for the different HMO options. Since in the earlier study I conditioned on coverage tier, the most appropriate results for making comparisons are those reported in Table 6. For this model, the estimated elasticity for the PPO plan is −0.21. In the UC study, the full sample elasticity is slightly lower than this (−0.14), though within the 95% confidence interval of the current estimate [−0.30 to −0.12].

The premiums observed in this data set are higher than in the data used by Atherly et al. (2004), where the mean premium is $10.17 and Dowd et al. (2003), where the mean is $8.16. This difference largely explains why their estimated elasticities (−0.13 and −0.14) are slightly lower than the ones found here. When compared in terms of the effect of a small price increase, the results look more similar. In my full sample, a $5 increase in premiums is predicted to reduce a plan's market share by an average of 0.004. The corresponding estimate from Atherly et al. is 0.003. This similarity suggests that in their study the bias from unmeasured plan attributes is not large.

The estimated price effects from this study are at the lower end of the results of studies based on non-elderly active employees, though they are not substantially different than those earlier results (and the confidence intervals overlap). Recall that Cutler and Reber (1998) estimate premium elasticities of −0.3 and −0.6. Measured relative to the full premiums in their sample, their price effects correspond to an insurer perspective elasticity of −2, which is smaller than the insurer perspective elasticities implied by my results.

In terms of similarities in research design, the cleanest comparison is with the paper by Strombom et al. (2002), which provides estimated price effects for different groups of active UC employees. They estimate conditional logit models using the full set of plans available to UC employees-a FFS plan, a PPO and several HMOs-and on a subsample that excludes individuals in the FFS plan. The full sample results are similar to those here: a $5 increase in premiums is predicted to reduce a plan's market share by 3% and the insurer-perspective elasticity ranges from −0.8 to −5.2, depending on plan, with an average of −2.5. However, the restricted choice set consisting only of managed care plans is more similar to the choice set in the current data. Those regressions imply stronger price effects: an insurer-perspective elasticity of −5.3 and a 7.6% loss of market share in response to a $5 price increase.

5.5. Implications for policy and research

These results have implications for the incentives that health plans would face if Medicare were restructured as a managed competition program. While the enrollee perspective elasticities may seem small, the insurer perspective elasticities and simulated market share effects indicate that health plans that raise premiums while their competitors hold premiums constant will lose a non-trivial share of their enrollment. The similarity between my results and those based on data from Harvard University and the UC suggests that the experiences of those employers are, in fact, relevant to Medicare. In both of those cases, the adoption of a fixed dollar premium contribution led to a reduction in health spending. Spending fell not only because employees shifted to lower cost plans but also because participating plans responded to this shift by reducing premiums (Cutler and Reber, 1998; Buchmueller, 1998). Taken together, the results from this and earlier studies provide some support for the arguments made by proponents of market-oriented reforms.[18]

The vigor with which plans compete on price will depend on how exactly the program is structured. If payments to plans are not fully risk adjusted, plans will have an incentive to attract healthier than average enrollees. To the extent that the price elasticity of demand is greater for healthier enrollees than for those in poor health, as has been found in previous studies, the incentive to compete on price will be even stronger. Put another way, the premium elasticities I estimate represent average effects, which understate the price sensitivity of those enrollees that plans are most interested in attracting. Further research on the relationship between health status and price sensitivity among Medicare beneficiaries – which also has implications for risk selection and market stability – would be valuable.

It is very useful for policy analysts to have evidence from a number of different settings and estimates based on different sources of variation. This is particularly true in the literature on health plan choice since most studies use data from a single employer. The fact that the elasticity estimates in this study are similar to those from other studies using different data sources is therefore important. Modelling the impact of a major policy change such as Medicare reform requires many assumptions, which introduce significant uncertainty. The tight range of premium elasticity estimates that has emerged from this literature suggests that this parameter is not a major source of uncertainty.

[18] Several caveats concerning this conclusion should be noted. First, it is not clear from those earlier studies whether adopting a managed competition approach affected the growth in spending or simply led to a one-time savings. In addition, applying this model to Medicare would involve significant transition costs, which were not an issue in those employer-sponsored programs. Plus, competition among multiple managed care plans may not be feasible in less densely populated parts of the country. Even where competition is feasible, there may not be a political willingness to expose traditional fee-for-service Medicare to competition, as occurred in the Harvard and UC examples. As a result, even if adopting the principles of managed competition did lead to more price competition in certain markets, the ultimate effect on Medicare spending is unclear. These details, while important, are beyond the scope of this study.

In addition to its relevance to Medicare policy, the results of this study have implications for research on the determinants of consumer health plan choice decisions. In multiple option health insurance programs, consumers make choices along several margins. In addition to the choice among competing plans there is the decision to take-up coverage at all and for married individuals there is also the decision to cover dependents. This study is the first in the literature to jointly consider the effect of price on all of these margins. The results suggest that for married retirees price has a similar effect on the decision to take-up coverage, the decision to cover a spouse and the choice among plans. As a consequence, models that condition either on the decision to accept coverage or the decision of whether or not to cover a dependent spouse yield similar price elasticities as a model that treats these two decisions as endogenous. For single retirees, the take-up decision is less sensitive to price than the choice among plans conditional on take-up. However, the difference is small. Therefore, models that condition on take-up and models that treat the decision to decline coverage as part of the choice set yield similar qualitative results.

6. Summary and conclusions

Prominent Medicare reform proposals call for a greater reliance on price-sensitive consumer demand as a force for driving competition and controlling costs. The premium elasticity of demand is a key parameter for understanding how Medicare beneficiaries would behave under such reforms. While previous studies provide elasticity estimates, this literature has important shortcomings. The studies with the strongest research design focus on younger, active employees, while the studies of older retirees have potentially important methodological limitations.

In this paper I provide estimates of health insurance premium elasticities that are directly relevant for understanding how Medicare beneficiaries would behave in a managed competition setting. I extend the literature by analyzing retirees rather than active employees and by using a quasi-experimental research design that exploits exogenous variation in health plan premiums. This research design produces more credible elasticity estimates than those derived from previous studies. The findings indicate that retirees do consider price when choosing among competing health plans and are willing to switch plans when relative prices change. The effect of out-of-pocket premiums on the health plan choice decision is negative, statistically significant, and very robust to different modelling strategies. For most estimation samples the implied enrollee perspective premium elasticity falls between -0.2 and -0.3. The regression results imply that a health plan that increased its premium by \$5 while its competitors held their prices constant would lose between 2 and 4% of its enrollees.

The results of this and other studies suggest that health plans competing in a reformed Medicare program will face considerable pressure to compete on price. Additional research at the level of the health plan would provide further insight on how exactly plans would compete under the incentives of a premium support program. The Medicare Modernization Act of 2003 included provisions that would establish competitive bidding demonstration projects in several markets. While the history of such projects in Medicare is not promising

– all previous demonstrations were cancelled before going into effect – these projects have the potential to provide important evidence on how health plans compete in a managed competition environment.

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