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Switching costs, price sensitivity and health plan choice

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Abstract

We investigate the extent to which sensitivity to health plan premiums differs across individuals according to characteristics related to the cost of switching plans. Our results indicate substantial variation in price sensitivity related to expected health care costs: younger, healthier employees are between two and four times more sensitive to price than employees who are older and who have been recently hospitalized or diagnosed with cancer. We also find evidence of status quo bias: estimated premium elasticities are significantly higher for new hires than for incumbent employees. Simulations combining our results with actuarial data illustrate the cost implications of risk-related differences in price elasticity. © 2002 Elsevier Science B.V. All rights reserved.

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1. Introduction

Price-conscious consumer choice is a cornerstone of market-oriented health care reform proposals and the purchasing strategies used by many large employers. In the managed competition model, individuals make health insurance choices during an annual open enrollment period. Because the default option for those previously enrolled is to remain in their current plan, the price sensitivity of incumbent employees depends on how costly it is

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to switch plans. To the extent that the cost of switching health plans varies across individuals, there will also be heterogeneity in consumers' sensitivity to health plan premiums. In particular, since switching health plans often requires changing medical providers as well, switching costs are likely to be higher—and hence price sensitivity lower—for individuals with higher expected medical expenses.

A number of studies have examined the effect of premiums and other factors on consumer health plan choices. A consistent finding from this literature is that, *on average*, consumers are quite sensitive to price. However, the existing literature provides very little evidence on how price sensitivity varies across consumers. Understanding how price elasticity varies with consumer characteristics is important for evaluating the potential effects of market-oriented health policies. If individuals who are in poor health are substantially less price-sensitive than healthier consumers, plans that increase in price relative to their competitors will not only lose enrollment, but will experience rising costs due to "adverse retention". Such a result has potentially serious implications for economic efficiency. Without effective risk adjustment, differences in enrollee health status will generate differences in premiums between more generous plans and less generous plans that exceed true cost differences. In the extreme, biased risk selection can attenuate the incentive for plans to compete on quality.

In this paper we investigate the extent to which health plan premium elasticities vary across individuals according to three factors related to the cost of switching health plans: health status, age and employment tenure. The analysis is based on a unique data set that combines 5 years of open enrollment data from a large multi-location employer with information from a statewide cancer registry and a statewide hospital discharge database. No previous study has used data on such a large sample of individuals over such a long period of time or has incorporated comparable information on health status.

Our results indicate large differences in demand elasticity associated with health status, age, and job tenure. While significant price effects are estimated for all groups, the elasticities for individuals presumed to face the lowest switching costs—younger, recently hired employees who are in good health—are roughly four times larger than those estimated for individuals for whom switching plans is likely to be more costly—older, incumbent employees who have recently been hospitalized or diagnosed with cancer. Combining our regression results with actuarial cost data, we show how this heterogeneity in price sensitivity affects the expected claims costs of a plan that raises its premiums and competing plans that do not. In some cases, nearly half the price increase is offset by the increase in costs resulting from the change in a plan's risk pool.

While based on data for active employees, our results have implications for recent proposals for restructuring the Medicare program according to the principles of managed competition. Our findings suggest that Medicare beneficiaries are likely to be substantially less price-sensitive than non-elderly workers, which corroborates the results of earlier work on the health plan choices of retirees (Buchmueller, 2000).

In Section 2, we review the previous literature relating to switching costs and consumer health plan choice decisions. Sections 3 and 4 describe our data and econometric models, respectively. We report our regression results in Section 5, and in Section 6 present several simulations that illustrate the cost implications of these results. Concluding comments are in Section 7.

2. Background and previous literature

2.1. Switching costs in health insurance

There is a growing interest among economists in markets where it is costly for consumers to switch among competing suppliers. A review by Klemperer (1995) identifies several categories of switching costs, most of which arise in connection with the purchase of health insurance. There are transactions costs, which in the case of insurance include not only the direct cost of enrollment, but also the cost of researching alternative plans. Since health care is an experience good, consumers will have better information on the quality of their current plan than on the quality of its competitors. This information asymmetry makes plan changes costly for risk-adverse individuals. In addition, because health plans vary in their rules and procedures, there are costs associated with learning a new system once someone changes plans. The one type of switching costs that is unique to health care is arguably the most important. Consumers who switch managed care plans often must sever relationships with previous health care providers and begin relationships with new ones. The link between providers and plans will generate greater “brand loyalty” than that which occurs in other product markets.

Persistence in plan choices arising from these factors can be thought of as a response to real, albeit non-pecuniary, costs. These costs, particularly those related to changing providers, are likely to be larger for individuals with greater medical care utilization. As a result, we would expect older consumers and those with serious health conditions to have lower elasticities of demand for health insurance than their younger, healthier counterparts. Assuming that patient–physician bonds grow stronger over time, price sensitivity will also decrease with the time a person has been in a plan. Persistence may also result from apparent deviations from rational behavior (Samuelson and Zeckhauser, 1988). That is, even when transactions costs are low and switching health plans does not require changing physicians, consumers may stay in the same health plan despite changes in the factors that influenced the initial choice. [1] In an employer–sponsored benefits program, “status quo bias” will result in differences in the choices made by new and incumbent employees. While these differences in behavior do not represent a direct effect of health status, because new employees tend to be younger than incumbent ones, status quo bias will also result in a negative correlation between expected medical expenditures and price sensitivity.

2.2. Previous studies on health plan choice and switching decisions

A number of studies have estimated the effect of price and other factors on health plan choice decisions, typically using data on one employer, or several employers in the same market (McGuire, 1981; Feldman et al., 1989; Barringer and Mitchell, 1994; Cutler and Reber, 1998; Chernew and Scanlon, 1998; Royalty and Solomon, 1999). While most of these studies estimate the effect of price (i.e. employee premium contributions), very little

[1] This type of inertia has also been shown to be prevalent in decisions concerning other employee benefits, such as flexible medical spending accounts (Schweitzer et al., 1996) and retirement savings plans (Samuelson and Zeckhauser, 1988; Madrian and Shea, 2000).

attention has been paid to how sensitivity to price varies across consumers. That is, they do not account for potential differences in behavior related to health risk and they treat initial plan choices and subsequent decisions of whether or not to switch plans as equivalent, ignoring the possibility that demand becomes less elastic as consumers become "locked in" to earlier choices. [2]

Another important limitation of much of the health plan choice literature stems from the use of cross-section data. With data from a single cross-section, the effect of premiums and other plan attributes are estimated on the basis of variation across plans. If unmeasured plan attributes are correlated with premiums—e.g. if premiums are higher for plans with more generous benefits or those contracting with more highly regarded providers—parameter estimates will suffer from omitted variable bias. This is a likely explanation for the positive effect of premiums indicated by some of Barringer and Mitchell's (1994) models, and for some of Chernew and Scanlon's (1998) counterintuitive findings regarding the effect of various proxies for plan quality.

One study that is not subject to these criticisms is a recent paper by Royalty and Solomon (1999), which uses 2 years of data on employees of Stanford University. They estimate a set of choice models in which the premium variable is interacted with employee age, job tenure and self-reported health status. Their results suggest larger premium elasticities for groups who are likely to face lower switching costs—younger employees, new hires and individuals reporting no chronic health conditions—though in many cases the differences across employee groups are not statistically significant. For individuals with self-reported health conditions, the estimated effect of price is not significantly different from zero, though this is mainly due to large standard errors. Because these factors were interacted with the premium variable separately, i.e. age was interacted with the premium in one model, health status in another—the results are difficult to interpret. For example, it is not possible to determine whether price sensitivity varies independently with both age and job tenure, or if only one of these factors matters and the different models where these variables are interacted with price are capturing the same effect.

Our analysis is similar to Royalty and Solomon's (1999), though three features of our data set allow for improvements over their work and the rest of the literature. First, with data on whether each employee and spouse has been diagnosed with cancer or hospitalized in the months immediately around each open enrollment period we are able to construct objective measures of health status that should be relevant to health plan choice decisions. Second, because we have 5 years of data from 11 locations, there is substantial variation in premiums within as well as across health plans. As we discuss in Section 3, this variation is driven largely by changes in the employer premium contribution and is therefore plausibly exogenous to plan characteristics and individual health plan choices. The third major strength of our data set is its size: we have information on over 100,000 distinct employees. To put this in perspective, Royalty and Solomon's (1999) models that include

[2] A related literature examines plan switching directly (Wersenger and Sorenson, 1982; Long et al., 1988; Buchmueller and Feldstein, 1997; Altman et al., 1998). A common finding is that switchers are younger and have lower prior utilization than continuing enrollees, which is easily explained by differences in switching costs. By definition, it is not possible in an analysis of switching behavior to compare the behavior of individuals who are making choices for the first time and those whose decision is whether or not to change plans. No study in this literature tests for differences in price sensitivity related to health risk.

health status are estimated on a sample of roughly 1200 observations. Our larger data set allows for precise estimates of differences across subgroups within the sample, even when certain groups that are very important from a policy perspective—such as those in poor health—represent a small fraction of the population. Moreover, we are able to allow the effect of price to vary simultaneously with age, job tenure and health status. We now turn to the details of our data.

3. Data and descriptive results

3.1. Data on UC employees

Our main source of data is the health benefits program of the University of California (UC), which has employees at 11 locations throughout California. [3] Our data span the period from 1993 to 1997, during which the number of employees offered a choice of health plans (and for whom complete data are available) ranged between 65,933 and 73,980 per year. Overall, we have complete data on 103,835 distinct employees.

The fact that we observe many employees in multiple years creates a potential problem for our estimation. If we were to use all these observations, the error terms in our regressions would be correlated across observations, causing estimates of the standard errors to be understated. To avoid this problem, we use a sampling technique that limits the estimation sample to one observation per employee. For individuals who were employed at the UC in more than 1 year between 1993 and 1997, we selected one observation at random and discarded the remaining observations for that person.

The UC administrative data files provide information on the following employee characteristics: gender, age, salary, job tenure, and job classification (academic, management or support staff). We also have information on the home ZIP code of each employee, which we use to identify those living in rural areas. Table 1 presents summary statistics on these employee characteristics for our estimation sample and, for comparison purposes, the population of employees for the year 1995. [4] Because of the way it is formed, our estimation sample has a higher percentage of new hires than a cross-section from any year. This causes the estimation sample to differ from the UC employee population in terms of some characteristics, such as age, salary and the presence of dependents, though the differences are small.

An important limitation of administrative data for analyzing health plan choices is a lack of information on health status or medical care utilization. To address this problem, we matched the UC data on employees and (for married workers) their spouses with data from the California Cancer Registry (CCR) and hospital discharge data from the California Office of Statewide Health Planning and Development (OSHPD). [5] From these auxiliary data sources we can identify UC employees and spouses diagnosed with cancer between

[3] The UC also operates the Los Alamos National Laboratory in New Mexico. Because our cancer and hospital discharge data are from California only, we exclude New Mexico employees from our analysis.

[4] Because employee demographic characteristics are fairly constant over the period, we report data for a single year for the sake of brevity.

[5] To ensure confidentiality, the linkage was performed by researchers at the CCR.

Table 1
Employee characteristics—summary statistics for the estimation sample and full sample from 1995[a]

	Estimation sample	Full sample, 1995
Age (years)	40.13 (10.21)	41.70 (10.28)
Job tenure (years)	7.11 (7.67)	8.70 (7.89)
Annual salary (US$)	51266 (35611)	53688 (36201)
Male (0, 1)	0.468	0.478
Lives in a rural county (0, 1)	0.011	0.011
Coverage tier		
Single coverage (0, 1)	0.481	0.406
Two party coverage (0, 1)	0.171	0.200
Family coverage (0, 1)	0.342	0.394
Job category		
Academic (0, 1)	0.208	0.204
Management/professional (0, 1)	0.033	0.034
Other (0, 1)	0.760	0.762
Diagnosed with cancer		
In 22 months prior to open enrollment (0, 1)	0.007	0.008
In 12 months after open enrollment (0, 1)	0.005	0.006
In 34 months around open enrollment (0, 1)	0.012	0.013
Hospitalized (other than maternity)		
In 22 months prior to open enrollment	0.064	0.069
In 12 months after open enrollment	0.040	0.043
In 34 months around open enrollment	0.091	0.103
Cancer or non-maternity hospitalization		
In 22 months prior to open enrollment	0.067	0.072
In 12 months after open enrollment	0.042	0.046
In 34 months around open enrollment	0.100	0.108
Number of employees	103835	72368

[a] Standard errors of continuous variables are reported in parentheses. Salary figures are in 1997 US$. The source of the cancer data is the California Cancer Registry. The source of the hospitalization data is the California Office of Statewide Health Planning and Development. The rural area indicator is based on ZIP code-level data from the 1990 census. All other variables are from the administrative files of the UC health benefits program.

1 January 1988 and 31 December 1997 or hospitalized between 1 January 1991 and 31 December 1997.

In our analysis we use the CCR/OSHPD data to identify "high-risk" individuals who are likely to face higher than average switching costs due to strong ties with particular providers. We create indicator variables that equal one for observations where an employee and/or spouse were hospitalized (excluding maternity cases) or diagnosed with cancer during several different intervals around the open enrollment period. The longest period of time prior to open enrollment for which we have both inpatient utilization and cancer data is 22 months. Some 6.4% of our estimation sample was hospitalized in the 22 months prior to open enrollment and 0.7% had been diagnosed with cancer. When we narrow the window to 12 months prior to open enrollment, 3.8 and 0.4% were hospitalized or diagnosed with cancer, respectively.

The rationale for defining risk based on prior utilization is that whatever information utilization provides for forecasting future health needs will be known by employees at the time they choose their plans for the subsequent year. Individuals with a recent hospitalization or cancer diagnosis will not only have used more health care services in the recent past, but may also expect higher-than-average future utilization and are therefore likely to have stronger ties to their medical providers. However, defining risk status solely on prior utilization will misclassify some individuals who at the time when insurance choices are made are anticipating future inpatient utilization. To account for this possibility, we also estimate models in which high-risk individuals are defined as those with a hospitalization or cancer diagnoses in either the 22 months prior to open enrollment or the subsequent 12 months. Approximately 10% of our estimation sample is classified as high-risk according to this broader definition. We report results using this broader definition, though as we describe in Section 5, our qualitative findings are not sensitive to the use of alternative risk measures.

3.2. The choice set

Table 2 provides summary information on the plan offerings, premiums, and enrollment patterns by plan type for the period of our analysis. During this time, UC employees were offered a choice of between four and seven health insurance options, depending on location.

Table 2
Plan offerings, monthly premiums and market share by plan type in the University of California health benefits program, 1993–1997[a]

	1993	1994	1995	1996	1997
Indemnity plan					
Number of plans	1	1	1	1	1
Monthly employee premium (US$)					
Min	69.05	123.59	140.42	279.71	390.64
Max	168.60	259.71	289.91	632.91	902.42
UC employees enrolled (%)	10.0	5.2	3.9	1.4	0.6
PPO/POS					
Number of plans	3	1	1	1	1
Monthly employee premium (US$)					
Min	0	0	4.17	9.62	9.39
Max	23.53	0	10.97	26.61	25.98
UC employees enrolled (%)	21.1	16.3	19.4	22.0	23.3
HMO					
Number of plans	7	7	6	6	5
Monthly employee premium (US$)					
Min	0	0	0	0	0
Max	0	28.85	47.32	27.09	1.54
UC employees enrolled (%)	68.9	78.5	76.7	76.6	76.1

[a] Premiums are expressed in 1997 US$. Premiums vary according to three coverage tier categories: single, two-party, and family. All minimum premiums reported are for single coverage and all maximums are for family coverage. The number of plans is total number offered statewide. The choice set at any particular campus never included more than one PPO. The number of HMOs offered at any one location ranges from 2 to 5.

At all locations, the choice set included at least two HMOs offering a standard set of benefits.[6] In 1993, 68.9% of UC employees were enrolled in an HMO. That figure increased to 78.5% in the following year and remained fairly constant thereafter. Throughout the period, the UC menu included two other plans, Prudential High Option, a managed indemnity plan[7] and UC Care, which was a PPO through 1995 and a three-tiered point-of-service (POS) plan thereafter.[8,9]

In 1993, the UC premium contribution was a weighted average of the premiums charged by the four plans with the greatest UC enrollment. Since this group included the most expensive option on the menu (Prudential), the contribution exceeded the total premium of all the managed care plans, which therefore required no employee contribution. Since 1994, the UC contribution has been based on the amount charged by the lowest cost plan available statewide. This policy change generated price increases between 1993 and 1994 for several plans. Since then, strategic pricing behavior by competing HMOs has led to changes in the plan that is the basis for the UC contribution, leading in turn to numerous changes in the out-of-pocket premiums facing employees. More details on this process and its impact on program spending have been reported previously (Buchmueller, 1998). What is important to note here is that the changes in employee contributions were driven largely by changes in the UC contribution. In several cases where an HMO became more expensive to employees between one year and the next, the total premium charged to the UC was either constant or even falling. Thus, it is reasonable to assume that the changes in the prices facing employees are uncorrelated with changes in plan quality, the nature of coverage offered or the characteristics of employees enrolled.[10]

Even if price changes are uncorrelated with these factors, we would expect the elasticity estimates to be sensitive to the inclusion of Prudential in the choice set. Since Prudential's share of enrollment was so low in the later years of our data, there is a limit to how much its enrollment could decline even in response to large price increases. As a result, including Prudential in the estimation sample can be expected to result in smaller estimated price effects.

[6] HMO enrollees were charged copayments of US$ 5 for physician visits and prescription drugs, and US$ 35 for emergency room visits.

[7] Prudential High Option has a US$ 200 deductible and a 10% coinsurance rate for a large network of providers and a 20% coinsurance rate for all other providers and for prescription drugs.

[8] Tier 1 has an HMO benefit design with copayments of US$ 10 per office visit. Patients can also opt to self-refer to a network of "tier 2" providers, and pay a US$ 250 deductible and 20% of charges thereafter. Tier 3 covers all other providers with a US$ 500 deductible and a 40% coinsurance rate.

[9] In 1993, other PPOs were offered at the Davis and Santa Barbara campuses instead of UC Care. Both plans were replaced by UC Care in 1994.

[10] The one possible exception to this is the case of Prudential, at least in the later years of our data. Between 1993 and 1995, Prudential's gross premiums remained constant but employee contributions for the plan increased as competition among the HMOs drove down the employer contribution. For these years, it seems reasonable to treat the change in Prudential's out-of-pocket price as exogenous. As Prudential's enrollment declined, its risk pool worsened, due to adverse retention. In 1996 and 1997, the plan responded by raising gross premiums, which caused the price to employees to increase as well. Because they were influenced by changes in enrollee characteristics, the exogeneity of these price changes may be questioned. This possibility led us to estimate a model using only data from the period 1993 through 1995 as summarized in Section 5.

Table 3
Differences in the distribution of plan enrollment for incumbent and newly hired UC employees, 1994 and 1996[a]

	Single coverage premium		Distribution of enrollment	
	Prior year	Current year	Incumbents	New hires
1994				
Prudential High Option	US$ 64.55	US$ 114.55	6.0%	1.3%*
UC Care	0	0	16.5	19.4*
Health Net	0	0	37.5	46.3*
PacifiCare	0	0	0.8	2.1*
Northern California Network HMOs	0	6.15	14.1	6.2*
Kaiser north	0	0	15.6	15.0
Kaiser south	0	0	9.5	9.7
Total (%)			100.0	100.0
1996				
Prudential High Option	US$ 133.55	US$ 273.12	1.7%	0.2%*
UC Care	3.97	9.39	22.7	23.2
Health Net	0	6.72	32.6	21.5*
PacifiCare	7.00	0	12.7	24.0*
Northern California Network HMOs	13.75	5.97	5.8	4.9*
Kaiser north	0	0	15.1	15.8
Kaiser south	0	0	9.4	10.5*
Total (%)			100.0	100.0

[a] The Northern California Network HMOs are Foundation and TakeCare (both years) and QualMed (1994 only). Premiums reported for these plans are enrollment-weighted average. New hires are workers with job tenure of less than 1 year.

* Difference between incumbents and new hires is statistically significant at the 0.01 level.

For this reason, in addition to estimating models for the full set of plans, we also analyze a sub-sample consisting of employees who chose one of the managed care plans (HMO, POS, or PPO) on the menu.[11] This sub-sample may offer more credible estimates of the "overall" elasticity that is relevant to settings where consumers are choosing from a menu of managed care plans. However, the full sample is more appropriate for investigating risk-related differences in price-sensitivity and making inferences about switching costs. Even before the change in the UC contribution policy, Prudential enrolled a disproportionate share of higher cost individuals and, independent of risk, switching from a less managed plan to an HMO may be more costly than switching among HMOs.

3.3. Descriptive evidence on persistence in plan choice

Descriptive statistics on enrollment suggest the importance of switching costs in health plan choice decisions. Table 3 shows the distribution of enrollment for recently hired UC employees, who were choosing their health plans for the first time, and incumbent

[11] We also conducted all analyses on a sample consisting of HMO enrollees. Since these results are quite similar to those for the managed care sample we do not report them.

employees, who were already enrolled in plans. We present data for 2 years, 1994 and 1996, for which differences between the two groups are especially instructive.

If switching costs are an important factor in health plan choice decisions, we would expect enrollment in plans that are new to a choice set to be greater among new hires who, by definition, cannot be locked into previous choices. The 1994 enrollment figures for PacifiCare, which was new to the UC program (and available only in southern California) that year, bear this out. The percentage choosing PacifiCare is greater for new hires than for incumbents: 2.1 versus 0.8%.

A more important question is whether persistence results in less elastic demand. Because 1994 was the first year after the change in the UC's contribution policy, Prudential and three network HMOs offered in northern California were more expensive to employees than they had been in the previous year. If switching costs translate to a lower price elasticity, we would expect these plans to have a greater market share among incumbent employees than among new employees. This is the case. In 1994, Prudential enrolled 6% of incumbents, but only 1.3% of new hires. [12] The percentage of incumbent employees in the HMOs requiring a monthly contribution was more than twice as large as the comparable percentage for new hires (14 versus 6%).

The data from 1996 provide more evidence of a negative relationship between switching costs and price sensitivity. From 1993 through 1995, Health Net was the only network model HMO available at all locations for a zero contribution, and by 1995 it had the highest share of UC enrollment of all the plans. In 1996, PacifiCare became available at all UC locations and undercut Health Net's premium to become the lowest cost plan in the program. PacifiCare gained significant market share with this strategy, largely at the expense of Health Net. While a significant number of employees switched from Health Net to PacifiCare, Health Net remained the most popular plan among incumbent employees with a much larger market share than PacifiCare (32.6 versus 12.7). New hires were more strongly influenced by the difference in price between these two plans, with a slight majority choosing PacifiCare.

The enrollment distribution for other years (not reported) reveals a similar pattern, which is consistent with the previous findings of Neipp and Zeckhauser (1985), and Royalty and Solomon (1999). Compared to incumbent employees, new hires are more likely to be in less costly plans and more likely to be in plans that are new to the menu.

4. Econometric model

4.1. Baseline model

The starting point for our econometric analysis is a conventional conditional logit model of health plan choice. This specification is similar to the models of most of the previous

[12] An obvious concern is that this result may be driven by differences in demographics between the two groups. For this reason, we also re-weighted the enrollment figures to control for differences in age, salary, and other observable characteristics using methods outlined by DiNardo et al. (1996). When the new hires are re-weighted to "look like" the incumbents, the percentage in Prudential increases, but only slightly. All the differences between the two groups remain statistically significant.

literature in that we constrain the effect of premiums to be the same for all individuals.[13] In addition to allowing clear comparisons with those prior studies, this specification is a useful benchmark for interpreting models that allow the effect of premiums to vary with key individual characteristics. This baseline model assumes that the probability of individual *i* choosing plan *j* is given by

$$\Pr(Y_i = j) = \frac{V_{ij}}{\sum_{k=1}^{J} V_{ik}} \tag{1}$$

where J is the number of options in the choice set. The utility individual i receives from plan j, V_{ij}, is given by

$$V_{ij} = \alpha P_{ij} + Z_j \gamma + X_i \beta_j + \varepsilon_{ij} \tag{2}$$

where, P_{ij} is the out-of-pocket premium that individual i would have to pay to enroll in plan j. The vector Z represents other plan attributes, including the number of years the plan had been offered at a given campus in a given year, which enters via three categories (1–2, 3–5 and more than 5 years on the menu). Because of the way benefits are standardized in the UC program, it is not possible to estimate the effect of benefit design on plan choice. To capture the effect of time-invariant plan attributes (e.g. reputation for customer service and other aspects of quality), Z includes a set of plan-specific intercepts.[14] With this specification, the effect of price is identified by intertemporal changes, differences across coverage tiers in required contributions, and interactions between the two. Thus, bias from omitted plan characteristics which may be a problem in other studies, should not be one here.[15]

The vector X represents individual-level characteristics that may be correlated with preferences and thus plan choices. Three variables that we hypothesize to be related to price sensitivity—age, job tenure, and health risk—enter the baseline model as controls. Age enters via three categories (30 and under, 31–45 and over 45) as does job tenure (new hire, job tenure of 1–5 years, and more than 5 years at the UC). Our health risk variable is dichotomous, based on whether or not the employee or his/her enrolled spouse was diagnosed with cancer or hospitalized around the time of open enrollment. The other individual-level controls in the model are six categories derived from interacting employee gender with coverage tier (single, two-party or family), three job categories (academic, management and

[13] One criticism of the conditional logit model is that it assumes the independence of irrelevant alternatives (IIA), which has led some researchers to use a nested logit model instead (see Feldman et al., 1989). However, with the nested logit model, IIA still holds within nests. Since most of the alternatives in the UC program would logically be grouped together, the nested logit model offers little improvement over a simpler model.

[14] Since the only variation in benefits is across plan type, an alternative, but more restrictive specification is to include dummy variables for each plan type (e.g. indemnity, PPO, POS and HMO), rather than for each specific plan. Such a specification yields estimated price effects that are essentially identical to those that we report. This indicates that, within a plan type, differences in plan characteristics are uncorrelated with differences in prices facing employees.

[15] This identification strategy assumes that there are no important "plan and time" effects that are correlated with changes in employee contributions. This is a reasonable assumption for the benefit program and time period we study. As noted, much of the variation in employee premiums is driven by changes in the UC contribution. In addition, a detailed inspection of HMO provider panels reveals that the composition of the panels was fairly stable over this period.

other staff), 11 indicator variables for employment location (e.g. campus or laboratory), an indicator for employees residing in largely rural counties, and the log of annual salary.

4.2. Allowing for a heterogeneous response to price

Our main interest is in models in which the premium variable is interacted with employee characteristics that proxy for differences in switching costs. We divide the sample into 18 mutually exclusive categories defined by crossing the three age categories with the three job tenure categories and the health status indicator based on inpatient utilization and cancer diagnoses. For this model, the utility associated with a given plan is expressed as

$$V_{ij} = \sum_{c=1}^{18}\{\alpha_c D_i^c P_{ij}\} + Z_j\gamma + X_i\beta_j + \varepsilon_{ij} \tag{3}$$

where the subscript c indexes the 18 distinct employee categories and the indicator variable D_i^c equals 1 if individual i is in group c, and zero otherwise. The variables in Z and X are the same as in Eq. (2).

The discussion of switching costs from Section 2 implies several clear predictions concerning differences across these categories in the effect of premiums. Since switching costs are likely to increase with medical care utilization, we expect the coefficient on premiums to decrease with age and, conditional on age, to be lower for individuals who we classify as high-risk based on the hospitalization and cancer data. Holding age and health status constant, we expect sensitivity to premiums to decline with the length of time an individual has been in the UC benefits program. We use job tenure rather than the time an individual has been enrolled in a particular plan because the latter is simply a lagged value of our dependent variable, making it endogenous. In addition to having stronger relationships with particular providers, we suspect that longer-term employees are more likely to exhibit a bias toward the default option during open enrollment. The estimated relationship between job tenure and price sensitivity combines both of these effects. Overall, we expect new hires in the youngest age category who do not meet our high-risk criteria to be most responsive to premiums and older, high-risk employees who have been at the UC for more than 5 years to be the least price-sensitive.

5. Estimation results

5.1. Baseline model

Table 4 presents selected results for the baseline model. [16] Results for the full sample and for the managed care sub-sample are reported in columns 2 and 3, respectively. In this specification our proxies for switching costs—age, job tenure and health risk—enter as control variables. The health risk variable we use is based on hospitalizations or cancer diagnoses

[16] For complete results for all the models described in this paper see Strombom (2000).

Table 4
Conditional logit regression results—baseline model[a]

	All plans	Managed care only
Selected coefficients		
Monthly premium	−0.0088 (0.0002)	−0.0241 (0.0083)
Plan offered 3–5 years	1.094 (0.037)	1.136 (0.037)
Plan offered >5 years	1.596 (0.042)	1.812 (0.044)
Implied price effects		
Loss of market share from a US$ 5 increase in premium		
Range (across plans) (%)	−1.2 to −3.7	−3.2 to −9.7
Enrollment-weighted average (%)	−3.2	−7.6
Insurer-perspective elasticity		
Range (across plans)	−0.84 to −5.18	−2.31 to −6.59
Enrollment-weighted average	−2.47	−5.27
Number of observations	103835	99692
Log-likelihood	−132943	−120338
Pseudo-R^2	0.270	0.233

[a] Standard errors in parentheses. The other explanatory variables are: natural log of annual UC salary plus categorical variables for age (30 and under, 31–45, over 45 years old), job tenure (new hire, 1–5 years, over 5 years), employee gender interacted with 3 coverage tier, 3 job classification categories, location of employment (11 categories) health status, residence in a rural area.

in the 34 months around open enrollment (22 months prior plus 12 months after).[17] For reasons of space, we do not report the estimated coefficients for the individual-specific variables; the most relevant results are summarized as follows.

Theoretical models by Cutler and Reber (1998), and Feldman and Dowd (2000) assume that individuals with greater expected utilization will have the strongest preference for the least tightly managed plans and the strongest aversion to integrated plans with closed provider panels. Our results support this hypothesis. We estimated the model treating UC Care as the omitted alternative against which the parameters for other plans are calibrated. Prudential High Option was the only option for which the coefficient on the high-risk indicator is positive and significant, indicating that high-risk individuals are more likely than low-risk individuals to prefer Prudential to all the managed care options. The corresponding coefficients were negative for all but one HMO, with the most negative being the one for Kaiser, which has the most limited provider panel. Similarly, older employees are most likely to prefer Prudential to UC Care and UC Care to any of the HMOs. The results also indicate that the probability of choosing the indemnity option is positively related to employee salary. The results for age and salary are similar to those of Cutler and Reber (1998).

[17] The results for the baseline model were not at all sensitive to alternative ways of defining health risk. Likewise, the pattern of results in the fully interacted model is not materially affected by the window used to identify high risk individuals. However, the use of smaller window results in a fairly small number of observations in certain cells. Thus, for the sake of greater precision, we report specifications that use the 34 month window around open enrollment to define health risk.

The single non-price attribute for which we can estimate an effect is the number of years a plan has been available at a particular location. Controlling for price, market share is higher for plans that have been on the menu for a longer period of time, which is consistent with the hypothesis that health insurance decisions are subject to persistence.

The coefficient on monthly premium is negative and statistically significant for both samples, though the magnitude differs considerably. In the full sample, the coefficient on the premium variable is −0.009; when we limit the sample to employees in managed care plans, the coefficient is −0.024. This difference can be explained by the lack of close substitutes for Prudential and by the possibility of a non-linear response to price.[18] Whatever the explanation, the difference between the two samples shows that estimates of health plan premium effects can be quite sensitive to the composition of the choice set and the range of premiums observed.

Since the logit coefficients are not directly meaningful, we present two alternative measures that help gauge the magnitude of the price effect. First, we simulate the loss of market share a plan would suffer if its cost to consumers were to increase by US$ 5 and the cost of all other plans remained constant. This is done by estimating for each individual the effect of the price change on the probability of choosing each plan and then averaging the differences for each plan across all individuals in the sample. Because of the non-linearity of the model, this simulation generates different effects for each plan. We calculate an "average" effect by weighting the result for each plan by its enrollment. Second, we calculate "insurer-perspective" demand elasticities. In the conditional logit model, a plan's own-price elasticity is a function of its price and market share:

$$\eta = \frac{\partial \ln \Pr(Y = j)}{\partial \ln P_j} = \alpha P_j \{1 - \Pr(Y = j)\}$$

where α is the coefficient on the price variable from Eq. (2) and P is the price at which the elasticity is calculated. The insurer-perspective elasticity is calculated using the plans' total premiums in this equation rather than the out-of-pocket premiums facing employees. Since elasticities will vary across plans depending on their premiums and market shares, we also calculate enrollment-weighted average elasticities.

Results for the simulations and the estimated elasticities are summarized in the middle panel of Table 4. The full sample model predicts that a premium increase of US$ 5 per month will reduce a plan's market share between 1.2 and 3.7%. For 9 of 11 plans the predicted reduction in market share is between 2.5 and 3.5%; the weighted average is a reduction of 3.2%. The estimated insurer-perspective elasticities for the full sample range from −0.8 to −5.2, though for 8 of the 11 plans the elasticity falls between −1.6 and −2.9. The average elasticity for the full sample is −2.5.

When we restrict the sample to individuals enrolled in managed care plans, the results imply a much stronger price effect: a US$ 5 increase in premiums is predicted to reduce

[18] In the full sample adding a quadratic term indicates a diminishing marginal effect of price and improves the fit of the model. In the managed care sample, however, including the quadratic term adds little explanatory power. We focus on the model that is linear in price to reduce the number of price parameters in our more complicated specification and to maximize comparability to the previous literature.

market share by between 3.2 and 9.7%, with an average decline of 7.6%. Likewise, the elasticities are larger in magnitude when we exclude Prudential, ranging from −2.3 to −6.6, with an average of −5.3.

This finding of a strong response to price is consistent with several recent studies. Dowd and Feldman (1994–1995) use plan-level data to estimate the effect of out-of-pocket premiums on plan market share. Their estimates imply that a plan that initially has 28% of the market and raises its premium by US$ 5 will see its market share fall by 11% points. This effect is nearly identical to the loss predicted in their earlier work using individual-level data (Feldman et al., 1989) and is quite similar to our average prediction from the managed care sub-sample. The range of elasticities we obtain from our full sample overlap considerably with those obtained by Royalty and Solomon (1999) using a conditional logit model. Our managed care only results are similar to ones they estimate using a fixed effect logit specification.

5.2. *Allowing for a heterogeneous response to price*

We now turn to the models that allow the effect of price to vary across 18 mutually exclusive employee groups, defined on the basis of age, job tenure and health risk. For each group, the full sample premium coefficient and the corresponding insurer-perspective elasticity are reported in Table 5.

The premium coefficient is negative and significant for all 18 categories. Further, the results indicate substantial heterogeneity in price sensitivity. Differences across groups are consistent with presumed differences in switching costs. As we hypothesized, young, newly hired employees who are classified as low-risk (columns 2 and 3) are the most price-sensitive, and high-risk employees in the oldest age category who have been at the UC for more than 5 years (row 12, columns 2 and 3) are the least sensitive to out-of-pocket premiums. There is a nearly seven-fold difference in the price coefficients for these two groups, −0.044 versus −0.007, and a greater than four-fold difference in insurer-perspective elasticities: −8.41 versus −2.03.

Between these two extremes, we find statistically significant marginal effects for age and job tenure. Controlling for job tenure and health status, the implied elasticity for employees in the youngest age category is roughly twice as large as the corresponding group in the oldest category. In all cases, the differences across age categories are significant at the 0.01 level. As suggested by the descriptive results, we find strong evidence that the elasticity of demand declines with job tenure, with the differences between new hires and employees with 1–5 years of tenure being more pronounced than the differences between the two categories of incumbents.

As in the baseline model, the results from the augmented specification indicate that individuals in the high-risk category are more likely than low-risk categories to prefer Prudential to the HMOs. However, controlling for job tenure and age, the differences in demand elasticity between the two groups are generally small. Most of these differences are not statistically significant.

Fig. 1, which presents the predicted loss of market share within each of the 18 cells in response to a US$ 5 increase in a health plan's monthly premium, offers an alternative perspective on our results. In addition to the fact that the patterns are easier to discern from a

Table 5
Conditional logit coefficients and elasticities for model allowing a heterogeneous response to price—all plans[a]

	Low risk		High risk	
	Number of observations	Results	Number of observations	Results
Age: 30 and younger				
New hire	6480	−0.0436 (0.0032) [−8.41]	235	−0.0302 (0.0083) [−6.72]
UC employee for 1–5 years	9580	−0.0200 (0.0018) [−4.20]	439	−0.0183 (0.0040) [−4.68]
UC employee for >5 years	997	−0.0141 (0.0022) [−3.42]	62	−0.0137 (0.0059) [−3.82]
Age: 31–45				
New hire	8148	−0.0236 (0.0016) [−6.21]	518	−0.0179 (0.0027) [−5.54]
UC employee for 1–5 years	19842	−0.0130 (0.0006) [−3.68]	1743	−0.0105 (0.0010) [−3.30]
UC employee for >5 years	18216	−0.0100 (0.0004) [−3.12]	2015	−0.0096 (0.0008) [−3.28]
Age: 46 and older				
New hire	2254	−0.0176 (0.0016) [−4.69]	306	−0.0108 (0.0019) [−3.23]
UC employee for 1–5 years	6896	−0.0094 (0.0006) [−2.66]	1225	−0.0073 (0.0008) [−2.22]
UC employee for >5 years	21080	−0.0069 (0.0003) [−2.02]	3799	−0.0065 (0.0004) [−2.03]

[a] Standard errors in parentheses. Insurer-perspective elasticities in square brackets. Elasticities are enrollment-weighted average over all plans. An observation is classified as high-risk if the employee or spouse was diagnosed with cancer or had a non-maternity hospitalization in the period from 22 months prior through 12 months after open enrollment. Other explanatory variables are: years the plan has been on the menu (same categories as in Table 4), the natural log of annual salary, employee gender interacted with three coverage tier categories, three job classification categories, location of employment (11 categories), and residence in a rural area. Log-likelihood $= -132{,}762$; pseudo-$R^2 = 0.271$.

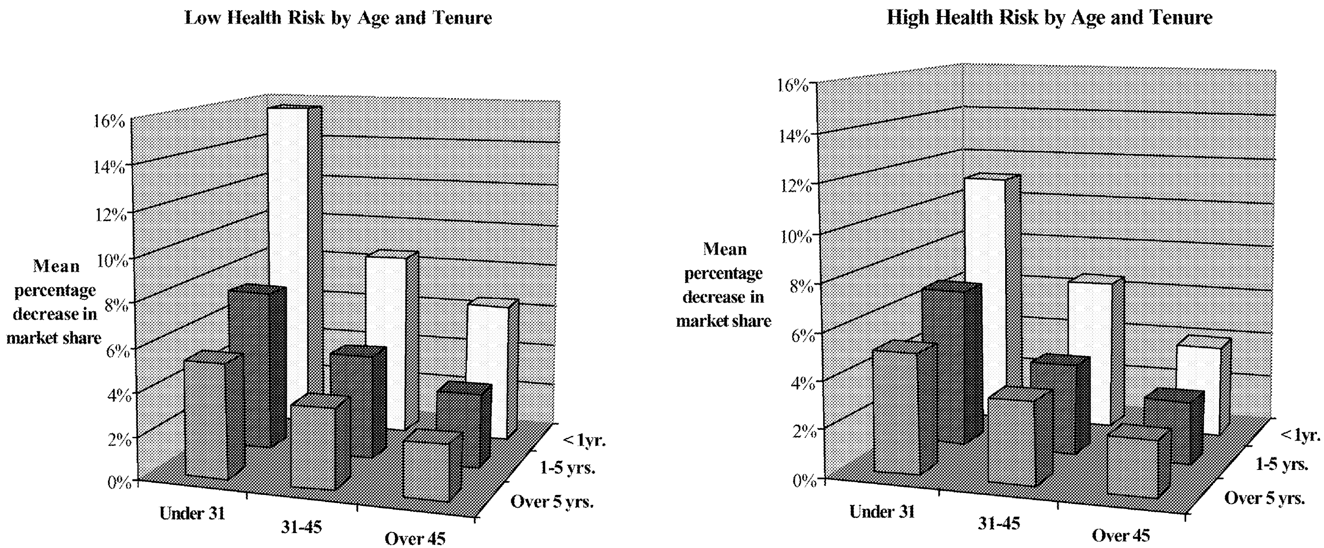

Fig. 1. Simulated loss of market share caused by a US$ 5 increase in monthly premiums, all plans. Simulations are based on estimates from the regression model reported in Table 5. For each employee category the loss of market share represents an enrollment-weighted average over all plans.

graphical presentation, these simulations have conceptual advantages relative to the results reported in Table 5. In a non-linear model where a variable such as price is interacted with individual characteristics, the differences in the effect of that variable will not generally be proportional to the magnitude of the interaction coefficients. In addition, as is well known, elasticity calculations are sensitive to the values at which the calculations are performed. The simulated loss of market share avoids both of these problems.

The results in Fig. 1 show that a US$ 5 increase in premiums is predicted to lead to a 15% reduction in market share among young, newly hired employees in the low-risk category and approximately a 11% reduction for young, new hires in the high-risk category (the back left corner of each figure). In contrast, roughly 2% of older workers who have been at the UC for more than 5 years are predicted to switch plans (the front right corner of each figure).

Table 6 and Fig. 2 present results for the managed care sub-sample. Differences across cells are somewhat less pronounced than in the full sample, though the overall pattern is quite similar. The price coefficient for the low-risk, young, new hires is four times as large as the coefficient for high risk, older, long-term employees (−0.077 versus −0.019). As in the full sample, controlling for health risk and job tenure, price sensitivity declines with age, though somewhat more gradually. As in the full sample, new hires are significantly more price sensitive than incumbents, but among incumbents price sensitivity does not decline with tenure. This result suggests that tenure effects may have more to do with status quo bias in open enrollment decisions than with employees becoming increasingly locked into to their choices over time as they develop ties to particular providers. Controlling for age and job tenure, the results from the managed care sample offer no evidence that poor health reduces price sensitivity.

The differences between the results for the full sample and managed care sample suggest that biased risk selection caused by differences in price sensitivity will tend to be a greater problem in markets where HMOs and non-HMOs compete directly and less of a problem for competition among HMOs. This inference is consistent with the experience of the Health Insurance Plan of California (HIPC), a state-sponsored purchasing cooperative designed according to the managed competition model. Several PPOs that were initially on the HIPC menu experienced adverse selection and ultimately withdrew from the cooperative. In contrast, during the first 5 years of the HIPC's existence, there was no evidence of biased selection among competing HMOs (Yegian et al., 2000).

5.3. Sensitivity tests

To test the sensitivity of our results, we considered several alternative specifications. First, we considered different approaches to constructing the health risk variable. The effect of using alternative measures is illustrated most concisely with a simpler model in which only the health variable is interacted with the premium. In this model, which is similar to the one used by Royalty and Solomon (1999), the coefficient on the interaction term represents the overall difference in the effect of price that is related to health risk, averaged over all age and tenure categories.

Key results are reported in Table 7. The figures in the first column are for the risk variable used throughout the paper so far. The coefficient on out-of-pocket premiums is 16% smaller in absolute value for employees who are defined as high-risk according to these criteria than

Table 6
Conditional logit coefficients and elasticities for model allowing a heterogeneous response to price—managed care plans only[a]

	Low risk		High risk	
	Number of observations	Results	Number of observations	Results
Age: 30 and younger				
New hire	6460	−0.0771 (0.0045) [−11.12]	233	−0.0715 (0.0173) [−12.02]
UC employee for 1–5 years	9495	−0.0271 (0.0032) [−4.31]	433	−0.0350 (0.0107) [−6.90]
UC employee for >5 years	978	−0.0359 (0.0074) [−6.74]	61	0.0000 (0.0194) [0.00]
Age: 31–45				
New hire	8070	−0.0525 (0.0026) [−10.53]	511	−0.0491 (0.0071) [11.69]
UC employee for 1–5 years	19345	−0.0214 (0.0015) [−4.66]	1674	−0.0208 (0.0040) [−5.15]
UC employee for >5 years	17466	−0.0211 (0.0014) [−5.21]	1907	−0.0272 (0.0033) [−7.42]
Age: 46 and older				
New hire	2228	−0.0447 (0.0042) [−9.00]	297	−0.0369 (0.0091) [−8.40]
UC employee for 1–5 years	6605	−0.0162 (0.0024) [−3.51]	1154	−0.0122 (0.0046) [−2.90]
UC employee for >5 years	19350	−0.0176 (0.0014) [−4.02]	3425	−0.0188 (0.0028) [−4.64]

[a] Standard errors in parentheses. Insurer-perspective elasticities in square brackets. Elasticities are enrollment-weighted average over all plans. An observation is classified as high-risk if the employee or spouse was diagnosed with cancer or had a non-maternity hospitalization in the period from 22 months prior through 12 months after open enrollment. Other explanatory variables are: years the plan has been on the menu (same categories as in Table 4), the natural log of annual salary, employee gender interacted with 3 coverage tier categories, 3 job classification categories, location of employment (11 categories), and residence in a rural area. Log-likelihood $= -120{,}181$; pseudo-$R^2 = 0.234$.

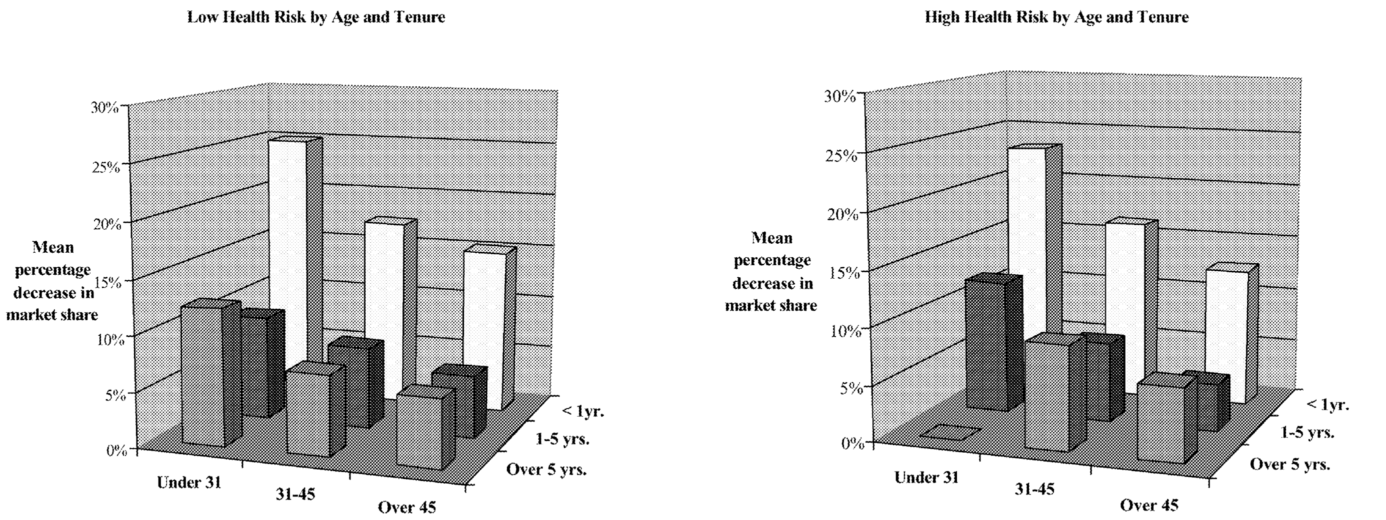

Fig. 2. Simulated loss of market share caused by a US$ 5 increase in monthly premiums, managed care plans only. Simulations are based on estimates from the regression model reported in Table 6. For each employee category the loss of market share represents an enrollment-weighted average over all plans.

Table 7
Health-related differences in the effect of premiums using alternative health status measures[a]

	Observations with health event (%)	Coefficient on premium		Difference (%)	P-value for difference
		No health event	Health event		
Cancer or non-maternity hospitalization 22 months prior, 12 months after OE	10.0	−0.0090	−0.0076	−16	0.004
Cancer or non-maternity hospitalization 22 months prior to OE	6.7	−0.0090	−0.0076	−16	0.015
Cancer or non-maternity hospitalization 12 months after OE	4.2	−0.0089	−0.0073	−18	0.000
Non-maternity hospitalization 12 months after OE	4.0	−0.0089	−0.0070	−21	0.003
Cancer diagnosis 22 months prior to OE	0.7	−0.0089	−0.0052	−42	0.000
Maternity hospitalization 22 months prior to OE	4.9	−0.0087	−0.0135	55	0.000

[a] Results are for models estimated on the full sample. Except for the interaction of the premium and health status variables, the specification is the same as in Table 4.

for those defined as low-risk: −0.0076 versus −0.0090. The results are identical when we limit the window to the 22 months prior to open enrollment, and virtually identical when the window is shifted to the 12 months following open enrollment.[19] When we define health risk based on cancer alone, the price coefficient for high-risk employees is 42% smaller in absolute value than that for low-risk employees. That the differential is greater when we limit our high-risk category to cancer patients is not surprising given that this group is more homogeneous than those defined based on being hospitalized. The one health risk variable that stands out from the rest is the one defined on the basis of a maternity-related hospitalization. The price coefficient for households in which there was a maternity-related hospitalization in the 22 months preceding open enrollment is actually larger in absolute value than that for all other households.

As noted, differences in the results for samples that include and exclude Prudential suggest that the effect of out-of-pocket premiums on plan choices may be non-linear. To test the importance of this possibility to our findings, we estimated models that included (1) a quadratic price term; (2) a "free" plan indicator equal to one for plans that require no out-of-pocket contribution by enrollees, and (3) both a quadratic term and a free plan indicator. Each of these specifications significantly improved the fit of the model over the linear specification for the sample that included Prudential, but provided only marginal improvement in fit for the managed care only sample. The overall pattern of price sensitivity across the 18 categories, as indicated by the change in market share from a US$ 5 increase in premium, was similar to that reported above for each of these three alternative specifications.

A final set of sensitivity tests involved using alternative estimation samples. Because prices vary by coverage tier and because individuals with employed spouses may have alternative sources of coverage, we estimated separate models for employees choosing single coverage and those insuring dependents. The results for the two sub-samples were qualitatively the same. Also, as noted, we estimated models for the period 1993 through 1995 in order to gauge the effect of the arguably endogenous price spiral of Prudential High Option in 1996 and 1997. As expected, eliminating these years increases the estimated price effect by some 50%. This is consistent with our results from the managed care only sample, and underscores the sensitivity of estimated price effects to levels and variation in prices in the particular study period. Finally, we also estimated models on a sample that excludes both UC Care and Prudential. The results for this HMO only sub-sample are comparable to the results presented in Table 6 for the managed care sub-sample.

6. Implications

6.1. Risk selection within markets

The correlation between health risk and price sensitivity implies that changes in relative prices will lead to changes in plans' expected costs arising from changes in the distribution of

[19] While these findings, along with our findings concerning the effect of age and tenure on price sensitivity, are consistent with the results of Royalty and Solomon (1999), a direct comparison of the magnitudes of the effects we find with this earlier study is not possible since they report only consumer-perspective price elasticities for the models in which price is interacted with personal characteristics.

risk. To gain some perspective on the magnitude of such effects, we combine our regression results with actuarial cost data to simulate the effect of a change in relative premiums on the risk pools and expected medical costs of health plans in a hypothetical market. For each individual in our sample, we imputed expected medical costs using a Diagnostic Cost Group (DCG) model calibrated using data from a national sample of commercial health plan enrollees.[20] Individual-level expected medical costs were aggregated to the level of our 18 age/tenure/health risk categories,[21] which we use to estimate how plans' average costs change in response to changes in enrollee mix.

Our simulation results are summarized in Table 8. The top panel describes a baseline case in which we assume consumers are distributed equally across four competing health plans, with characteristics reflecting the actual distribution of all UC enrollees. The other panels show how enrollee characteristics and expected costs change for a plan that increases in price and for its average competitor, for which we assume premiums remain constant. For each estimation sample we simulate the effect of premium increases of US $5 and US $25 per month.

Using the parameter estimates from the full sample regressions, a plan whose premium increases by US $5 would see its market share decline by 5%, from 25.0 to 23.7%. The mean age of enrollees would increase by roughly 1%, from 38.4 to 38.6, and the percentage of enrollees falling into our high-risk category would increase by 2%, from 9.96 to 10.13. These changes in enrollee characteristics would cause the plan's expected claims expenses to increase by US $2.12 per household per month. A larger premium increase would clearly result in a greater loss of market share and a further deterioration of the risk pool. The results for the full sample imply that a plan increasing its premium by US $25 (relative to its competitors) would lose roughly one-quarter of its initial enrollment and the expected claims cost for those remaining in the plan would increase by US $10.36, or 3.9% of the baseline average.

The other plans that held their premiums constant would experience a slight decline in average expected claims costs, due to favorable selection. In response to a US $25 increase in premiums by its competitor, the three other plans would experience a decline in expected claims of US $2.47 per household per month. As a result, expected claims for these plans would be 5% lower than for the plan that increased its price.

In Table 4 we saw that our estimation results based on the sub-sample of UC employees in managed care plans implied larger premium elasticities. The results in Table 8 indicate that these greater elasticities do not imply greater change in the distribution of risk. For example, as shown in the fourth column of Table 8, based upon the sample of managed care plans only, a US $5 increase in premiums reduces a plan's expected market share by 9.8% and increases its expected benefit cost by US $2.27. By comparison, the results from the full sample shown in the second column indicate an expected reduction of market share of 5.2%, and an increase in expected benefits cost of US $2.12.

[20] The model predicts medical costs (including pharmacy costs) based upon the sex, age and principal inpatient diagnosis, if any, in the year of enrollment and preceding year. For a complete description of the model and calibrating data see Ellis et al. (1996). Our thanks to DxCG Inc. who supplied the model.

[21] Average expected medical costs ranged from US$ 157 per household per month for those 30 and under with no hospitalizations, to US$ 893 for those over 45 with at least one hospitalization in our reference period.

Table 8
Simulation analysis: the effect of a premium increase on plan market share and expected claims costs[a]

	All plans (US$)		Managed care plans only (US$)	
	5	25	5	25
Premium increase				
Baseline				
Market share (%)	25.0	25.0	25.0	25.0
Mean age of enrollees	38.4	38.4	38.4	38.4
Percent of enrollees in high-risk category (%)	9.96	9.96	9.96	9.96
Expected monthly claims costs per household (US$)	268.83	268.83	268.83	268.83
After price change—plan raising premium				
Market share (%)	23.7	19.3	22.7	15.7
Mean age of enrollees	38.6	39.4	38.6	39.3
Percent of enrollees in high-risk category (%)	10.13	10.81	10.10	10.35
Expected monthly claims costs per household (change relative to baseline, US$)	270.95 (+2.12)	279.19 (+10.36)	271.10 (+2.27)	277.00 (+8.17)
After price change—other plans				
Market share (%)	25.4	26.9	25.8	28.1
Mean age of enrollees	38.3	38.1	38.3	38.2
Percent of enrollees in high-risk category (%)	9.91	9.76	9.92	9.89
Expected monthly claims costs per household (change relative to baseline, US$)	268.17 (−0.66)	266.36 (−2.47)	268.16 (−0.67)	267.32 (−1.51)

[a] The baseline case assumes that enrollees are distributed identically across the four plans. Enrollee characteristics are based on the actual distribution of characteristics for all UC employees. Simulated changes in enrollee characteristics are based on the models summarized in Table 5 and Fig. 1 (all plans) and Table 6 and Fig. 2 (managed care plans). Average monthly claims costs for each plans are calculated by assigning cost weights derived from a Diagnostic Cost Group (DCG) model as described in the text.

These cost figures are meant to be illustrative and several caveats regarding them should be noted. First, because our health risk variable captures only one aspect of enrollee health status, our results do not reflect the full effect of a change in relative prices on the risk pool. Our simulations assume that both price sensitivity and expected costs are uniform within each of our 18 categories and ignore the effect of chronic health conditions not requiring hospitalization and the health status of dependent children. To the extent that these unmeasured factors lead to additional selection effects within categories, our results understate the degree to which the distribution of risks will change in response to changes in relative prices.

Second, the starting point for the simulations—four plans with equal and identical enrollment—is artificial in a way that produces more stability than would likely be observed in a real market. In particular, this set-up does not account for the fact that high-risk consumers are not only less price-sensitive, but have a strong preference for less managed plans that are inherently more costly. Third, the simulations are static and do not account for strategic interactions among health plans or other ways that plans may respond to changes in their risk pool (i.e. changes in benefit design).

For these reasons, our simulations probably understate the extent to which certain types of plans are susceptible to adverse selection and do not give a full sense of how that dynamic plays out over time.[22] Indeed, the actual experience of Prudential suggests that this is the case. In 1993, the share of Prudential enrollees hospitalized for a non-maternity diagnosis during the enrollment year was 43% higher than the corresponding figure for all other plans (6.4 versus 4.5%). After the change in the UC contribution policy caused an exogenous change in the price of Prudential to consumers, this differential increased to 55% in 1994 and 78% in 1995. By 1997, the percent hospitalized in Prudential was over twice as great as for the other plans (9.9 versus 4.5%). Similarly, the average age of employees enrolled in Prudential increased by 7.4 years, compared to an increase of only 1.9 years for all other plans.

The time path of Prudential's adverse selection spiral is very similar to that of the single PPO in the Harvard University health benefits program following a similar change in contribution policy (see Cutler and Reber, 1998). In both cases, as the least managed plan on the menu became more expensive it not only lost significant enrollment, but experienced an increase in costs as the members who remained in the plan were older and sicker than individuals who left. Premium increases intended to cover these higher costs led to further declines in enrollment and worsening of the risk pool. In both programs, the death spiral was completed within 3 years.[23]

6.2. *Differences across markets in average demand conditions*

Another implication of our results pertains to differences across markets in the demand conditions facing health plans. In particular, our results are pertinent to recent proposals

[22] Forecasting longer term dynamics would also require modeling the flow of employees in and out of the health benefits program, something that is beyond the scope of this paper.

[23] Another striking similarity between the two programs is that in both cases HMOs responded to the new contribution scheme by lowering their premiums. See Buchmueller (1998), and Cutler and Reber (1998) for more details.

to redesign the Medicare program with a greater emphasis on market competition and consumer choice. Advocates of such market-oriented reform proposals typically point to the experience of large public and private employers, where employees have proven to be quite price-sensitive and competitive strategies appear to have been successful in controlling health care spending (see, e.g. Butler and Moffit, 1995). However, it is not clear how well these results generalize to Medicare.

Our results suggest reasons why Medicare beneficiaries are likely to be significantly less price-sensitive than non-elderly workers. Not only are Medicare beneficiaries older and in poorer health, but also there is less turnover among the enrolled population than in the typical employer–sponsored program. As a result, health plans competing within the Medicare program are likely to face a less elastic demand than in commercial markets.

In thinking about how our results generalize to other settings, it is important to note that several features of the UC program have the effect of minimizing switching costs. In particular, during the period of our analysis, most UC employees were in HMOs and there was a high degree of overlap in the provider panels of competing plans. As a result, switching among HMOs did not generally require switching physicians. In markets where the provider panels of competing plans are more distinct, we would expect a lower level of price-sensitivity overall and a stronger gradient with respect to age and health status.

7. Conclusions

This study adds to the growing body of evidence that price is an important determinant of employee health insurance choice decisions. Our finding of substantial average price elasticities among employees of the University of California provide support for the argument that requiring consumers to bear the marginal cost of their health insurance decisions can significantly alter their behavior and, in turn, the incentives facing health insurance plans.

More importantly, this study extends the literature by investigating differences in price sensitivity within an insured population. We find strong evidence that price-sensitivity declines with age. Controlling for health status and job tenure, elasticities for employees age 30 and under are roughly twice the magnitude of elasticities for employees over age 45. We also find that newly hired employees are substantially more sensitive to price than incumbent employees, for whom the default option is to remain in a previously chosen plan. It is not possible to determine whether this result is due to "status quo bias", or if individuals become increasingly locked in to their chosen plan as their relationships with particular providers becomes stronger over time. However, the fact that we find little relationship between job tenure and price sensitivity among incumbent employees casts doubt on the latter explanation. It seems more likely that many individuals pay closer attention to enrollment materials when they first sign up for coverage than they do later.

Consistent with prior studies, we also find that high-risk employees have a stronger preference for less tightly managed plans. We find weak evidence of an independent effect of health status on price-sensitivity after controlling for age and job tenure. In our full sample, we find smaller price effects for individuals who were diagnosed with cancer or hospitalized in the months around open enrollment, but the differences between high- and low-risk employees are not generally significant. This lack of significance may be due to noise in our measure of risk. When we estimate simpler models in which only health status is interacted with plan premiums, risk-related differences are largest when risk is defined solely on the basis of being diagnosed with cancer.

The finding that individuals who are more costly to insure—older employees and those who have recently experienced serious health events—are less willing to switch health plans in response to a change in relative premiums than are younger, healthier employees means that plans that increase in price relative to competitors will not only lose market share, but will experience an increase in costs, due to adverse retention. The fact that, controlling for price, high-risk individuals tend to have a stronger preference for less tightly managed plans suggests that such plans are at a significant competitive disadvantage in environments where there is a menu of plans and differences in gross premiums are translated into differences in the prices facing consumers. Indeed, in the data we analyze, the adoption of a fixed dollar contribution policy pushed the single indemnity option on the menu into a rapid adverse selection death spiral.

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