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ABSTRACT

This study examines the effect of price on the demand for health insurance by early retirees between the ages of 55 and 64. The analysis is based on administrative data from a medium sized employer and takes advantage of a natural experiment created by the firm's health insurance contribution policy. The amount the firm contributes toward retiree health insurance coverage depends on when a person retired and her years of service at that date. As a result of this policy, there is considerable variation in out-of-pocket premiums faced by individuals in the data, but this variation is independent of the non-price attributes of the health insurance plans offered, and plausibly exogenous to individual characteristics that are likely to affect the demand for insurance. We find that price has a statistically significant but small effect on the decision to take up coverage. The implied elasticities are very similar to results found in previous studies using very different data.

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1 Introduction

There is considerable concern among US policy makers about the insurance coverage of "near-elderly" adults, i.e., those between the ages of 55 and 64. Because attachment to the labor force weakens as individuals approach the normal retirement age of 65, individuals in this age group have lower rates of employer-provided health insurance than younger adults (Monheit, Vistnes and Eisenberg 2001). Many older workers who lose their jobs are unable to find new jobs that provide health insurance, while others withdraw from the labor market altogether. Whereas many early retirees could once continue to receive insurance through their former employer, in the past decade there has been a dramatic decline in the percentage of employers offering such coverage (McCormack et al. 2002; Weller, Wenger and Gould 2004). Many firms that continue to provide retiree health benefits have substantially increased the cost to early retirees (US General Accounting Office 2001; Neuman 2004), and a growing number of firms offer "access only" plans, where the employer requires retirees to contribute the full premium. These trends are likely to continue. Retiree health benefits have also been eliminated as part of several high profile bankruptcy proceedings (for example, Polaroid, Bethlehem Steel). According to a 2003 survey of private sector employers, 10% of firms offering retiree health benefits have eliminated coverage for future retirees, and an additional 20% of firms are considering doing so (McArdle et al. 2004).

In light of these trends, recent policy proposals aimed at increasing insurance coverage have directly targeted the near-elderly. In each of his last three State of the Union Addresses, President Clinton proposed allowing the near-elderly to buy into Medicare at actuarially fair prices (Short, Shea and Powell 2001a). In 2000 and 2004, Democratic Presidential candidates Al Gore and John Kerry proposed a subsidized buy-in for 55 to 64 year olds. Congressional Democrats have proposed similar policies. The preferred strategy among Republicans for expanding health insurance relies on tax credits for non-group coverage. While this approach is not specifically targeted at particular age groups, the impact of a tax credit policy is likely to be most pronounced on the near-elderly as they tend to have a stronger demand for insurance and are more likely to rely on the non-group market than younger consumers.

In order to better understand the implications of the decline in employer payments for retiree health insurance and to evaluate these policy proposals, it is necessary to have good estimates of the price elasticity of health insurance demand for this segment of the population. In this paper we use data from an employer-sponsored retiree health insurance program to estimate the effect of out-of-pocket premiums on the insurance take-up decisions of early retirees between the ages of 55 and 64. Like many employers, this firm altered its retiree health benefits program in the mid-1990s in an attempt to control spending. This policy change created an excellent natural experiment for estimating the effect of price on early retiree health insurance decisions. Under the current system, the amount the firm contributes towards the insurance coverage of retirees depends on when a person retired and how many years she had been with the company. Specifically, for individuals who retired after January 1993, the employer's premium contribution depends on how long the person had been employed at the firm. Thus, two otherwise similar individuals who retired at different points in time–i.e., before or after January 1993– face very different prices. Similarly, for post-1993 retirees, prices also differ for two people who retired at the same time but with different years of service. This variation is ideal for identifying the effect of price on the demand for insurance since it is independent of any features of the plans offered (i.e., benefit generosity or the perceived quality of affiliated providers) and individual characteristics that are likely to be related to the demand for insurance.

We use these data to estimate probit regression models of the decision by early retirees to take up health insurance coverage offered by the firm. The regression results indicate a small but statistically significant effect of price on the take-up decision. The implied price elasticities range from -0.10 to -0.16. Our results are robust to various model specifications and sample definitions and are in the range of previous studies that use different data and estimation strategies. We use these regression results to simulate the effect of policy proposals for a Medicare buy-in and a non-group tax credit on coverage rates. Because the estimates imply that the take-up decision is fairly price inelastic, the simulations indicate small effects on take-up by near elderly retirees.

The paper is organized as follows. The next section summarizes several previous studies that also estimate take-up elasticities using data on workers offered coverage by their employers. Section 3 describes our data, presents descriptive evidence on the relationship between price and the take-up decision, and lays out our econometric strategy. We present our regression results in Section 4. Based on these regressions, results from simulations are reported in Section 5. The final section discusses limitations of the analysis and identifies possible directions for future research.

2 Previous Literature

The estimation of premium elasticities requires good data on the insurance options available to individuals and the prices charged for them. Population surveys that are commonly used to study insurance coverage lack this information. Thus, most research on the elasticity of demand for health insurance uses data on employees who are offered insurance by their employer but are required to contribute toward that coverage. The strengths and weaknesses of these studies reflect the advantages and limitations of each data source.

Chernew, Frick and McLaughlin (1997) use data from a survey of small employers in seven cities to model employee take-up as a function of out-of-pocket premiums. Focusing on single, lower income workers, they find a small but statistically significant effect of price. Their results imply take-up elasticities ranging from -0.03 to -0.095. Blumberg, Nichols and Banthin (2001) take a similar approach, using data from the Medical Expenditure Panel Survey (MEPS). While the MEPS is a nationally representative survey, in order to use information on the out-of-pocket premiums faced by employees, Blumberg, Nichols and Banthin use a special sub-sample of the data. For workers with dependents, their estimated take-up elasticities range from -0.03 to -0.08, depending on the econometric specification. Estimated price effects for single workers are smaller.

The main limitation of both of these studies is that the variation in price comes entirely from differences across employers. As a result, the results may be biased by unobserved heterogeneity. The direction of this bias is unclear. On one hand, for firms that set employee contributions as a fixed dollar amount or a fixed percentage of total premiums, plans that have a higher actuarial value (and are therefore more attractive) will be more expensive to employees. This will cause the price effect to be biased toward zero. On the other hand, if firms that pay higher compensation in general offer better health benefits and subsidize them more fully than firms that pay less overall, out-of-pocket premiums will be negatively correlated with plan quality, causing the partial effect of price to be overstated.

Two other studies each use data from a single employer. Gruber and Washington (2005) analyze a natural experiment caused by a change in the tax treatment of employee premium contributions that affected some federal workers and not others. Using group-level data, they exploit cross-sectional and intertemporal variation in the after-tax cost of insurance generated by this policy change and changes in marginal tax rates. They estimate a take-up elasticity of -0.02. Royalty and Hagens (2005) analyze a real experiment conducted by a large employer as part of an effort to redesign its fringe benefit offerings. Participating employees were asked to choose from a menu of hypothetical benefits, including health, dental and long term care insurance as well as vision and wellness benefits. For each type of benefit there were several alternatives including the option of declining coverage. The prices of the various options were manipulated in order to estimate the impact of price on employee choices. For all benefits, price is found to have a small negative impact on the decision to take up any coverage, though for health insurance the effect was not statistically significant.

An important advantage of the studies by Gruber and Washington (2005) and Royalty and Hagens (2005) is that they exploit within-plan variation in prices that is independent of other attributes of the health insurance offered by employers and plausibly exogenous to characteristics of employees that affect the demand for insurance. However, each study has its own shortcomings. Since Gruber and Washington's price variable is a function of the employee's marginal tax rate, which is not directly observed, they must impute marginal tax rates from other sources. The accuracy of this imputation and its impact on the results are not clear. The main limitation of Royalty and Hagens' analysis is that it is based on hypothetical, rather than actual, choices.

Our research design is similar to these two studies in that we use data from a single employer, and our identification strategy is based on within-plan variation in price that is plausibly exogenous. As in the Gruber and Washington study, the variation is driven both by rules that generate cross-sectional price differences across different classes of individuals, as well as from changes in prices over time. The main advantage of our data relative to theirs is that the price variable is observed directly in the data and measured without error.

The main advantage compared to the data used by Royalty and Hagens is that we analyze actual, rather than hypothetical, choices. A final feature that distinguishes our analysis from all the others in this literature is our focus on near-elderly retirees who, as noted in the introduction, are an important population from a policy perspective.

3 Data and Methods

3.1 Data Source and Sample Construction

The administrative data we use come from a medium-sized employer (roughly 2,700 employees) located in the Southwestern US and pertain to the health insurance choices made by early retirees from 1998 to 2003. Since we are interested in estimating take-up elasticities that are relevant to Medicare buy-in proposals, we focus on retirees between the ages of 55 and 64. In order to minimize the impact of unobserved heterogeneity, we limit the analysis sample to people who retired after 1990. The main reason is that individuals who retired in the 1980s and are still under the age of 65 by the late 1990s must have retired at a very young age. In some cases, the reason may have been a serious health problem; others may have taken early retirement only to start a second career elsewhere. In either case, there is reason to think that they are quite different than the average retiree in this age group.

Because we use multiple years of data, individuals can contribute between one and six observation to the sample. Overall, we have a sample size of 1,760 observations on 510 individuals.

3.2 Health Insurance Options and Prices

The firm offers four different health insurance options: two Health Maintenance Organizations (HMOs), one Preferred Provider Organization (PPO), and a cash payment for declining coverage. The employer contribution toward coverage is less than the full premium for the least costly plan; retirees are required to pay the difference between the employer contribution and the full premium for their chosen plan. For all plans, a higher contribution is required for two-party and family coverage than for single coverage.

The exact amount the employer contributes depends on whether an individual retired before or after January 1, 1993. Pre-1993 retirees receive a more generous subsidy. In 1998, they were charged just over $5 per month for single HMO coverage and $12 per month for single coverage under the PPO. Out-of-pocket prices increased during the period analyzed, especially for the PPO. By 2003, pre-1993 retirees were charged $53 per month for one HMO, $57 for the other, and $206 for the PPO. Throughout the period, the payment for declining coverage was constant at $75 per month for this group.

For individuals who retired after January 1 1993, out-of-pocket premiums depend on the person's years of service at the time of retirement. Those who had worked for the company for at least 25 years face the same prices as pre-1993 retirees. Retiree contributions increase by a fixed percentage for each year of service less than 25.[1] So, for example, in 2003, the monthly out-of-pocket cost for the cheaper HMO was $53 for a post-1993 retiree with 25 years of service, $69 for a post-1993 retiree with 20 years of service, and $130 for someone who retired with 10 years of service. For post-1993 retirees with less than 25 years of service the payment for declining coverage increases by $3 for every year of service.

This price variation is the greatest strength of these data. Most of our observations are on post-1993 retirees. There is little reason to think that someone who started working for the company 20 years prior to when we observe him should have a weaker or stronger demand for health insurance than, say, someone who had started with the company 25 years earlier. It is, perhaps, less obvious that the price differences between individuals who retired just before and after January 1993 are uncorrelated with the demand for insurance. In principle, someone with a very strong demand for coverage may have retired just before that date to lock into lower premiums. However, such strategic behavior is not a factor in these data because the employer's policy was determined retroactively. So, even if some employees might have been inclined to retire earlier to take advantage of a more generous subsidy, this was not possible. Indeed, an examination of the pattern of retirements over time suggests that the company may have chosen the January 1993 cut-off because a large number of workers retired in early 1993 (see Appendix Figure A-1).

[1]For individuals retiring after 1993 with less than 25 years of service, the out-of-pocket premium is defined by $P = F - C(1 - 0.04(25\text{-}s))$, where F is the full premium (the insurer price charged to the employer), C is the employer contribution for pre-1993 retirees, and s is the individual's years of service at retirement.

3.3 Descriptive Statistics

Table 1 presents summary statistics for our data. The mean of the dependent variable matches up closely with published sources and prior studies. For example, according to Monheit, Vistnes and Eisenberg (2001), in 1996 82% of US workers between the ages of 55 and 64 who were offered employer-sponsored health insurance took up that coverage. The take-up rate in our data is slightly higher than this figure (86.7%), but is essentially identical to the take-up rates in the studies by Gruber and Washington's (86.8%) and Blumberg, Nichols and Banthin (86.4%).[2]

Since the firm offers a choice of health insurance options, retirees face a schedule of prices, corresponding to the different plans available and different coverage tiers (i.e., single, two-party or family coverage). Following Blumberg, Nichols and Banthin (2001), we use the price for the least costly option available, which corresponds to a single coverage option for all individuals. For most observations in our data, the lowest cost option is single coverage through one of the HMOs. In the full sample, the mean for the lowest premium available is $59.81 per month .[3] However, this price does not represent the full cost of taking up coverage, as it does not account for the fact that individuals who take coverage forego a cash payment of up to $75 per month. Thus, the true cost of coverage is the lowest premium plus this foregone payment. The mean for this variable, which we use in our preferred regression specification, is $130.57. As a sensitivity test, we also estimate models using the mean price over all (single coverage) options available to an individual summed with the cost of not waiving coverage. The full sample mean for this price variable is $148.26.

As is typically the case with administrative data, there is relatively little information on individual characteristics. We observe each individual's age, gender, and marital status.[4] Surviving spouses of deceased former employees are entitled to health insurance coverage through the firm. They represent 5% of our retiree sample. We do not have data on income. As a proxy, we use ZIP code level data from the 2000 Population Census for the median income of households headed by adults between the ages of 55 and 64. The sample mean

[2] In Royalty and Hagen's experiment 93% of participants said that they would take up health insurance.

[3] All prices are expressed in 2003 dollars using the Consumer Price Index.

[4] Unlike most studies using health plan enrollment data, we observe actual marital status, as opposed to whether the individual chooses to cover a dependent spouse.

for this variable is \$43,056, which is slightly lower than the national average for this age group (\$47,203). Information on each person's ZIP code is also used to identify people living in rural areas. We use this as a control variable to account for the possibility that other insurance options may be more limited in such areas. An important limitation of this type of administrative data for analyzing health insurance demand is the lack of information on health status. As a consequence, we cannot test for differences in price elasticity related to health risk[5] or address questions related to adverse selection.

Table 2 presents several key variables for three subsamples: (1) pre-1993 retirees, (2) post-1993 retirees with 25 or more years of service, and (3) post-1993 retirees with less than 25 years. Comparisons among these groups give a sense of the price variation generated by the firm's contribution policy, provide an informal check on our identification strategy, and foreshadow our regression estimates of the effect of price. The first thing to note is that the mean age of each group is essentially the same. This is important given that of the demographic variables that we observe, age is most closely related to expected medical expenditures, which in turn will affect the demand for insurance.

There are significant differences across the groups in years of services, which is to be expected given the way the groups are defined. The mean is 30 years for group 2 (post-1993 retirees with 25 or more years of service) and about 18 for the other two groups. As described earlier, groups 1 and 2 face the same contribution rules and therefore the same prices. Thus, a comparison of the take-up rate for these two groups provides a test for a key assumption of our estimation strategy, which is that years of service at the time of retirement affects take-up only through its effect on out-of-pocket premiums. If this assumption is valid, we should see similar take-up rates for groups 1 and 2. Group 3 faces significantly higher prices than 1 and 2. If there is a negative effect of price, we should see a lower take-up rate for this cohort compared to the other two. The data in Table 2 are consistent with both of these predictions.

[5]Two recent studies on health plan choice find that individuals in poorer health have a less elastic demand (Royalty and Solomon 1999; Strombom, Buchmueller and Feldstein 2002).

3.4 Econometric Specification

To fully account for the variation in price, and to control for other observed factors that are likely to affect the demand for insurance, we estimate a reduced form probit model in which the propensity to take up coverage (T*) is a function of the price of coverage (P) and a vector of individual characteristics (X):

$$T^*_{it} = \alpha P_{it} + X'_{it}\beta + u_{it}. \tag{1}$$

The observed analog to T^* is a binary variable, T, that equals one if a person takes up coverage through the firm and equals zero otherwise.[6]

The variables in X include several demographic characteristics: age, gender, marital status and whether the individual qualifies for health benefits as a surviving spouse of a former employee. We interact gender and marital status to account for the possibility that gender differences in take-up behavior may be different for married and single individuals. Since married and single individuals have very different outside options for health insurance, their demand for coverage may be different. In particular, some married retirees will have the option of being covered through their spouse's employer or former employer, an option that will generally not be available for single individuals.[7] Therefore, in addition to estimating the model on a pooled sample with marital status as an independent variable, we also estimate models on separate married and single sub-samples.

Given the source of price variation in our data, it is important to control for when a person retired. We do this with four dummy variables corresponding to the following periods: 1993-1995, 1996-1998, 1999-2001 and 2002-2003. Pre-1993 retirees are the omitted category. Additional controls include the ZIP code level median income from the 2000 Census (as a proxy for income), year dummies and indicator variables for individuals who no longer live in the state where the company is located and for individuals living in rural areas.

[6] Since less than 2% of our sample switches their insurance status over the years we study, we present estimates from pooled probit regressions. We are reassured by this approach since the results are the same for models that that account for the panel structure of our data.

[7] It is worth noting that simple cross-tabulations do not suggest the importance of such differences in outside options. The take-up rates for married retirees (87.2%) and single retirees (87.8%) are not significantly different from each other.

4 Regression Results

Table 3 present the probit results. The price variable is the out-of-pocket premiums for the least costly plan available. The first column is for the full sample; in the next two, the sample is stratified by marital status. Since probit coefficients are not directly meaningful, we report marginal effects (i.e., probability derivatives) evaluated at the mean of the particular estimation sample. The standard errors for these effects are in parentheses.[8] For the price coefficient, we also report an estimate of the mean elasticity evaluated over the estimation sample.

Before turning to the estimated price effects, we will briefly summarize the coefficients on the control variables. There is a strongly positive and statistically significant effect of age in the full sample and the married sub-sample. The results imply that, all else equal, the take-up rate for the oldest individuals in our sample (64 year olds) is 17.7 percentage points higher than the take-up rate for the youngest individuals (55 year olds). This result is qualitatively similar to Chernew, Frick and McLaughlin (1997) and Gruber and Washington (2005). We also find that married men are more likely to take up coverage than married women. The difference is statistically significant (p-value= 0.02). Among single workers, however, the gender difference goes the other way. This pattern is consistent with previous research on take-up (not controlling for price) using nationally representative data (Buchmueller 1996/1997) and with the results of Blumberg, Nichols and Banthin (2001). Individuals who qualify for health benefits because they are the surviving spouse of a former employee are significantly less likely to take up coverage, perhaps, because they have a weaker attachment to the firm.

Controlling for other factors, take-up is higher for individuals living in a rural area and lower for individuals who have moved out of state. The former effect may reflect the dearth of lower cost managed care insurance options in rural areas; the latter may be explained by the fact that people who have left the state are no longer in the service area of the company's insurance plans. The coefficient on our income proxy is statistically significant at the 0.05 level in the full sample and at the 0.10 level in the married sub-sample. This result is also

[8]The calculation of the standard errors takes into account the fact that we have multiple observations on most individuals.

consistent with other studies that find a positive income effect on take-up (Chernew, Frick and McLaughlin 1997; Blumberg, Nichols and Banthin 2001; Gruber and Washington 2005).

In the full sample and the married sample, the retirement year coefficients do not follow any systematic pattern; only one of the coefficients is significant at the 0.10 level, and none are significant at the 0.05 level. In the single retiree sample, none of the year of retirement coefficients are significant. In fact, only one of the four has a t-statistic greater than one. This pattern supports our identification strategy since some of the price variation is coming from differences in when individuals retired.

The results for all three estimation samples suggest that higher premiums reduce take-up. In Table 3, the marginal effect for the full sample is -0.0007, which implies that \$10 increase in price reduces take up by 0.7 percentage points. The estimated price effect for the married sub-sample is similar, which is not surprising given that over three-quarters of the sample is married. Both of these point estimates are statistically significant at the 0.01 level. Like Blumberg, Nichols and Banthin, we find a weaker price effect for single individuals. For that sub-sample, the marginal effect of price is -0.0003 (p-value = 0.12). This difference between married and single retirees may be explained by the fact that married workers are more likely to have other insurance options, most importantly the option of obtaining coverage through their spouse's employer or former employer. Single individuals, in contrast, have fewer substitutes and thus have a less elastic demand.

Evaluated at the sample means, these price effects imply take-up elasticities ranging from -0.10 for singles, -0.15 for married individuals, and -0.16 the full sample. These elasticities are larger in magnitude than those estimated in prior studies. However, such comparisons must be made with caution as the average prices in our data are higher than in those studies. For example, in the data used by both Chernew, Frick and McLaughlin and Blumberg, Nichols and Banthin, the mean employee contribution is about \$20 per month.[9] The mean price for our full sample is \$130.57 per month. When we calculate the elasticity for the full sample at a price of \$23, which is Blumberg, Nichols and Banthin's mean expressed in 2003 dollars, and leave all other variables at the sample mean, we obtain an estimate of -0.03, which is

[9]Comparisons with the other papers are less straightforward. The main independent variable in Gruber and Washington's study is the employee's share of premiums. Royalty and Hagens do not provide enough information to compare the mean prices in their data to ours.

very similar to the elasticity estimates of those prior studies.

4.1 Sensitivity Tests

To test the robustness of these results, we estimated a set of models using alternative specifications and sample definitions. First, we considered the impact of using an alternative price variable. Instead of using the premium for the lowest cost plan, we use the average premium facing an individual. (As before, we add to this premium the payment foregone by not waiving coverage.) The results based on this price variable are qualitatively similar to those from our preferred specification. For the full sample and the married sub-sample, the marginal effect of price is slightly higher when we use the mean price (-0.0012 vs. -0.0007), though the confidence intervals for the two estimates overlap considerably. For singles, marginal effect of price for the mean price model is essentially the same as when we use the minimum price (-0.0004). These similarities are not surprising given that the contribution rules that are the source of identifying variation shift the whole menu of prices. In addition, in the early years of the data the premium differences across plans were small, so the least costly premium and average premium are highly correlated.

Next, we altered the sample inclusion criteria to make the estimation samples more homogeneous. One potential criticism of our main analysis is that part of price variation comes from differences between individuals who retired before and after January 1993. While we explicitly control for the main effect of retirement cohort, it is possible that this does not fully account for behavioral differences among these two groups. Therefore, we re-estimated the models on a sample that excludes individuals who retired before 1993. In this sample, the cross sectional variation in price comes mainly from differences in the number of years of service.[10] The results for this restricted sample are essentially identical to those in Table 3. For example, when we pool married and single retirees, the marginal effect of price is -0.0008. Because average prices are slightly higher for the post-1993 sample, the corresponding elasticity is slightly larger in magnitude (-0.19 vs. -0.16). The point estimates for the single retirees are identical for the two samples.

[10] There is some additional variation caused by the fact that not all retirees live in the service area of the lowest cost HMO.

Finally, we tested to see whether the estimated price effects were sensitive to the way that other variables enter the model. Since the variation in premiums is related to the timing of retirement decisions, we need to ensure that the premium coefficient is not sensitive to different parameterizations of age and retirement year. Therefore, we estimated models using different retirement year groupings. We also tested models in which age enters quadratically or is measured in discrete categories. The main results were robust to these changes.

5 Policy Simulations

The elasticities we estimate can be used to predict the response to the further reductions in employer premium contributions that retirees are likely to face in coming years as well as the impact of different policy proposals for subsidizing coverage. While a full micro-simulation of these policies is not possible with our data, by simulating the percent of retirees taking up coverage at different levels of out-of-pocket premiums, we can provide a sense of how the effect of different policy proposals will depend on the degree to which coverage is subsidized. To this end, in Table 4 we report several simulations based on our regression results.

We begin by estimating the percentage of near-elderly retirees taking up insurance when the coverage is subsidized at different levels. This simulation is most directly relevant to the question of what would happen to coverage if employers reduced the amount they paid on behalf of their retirees. Since the estimated actuarial cost of extending Medicare to this population is similar to the premiums for the plans charged by this employer, these simulations are relevant to Medicare buy-in proposals.[11] For each hypothetical subsidy rate, we report two take-up rates. The first is based on our full sample regression results (Table 3, column 1). One potential criticism of these results is that these results are strongly influenced by the behavior of married retirees, who make up over 80% of our sample. It is likely that some of these married retirees who drop the coverage offered by this employer are covered through their spouse. As a result, the full sample elasticity will overstate the effect of subsidies on the number of people with *any* coverage. Therefore, we also report

[11] According to the Congressional Budget Office (1999), the premium for a Medicare buy-in policy would be roughly $300 per month. This is slightly higher than the average full premium for plans in our data ($275).

simulations where the parameter estimates from the single retiree subsample are applied to all retirees. These results provide a conservative estimate of the responsiveness of coverage to subsidies.

The first row in panel A reports the average monthly cost to retirees and the take-up rate that would result if this employer converted to an "access only" plan whereby it made retiree coverage available but provided no financial contribution. Under this scenario, the least expensive plan option facing the average early retiree in our data would cost $275 per month and between 72% and 80% of eligible retirees would choose to take up coverage.[12] Assuming that the coverage offered by this employer is roughly comparable to Medicare coverage, the results in row 1 of panel A also pertain to a policy allowing 55 to 64 year olds to purchase Medicare coverage at actuarially fair premiums. If the employer or the Federal government were to pay a quarter of the full premium, we estimate that between 80% and 85% of individuals would enroll. A 50% subsidy (which is quite close to the mean subsidy in these data) corresponds to a take-up rate of 87% to 89%, depending on whether we use the coefficient estimates from the full sample or the single subsample.

In the second panel of the table we simulate what would happen if the employer were to drop retiree coverage altogether, leaving retirees to purchase coverage in the non-group market. Because of higher administrative and marketing costs–i.e., "loading" fees–non-group premiums will be even higher than the "access only premiums" in panel A. We adjust premiums by assuming a loading fee of 5% for a group of this size and 30% for non-group plans.[13] Based on this adjustment, coverage comparable to what this employer offers would cost $339 per member per month. Because this amount is much higher than the sample mean, there is a larger difference in the simulated take-up rates corresponding to the full sample and single sample coefficients. The full sample results predict that slightly less than two-thirds of 55 to 64 year od retirees would choose to purchase non-group coverage. Using the single retiree coefficients yields a prediction of just over three-quarters.

In its 2006 Budget, the Bush Administration proposed a $1000 annual tax credit for

[12]An important caveat with this simulation is that this premium amount is outside the range of prices observed in our data. Only 5% percent of the retirees in our sample face monthly out-of-pocket premiums higher than $170.

[13]These estimates are based on typical loading rates reported in the literature. See, for example, Phelps (1997) and Pauly, Percy and Herring (1999).

the purchase of non-group coverage. Two bills by Congressional Republicans (H.R.765 and S.160) propose subsidies of the same amount. The second row of panel B simulates the effect of such a policy. The $1,000 credit would reduce the cost of this type of coverage by roughly 25%. Based on the full sample results, this would in turn increase the percentage of near elderly retirees with non-group coverage by 11.4 percentage points, or by 18% relative to the simulated non-group coverage under the assumption of no subsidy. When we use the smaller elasticity estimate from the single retiree sample, the change in the number of insured is smaller–a gain of 6 percentage points.[14]

6 Discussion

Estimates of the price elasticity of health insurance take-up are necessary for predicting how consumers will respond to policies that subsidize the purchase of health insurance and for making comparisons among such policies that differ in the extent of the subsidy. The best evidence on take-up elasticities comes from studies that use data on employees who are offered health insurance by their employer and are required to make premium contributions toward that coverage.

In this study, we estimate take-up elasticities using unique data from an employer-sponsored retiree health benefit program. Our research design takes advantage of a natural experiment generated by the employer's policy on contributing to retiree health insurance coverage. There are advantages and disadvantages of this approach. The most important advantage is that the employer's contribution policy generates price variation that is uncorrelated with the quality of plan offerings and are plausibly exogenous to individual characteristics that influence the demand for health insurance coverage. Therefore, bias from endogeneity or unobserved heterogeneity is much less of a concern than it is in studies where the price variation comes from differences across employers. A limitation of using data from

[14]The simulations of the tax credit policy are subject to another important caveat. Our results assume that everyone who seeks non-group coverage is able to buy it at essentially community rated premiums. This is not generally the case under current regulations in most states (Simantov, Shoen and Bruegman 2001; Shea, Short and Powell 2001). Therefore, in absence of other policy developments (e.g., non-group market underwriting reforms, the expansion of subsidized high risk pools), consumers who are deemed to be "high risk" may be unable to use the tax credit.

a single firm is that the early retirees that we analyze may not be representative of the entire near-elderly population. For example, employees with access to retiree health benefits tend to work for larger firms and have higher than average incomes (Weller, Wenger and Gould 2004; KFF/HRET 2003).[15]

One striking finding from this study is that despite differences in methodology and the populations studied, our results are quite similar to the results of prior studies based on survey data (Chernew, Frick and McLaughlin 1997; Blumberg, Banthin and Nichols 2001) as well as those that use other types of natural experiments (Gruber and Washington 2005; Royalty and Hagens 2005). Like those studies, we find that the out-of-pocket price of insurance has a small, but statistically significant impact on the decision by early retirees to accept coverage offered by the employer. The implied elasticities range from -0.10 to -0.16, depending on the sample. These elasticity estimates imply that near-elderly consumers would respond to policies that either subsidized non-group insurance or allowed access to Medicare prior to age 65, but this response would be modest. Another limitation of our data (and the data used in the prior studies) is that we do not observe whether individuals who decline coverage through their own firm have alternative sources of coverage, such as through a spouse's employer or former employer. As a result, while we can simulate the number of individuals who would take up coverage that is offered to them at different prices, we cannot assess what fraction of newly enrolled individuals would have otherwise been insured and what fraction would merely switch from one type of insurance to another.

A final limitation of our study is that we are not able to investigate differences in consumer behavior and coverage outcomes related to health risk. These differences have potentially important implications for the cost of different policy initiatives and their impact on coverage. In particular, the effectiveness of tax subsidies and other policies aimed at increasing non-group health insurance coverage will depend on the prices and options available to high risk consumers in the non-group market. These outcomes will, in turn, depend on market rules pertaining to insurer underwriting practices. These questions represent an important direction for future research.

[15] Although the ZIP code average income for our sample is quite close to the national average for this age group, without individual-level data on income we cannot definitively rule this out as an important difference between our sample and the population.

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Table 1. Sample Summary Statistics

	Mean	(Std. Dev)
Dependent Variable		
Take-up of Coverage (0,1)	0.87	(0.33)
Price Variables		
Minimum Premium ($/month)	59.81	(51.76)
Minimum Premium + Amount Foregone by Not Waiving Coverage ($/month)	130.57	(45.56)
Average Premium + Amount Foregone by Not Waiving Coverage ($/month)	148.26	(53.13)
Other Covariates		
Age (years)	60.7	(2.87)
Male (0,1)	0.75	(0.43)
Married (0,1)	0.83	(0.38)
Retired after January 1993 (0,1)	0.87	(0.33)
ZIP Code Level Median Income ($000)	43.06	(12.26)
Surviving Spouse of Former Employee (0,1)	0.05	(0.22)
Lives in a Rural Area (0,1)	0.41	(0.49)
Lives Out of State (0,1)	0.12	(0.32)
Number of Observations		1760
Number of Individuals		510

Note: The minimum premium is the out-of-pocket price of the least expensive single coverage plan available. The average premium is the mean out-of-pocket price of single coverage plans available.

Table 2. Differences in Selected Variables Among Retirement Cohorts

	Pre-1993 Retirees	Post-1993 Retirees By Years of Service	
		≥ 25 Years	< 25 Years
Minimum Monthly Premium	37.65	41.04	82.40[ab]
	(34.89)	(28.07)	(64.46)
Take-Up Rate	0.94	0.92	0.79[ab]
	(0.24)	(0.26)	(0.40)
Age	60.30	59.65	60.46
	(2.86)	(2.77)	(2.92)
Years of Service	18.10	30.08[a]	18.69[b]
	(9.61)	(3.50)	(3.78)
Number of Observations	225	809	726
Number of Individuals	56	232	222

[a] Significantly different from pre-1993 cohort.
[b] Significantly different from post-1993/≥ 25 years of service cohort.

Table 3. Take-up Probit Regression Results

	Full Sample	Married	Single
Relative Minimum Premium	-0.0007**	-0.0007**	-0.0003
	(0.0002)	(0.0002)	(0.0003)
	[-0.158]	[-0.146]	[-0.101]
Married	-0.1280**		
	(0.0274)		
Male	-0.1177	0.0597	-0.1764*
	(0.0458)	(0.0484)	(0.0904)
Married*Male	0.2872*		
	(0.1538)		
Surviving Spouse	-0.5939**		-0.4108**
	(0.1962)		(0.1478)
Age	0.0177**	0.0204**	0.0070
	(0.0038)	(0.0041)	(0.0076)
Ln(ZIP-Level Median Income)	0.0894*	0.0883	0.0814
	(0.0451)	(0.0484)	(0.1183)
Nonmetro Residence	0.0645*	0.0663*	0.0415
	(0.0275)	(0.0309)	(0.0472)
Out of State Residence	-0.0559	-0.0680	-0.1646
	(0.0614)	(0.0679)	(0.1946)
1993-1995 Retiree	-0.1107	-0.1909	0.0365
	(0.0673)	(0.1105)	(0.0568)
1996-1998 Retiree	-0.0175	-0.0494	-0.0209
	(0.0620)	(0.0970)	(0.0953)
1999-2001 Retiree	0.0120	-0.0253	0.0431
	(0.0520)	(0.0864)	(0.0485)
2002-2003 Retiree	0.0679	0.0562	0.0526
	(0.0293)	(0.0526)	(0.0271)
Year 1999	0.0038	0.0001	0.0090
	(0.0125)	(0.0152)	(0.0119)
Year 2000	-0.0033	-0.0110	0.0277
	(0.0163)	(0.0189)	(0.0229)
Year 2001	-0.0060	-0.0083	-0.0196
	(0.0199)	(0.0218)	(0.0473)
Year 2002	-0.0393	-0.0433	-0.0516
	(0.0251)	(0.0279)	(0.0624)
Year 2003	-0.0474	-0.0547	-0.0514
	(0.0299)	(0.0340)	(0.0611)
Number of Observations	1760	1458	302
Number of Individuals	510	416	94
Log Likelihood	-558.42	-462.71	-82.91

Notes: ** = statistically significant at the 0.01 level; * = statistically significant at the 0.05 level. The dependent variable equals 1 if the individual takes up coverage; zero otherwise. Robust standard errors are in parentheses. The figures in brackets are the mean elasticities evaluated over the estimation sample. Pre-1993 retirees and the 1998 year indicator are omitted reference groups.

Table 4. Simulation Results with Reduced Employer Sponsored Health Insurance Benefits

		Predicted Percent of Take-Up	
Market	**Monthly Premium ($)**	**Full Sample Response**	**Single Sample Response**
1. Baseline- Employer Sponsored Insurance	$130.57	87.3%	88.7%
2. Full Premium/Medicare Buy-In			
No Subsidy	274.53	72.2	80.4
25% Subsidy	205.90	80.5	84.9
50% Subsidy	137.26	87.1	88.7
3. Non-Group Market			
No Tax Credit	339.04	63.1	75.6
$1000 Tax Credit	255.71	74.5	81.6

Notes: The monthly premium is the average monthly minimum premium relative to the waive payment for the available health insurance options. The second column is based on the predicted percent of take-up of the full sample from column 1 in Table 3, and the third column is based on the predicted percent of take-up of the full sample from the results in column 3 in Table 3.

Appendix

Figure A-1: Distribution of Employees by Retirement Year