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HEALTH INSURANCE AND EARLY RETIREMENT: EVIDENCE FROM THE AVAILABILITY OF CONTINUATION COVERAGE

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ABSTRACT

Although the vast majority of working individuals aged 55-64 receive health insurance coverage through their employment, many of these individuals face the prospect of losing such coverage should they retire before becoming eligible for guaranteed public coverage through Medicare at age 65. Because the expected medical expenses of this group are large and uncertain, the availability of health insurance coverage after retirement could be a key factor in the retirement decision of older workers. We examine the effect of health insurance on retirement by looking at variation in state and federal "continuation of coverage" mandates, laws which allow individuals to continue purchasing health insurance through a previous employer for a specified number of months after leaving the firm. By allowing individuals to maintain their employer-provided health insurance after retirement, these laws decrease the cost of early retirement for those who do not have other retiree health insurance available. Using data on 55-64 year old men from the Current Population Survey, we find that one year of continuation benefits increases the probability of being retired by 1 percentage point; this represents a 5.4 percent increase in the baseline probability of being retired for this group. We also find that continuation mandates increase the likelihood of being insured after retirement.

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The dramatic postwar decline in the labor force participation of older men in the U.S. has motivated a sizeable body of literature on retirement behaivor. Three factors, in particular, have been studied extensively: the growth of the Social Security program (see for example Burtless 1986; Burtless and Moffitt 1984; Diamond and Hausman 1984; Hausman and Wise 1985; and Sueyoshi 1989), the increased availability and generosity of private pensions (Stock and Wise 1990a and 1990b), and the expansion of federal disability insurance (Bound and Waidmann 1992). One potentially important factor which until recently has not received much attention is the availability of health insurance for retirees. This oversight is especially surprising given the rather consistent evidence that health status is an important determinant in the retirement decision (Bazzoli 1985; Diamond and Hausman 1984). If health status matters in the decision about when to retire, it seems quite natural that health insurance should matter as well.

The increased availability of health insurance for older Americans, especially retirees, has come in several forms. First among them is the introduction in the mid-1960s of Medicare, a federal program that provides near universal health insurance coverage for those over age 65. A second source of health insurance that has grown in importance, particularly for those under age 65 who are not yet eligible for Medicare, is employer-provided post-retirement health insurance. While only 30% of men who retired in the early 1960s received health insurance from their former employers, this fraction increased to almost half for those retiring in the 1980s (Madrian 1993).

This paper looks at the effect on retirement of a third source of health insurance for early retirees, namely continuation of coverage benefits. During the late 1970s and early 1980s many states mandated that employers allow employees who leave their jobs to continue purchasing their group health insurance for a specified number of months. These continuation of coverage beneifts were then extended to all workers in 1986 as part of the federal Consolidated Omnibus Budget Reconciliation Act (or COBRA) legislation. Although this coverage is available to all workers regardless of age, it should be particularly attractive to older workers who face a relatively high price for health insurance in the private market and who are more likely to be subject to the preexisting conditions exclusions that are characteristic of such policies.

To identify the effect of continuation benefits on retirement, we exploit the fact that these benefits were mandated at different times by different states (and finally the federal government) and that the generosity of the mandates varied across states as well. Using data from the Current Population Survey (CPS), we find a strong correlation between the availability of continuation benefits and the likelihood that individuals are retired. Our key finding is that among men aged 55-64, one year of continuation benefits increases the probability of being retired by 1 percentage point; this is 5.4% of the baseline probability of being retired for this group. Furthermore, we find that although the estimated percentage point effects are strongest near the age of Medicare eligibility, as a fraction of baseline retirement probabilities they actually decline with age. Although this latter result is somewhat counterintuitive, it is consistent with other work which examines the effect of continuation coverage on flows into retirement (Gruber and Madrian 1993). We also find that continuation mandates significantly increase the likelihood that early retirees are covered by employer-provided health insurance after retirement. This effect is much larger than the implied effect on retirement, suggesting that much of the increase in coverage is occuring among those individuals who would have retired even in the absence of such benefits.

The organization of the paper is as follows. Section I provides some motivation for why health insurance should matter in the early retirement decision. Section II then outlines the state and federal continuation of coverage laws which we use to identify the effect of health insurance on retirement. This is followed in Section III by a model which formalizes the effect of health insurance on retirement. The data and regression framework are presented in Section IV, and the results follow in Section V along with a comparison to our findings from dynamic models of retirement behavior. Section VI then considers the impact of continuation mandates on insurance coverage. The paper concludes in Section VII with a discussion of the methodological and policy implications of our results.

I. Health Insurance and Retirement--Should It Matter?

The high and variable level of medical expenditures for persons aged 55-64, without the guarantee of public coverage through Medicare for those over age 65, means that the availability of health insurance coverage could be a key factor in determining the timing of retirement. However, until recently there has been little study of the effect of retiree health insurance coverage on retirement patterns. Two recent papers have attempted to model the role of health insurance in the retirement decision. Lumsdaine, Stock and Wise (1992a) incorporate the value of Medicare into an option value model of retirement and find no effect of Medicare eligibility on the retirement hazard. Their result is not surprising, however, as they estimate their model on a sample of workers from the same firm, all of whom have employer-provided post-retirement health insurance which is much more generous than Medicare. Gustman and Steinmeier (1992) use information from the Retirement History Survey, a longitudinal survey from the 1970s, to ascertain whether individuals have employer-provided retiree health insurance, and data from the 1977 National Medical Care Expenditure Survey, to impute the value of that insurance based on individual characteristics. They also find very small effects of retiree health insurance on retirement decisions.

The results of these two studies are at odds with both intuition and with what individuals report about the importance of health insurance in the retirement decision. In a recent Gallup poll, 63 percent of working Americans reported that they "would delay retirement until becoming eligible for Medicare [age 65] if their employers were not going to provide health coverage" despite the fact that 50 percent "said they would prefer to retire early--by age 62" (Employee Benefit Research Institute 1990). The apparent contradiction between the importance of health insurance as stated by individuals and that estimated by these two previous studies provides a further motivation for our research.

Health Status of Older Individuals

That individuals should cite health insurance as an important consideration in the retirement decision is not surprising, as older persons are fairly likely to need expensive medical care. Tables 1-4 compare the health status of individuals by age along a number of dimensions. The simplest measure, self-reported health status, is shown in Table 1. The fraction of individuals who report being in fair or poor health increases markedly from ages 45-54 (19.7%) to ages 55-64 (31.3%). While recent research has suggested that self-reported health status may be a poor indicator of the actual severity of an individual's clinical conditions (Bazzoli 1985), it may be the most accurate measure of an individual's valuation of health insurance coverage. Thus, these figures suggest that insurance valuation will rise dramatically with age.

Furthermore, as Table 2 shows, health status as measured by doctor-diagnosed health problems deteriorates with age as well. The incidence of many of the health problems listed (stroke, cancer, heart attack, arteriosclerosis, emphysema, and heart disease) more than doubles between ages 45-54 and ages 55-64. Furthermore, almost three-quarters of those aged 55-64 have been diagnosed with at least one of the 11 conditions listed. Not surprisingly, relative to those aged 45-54, individuals 55-64 are more likely to be admitted to the hospital over the course of a year and spend more time there once admitted (Table 3).

The most direct evidence that health insurance should be valued relatively highly by older workers, however, is that the actual medical expenses incurred by those aged 55-64 are much higher than those of younger individuals (Table 4). In every category, not only do expenditures rise with age, but the variance increases as well. In 1990 dollars, total medical expenditures of those 55-64 averaged $2144. This represents 5.4% of average total family income for this age group, 6.9% of average total family income for retired individuals, and 30% of the average pension income of early retirees.[1] A one standard deviation increase in expenditures for a 55-64 year-old would represent an additional 16.5% of family income. Total *family* medical expenditures would naturally constitute a much higher fraction of income. Thus, it is easy to see why older individuals should be concerned about their health insurance coverage after retirement.

Health Insurance Coverage and Costs

Given the costs of health care for older workers, it should not be surprising that older individuals are no more likely to be uninsured than their younger counterparts as is shown in Table 5. The sources of health insurance coverage, however, differ with age. Even though employment-based health insurance is the predominant source of coverage regardless of age, older individuals are less likely than younger persons to have employment-based health insurance and much more likely to be covered by a nongroup (individual) or other group policy. This

[1] Expenditures as a fraction of income are calculated using income data from the March 1990 Current Population Survey.

suggests that individuals who retire early but who do not have access to employer-provided health insurance turn to the individual market for insurance.

The bottom two panels of Table 5 break down the sources of health insurance coverage by employment status. There are three major differences between the sources of health insurance coverage for those who are and are not employed. First, one-fifth of non-working older persons are insured through Medicare or Medicaid, while only 1% of the older employed receive coverage from one of these two sources. Second, older non-working individuals are 40% less likely to be uninsured than their younger counterparts. Third, relative to the young, the older non-working are six times more likely to be covered by employer-provided health insurance in their own name.

These last two differences are explained in large part by the availability of employer-provided post-retirement health insurance. 45% of individuals work in firms that provide retiree health insurance benefits[2]. The older non-working, who are more likely to be retired than the young non-working, are therefore more likely to be covered by employer-provided retiree health insurance.

There are, nevertheless, a substantial number of older individuals who are not covered by either employer or government-provided health insurance. It is these individuals who find themselves in the market for individual health insurance and who we would therefore expect to benefit from the availability of continuation coverage. The reason is simple--insurance in the individual market is typically quite expensive.

[2] See Madrian (1993) for background on the structure and availability of post-retirement health insurance.

Employers have significant cost advantages in providing health insurance. By pooling the risks of many individuals, they are able to lower administrative expenses and reduce adverse selection. These two factors alone are estimated to reduce the cost of providing insurance in large (10,000 or more employee) firms relative to small (1-4 employee) firms by 40% (Congressional Research Service, 1988). For older individuals, the cost differential between employer-provided and individual health insurance is exacerbated by the fact that policies in the individual market are typically age-rated, while within the firm younger workers subsidize the health insurance costs of their older co-workers. The Congressional Research Service (1988) reports that the cost to employers of providing insurance coverage for 55-64 year old males is three times that of providing coverage to males under 40; for females, the ratio is two to one.[3]

In Massachusetts, the average cost of family health insurance coverage per employee in 1989 was $3882.[4] When inflated by the medical care component of the Consumer Price Index, this is equivalent to $5047 in 1993 dollars. In contrast, a New England commercial insurance company is offering a family policy for a 58 year-old male with a one-year preexisting conditions exclusion at a price of $8640. This represents 26% of the average family income of retired individuals aged 55-64 in Massachusetts. Individual policies may also be medically underwritten so that sick individuals may face substantially higher prices or may not be able to purchase a policy at all.

The coverage available in the private market not only is expensive, but is typically less generous than employer-provided health insurance as well. Table 6 compares the health

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[3] Of course, to the extent employer costs can be shifted to the wages of employees in an age-specific fashion, older individuals will bear these higher costs. See the discussion in Section III.

[4] Authors' calculation using unpublished data from the Health Insurance Association of America.

insurance benefits of individuals covered under group and nongroup policies in 1977. In every category, those covered under nongroup policies receive more limited benefits. Relative to those with nongroup coverage, those with group policies are more than twice as likely to receive major medical coverage or coverage for physician office visits and prescription drugs, and more than 50 percent more likely to receive ambulance, mental health, and outpatient diagnostic service coverage. Furthermore, nongroup policies generally feature both higher deductibles and higher copayments. Thus, relative to the individual market, group coverage offers individuals higher quality insurance coverage at a significantly lower price.

II. Continuation of Coverage Laws

For those individuals whose employers do not offer retiree health insurance, an alternative to purchasing health insurance in the individual market is provided by various state and federal continuation of coverage laws. These laws mandate that employers sponsoring group health insurance plans offer terminating employees and their families the right to continue their health insurance coverage through the employer's plan for a specified period of time. The laws generally apply to all separations (except those due to an employee's gross misconduct), although in some states benefits are restricted to those who leave their jobs involuntarily.[5] They often also provide benefits to divorced or widowed spouses and their families. The first such law was implemented by Minnesota in 1974. More than 20 states passed similar laws over the next decade before the federal government, as part of its 1985 Consolidated Omnibus Budget

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[5] Because retirement is a voluntary separation, we treat those states whose laws apply only to involuntarily terminated employees as states without laws.

Reconciliation Act (COBRA), mandated such coverage at the national level. Continuation coverage is now commonly referred to as COBRA coverage, a nomenclature we will also use.

The various state statutes are summarized in Table 7.[6] The length of coverage is generally quite short, from 3-6 months, although 10 states mandate coverage of nine months or more. Although most state laws stipulate that an employee must have been covered by an employer's insurance for 3-6 months before being eligible for continuation coverage, this requirement is not likely to be binding on older workers, most of whom have been with their current employer for many years.[7] The states laws also apply only to firms that actually purchase insurance through an insurance company; self-insured firms, under the 1974 Employee Retirement Income and Security Act (ERISA), are not subject to these (or any other) state mandates.[8]

Although similar in spirit, the state and federal laws differ in a number of important ways. First, the length of coverage mandated under the federal law, 18 months, equals or exceeds that mandated by all but one state (as of January 1987, Connecticut law provides for up

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[6] Details on state laws are from Hewitt (1985) and Thompson Publishing (1992) and have been cross-checked against the actual state statutes. Table 7 lists only those states with laws that apply to employees who terminate their employment voluntarily. There are, in addition, several states with laws that apply only to involuntarily terminated employees.

[7] Almost 95% of retirees have job tenure of at least ten years by the time they retire (Madrian 1993).

[8] In a related paper, we incorporate a correction factor which accounts for the exclusion of some firms from the effects of these laws (Gruber and Madrian 1993). This has little effect on the significance of the estimates of the effect of continuation coverage on retirement, although the magnitude increases two to three-fold.

to 20 months of coverage).[9] Second, there is no minimal length of time for which an employee must be covered under an employer's plan before being eligible for continuation benefits. Third, the federal law applies to self-insured firms, who are exempt from the state laws, as well as to those who purchase their coverage from insurers. The federal law, however, does not apply to small firms employing less than 20 workers. Finally, employees of religious organizations and the federal government were exempt from COBRA, although federal employees have subsequently been included (beginning in 1990). When the specific details of the state and federal statutes are at odds, firm provision of continuation benefits is governed by the law which provides for more generous coverage.

The effective dates of the state laws are listed in Table 7. The federal coverage mandated under COBRA was phased in. Beginning in July 1986, firms had to offer continuation benefits at the start of their next plan year. For workers provided health insurance under union contracts, such benefits did not have to be offered until the next contract negotiation after January 1987.

Both the state and federal laws stipulate that the employee must pay the full cost of the coverage. At the federal level, this is defined specifically as 102% of the average employer cost of providing coverage. The coverage must be identical to that provided to similarly situated active employees, including the option to continue enrollment in supplemental insurance plans (such as for vision or dental care) if these are available. Although 102% of the employer's cost is typically much more than individuals pay as active employees, it is, as already noted,

[9] 18 months is the maximum length of coverage available following the voluntary or involuntary termination of employment. COBRA also provides up to 36 months of coverage for family members who would otherwise lose their insurance coverage through events such as an employee's death, divorce from the employee, or the employee's eligibility for Medicare.

substantially less than the cost of buying equivalent coverage in the private market, especially for older workers.

Because continuation coverage is a relatively new phenomenon (at least at the national level), information on the extent of continuation coverage is somewhat scarce. Zedlewski (1993) estimates that in 1988, 5.2% of retired workers aged 55-64 were covered by COBRA health insurance. This figure must be interpreted relative to the number of individuals who could be expected to take up such coverage. The 52% of individuals aged 55-64 with retiree health insurance are not likely to be covered, and the 21% of individuals who were not insured through their former employer are not eligible. Similarly, those who have been retired for more than 18 months will have exceeded their potential eligibility. Tabulations from the 1987 National Medical Expenditure Survey indicate that one-third of retired individuals aged 55-64 have been retired for less than 18 months. If we take the group who could potentially be affected by COBRA to be one-third of retired individuals between ages 55 and 64 who worked in firms that provided health insurance but did not provide retiree health insurance, we would expect at most 9% of early retirees to be covered. That 5.2% receive continuation benefits suggests that 58% of the retired population who would be at all likely to be covered by COBRA actually are. As knowledge about the availability of such coverage has become more widespread since 1988, this fraction may be higher today.

An alternative calculation is possible using figures reported in Flynn (1992). She uses data from a large firm that administers COBRA claims to estimate that 23% of individuals who qualified for COBRA coverage because of retirement elected to receive benefits. If we would only expect the 30% of individuals in firms that offer health insurance but do not offer retiree health insurance to even consider purchasing COBRA insurance, this take-up rate implies that

75% of those most likely to be covered by continuation benefits actually are. Both of these calculations, therefore, suggest that retirees without an alternative source of health insurance coverage are quite likely to elect continuation coverage.

For all COBRA beneficiaries, the average length of time on COBRA was 7 months (Flynn 1992). Individuals over age 61, however, maintained their coverage for a much longer period of time--about 12 months on average. This finding is not surprising. First, younger individuals are more likely to find alternative coverage through a new job or a spouse's employment. Second, COBRA coverage provides a larger subsidy for older workers; with a lower relative price, they should therefore demand more coverage.

Table 8 compares the distribution of health insurance coverage in 1984, two years before COBRA was first implemented, and in 1989, two years after it had been phased-in. Note that employment-based health insurance coverage is more prevalent after COBRA, and that this effect is confined to those who are not employed, exactly the group whom we would expect to be insured under COBRA. This finding is similar to evidence presented in Rogowski and Karoly (1992) who examined the primary source of insurance coverage after retirement based on the source of insurance coverage before retirement before and after COBRA. They find that in the pre-COBRA period, 72% of individuals who retired from jobs with employment-based health insurance continued to be covered by that insurance upon retirement. After COBRA, this figure rises to 78.5%.[10] Taken together, the evidence on take-up rates and the increase in the extent of employer-provided health insurance coverage among early retirees after COBRA suggests that

[10] We present a stronger test of the effect of continuation mandates on insurance coverage in Section VI.

older workers who retire early and who do not have an alternative source of coverage actually avail themselves of the continuation benefits to which they are entitled.

III. Modelling the Effect of Health Insurance on Retirement

We present a simple graphical exposition of the effect of health insurance benefits on the retirement decision, along the lines of Burtless (1986) and Burtless and Moffitt (1984). We consider both retiree health insurance in general and continuation benefits more specifically. Figure 1 shows the budget constraint facing an older worker between the ages of 55 and 65. The horizontal axis represents the age of retirement. The vertical axis measures the certainty equivalent (CE) of consumption from age 55 onward. This differs from the earlier literature which has typically considered the relationship between the age at retirement and the actual level of future consumption rather than the certainty equivalent of future consumption. This departure is necessitated by our focus on the effect of insurance coverage.

We assume that workers receive health insurance on their current job but that they may or may not have retiree health insurance coverage. Firms that provide post-retirement health insurance do so on the same basis for both workers and retirees, and these benefits cease upon eligibility for Medicare.[11] We also assume that once a worker leaves his current job, he will remain retired for the rest of his life. To simplify the analysis, we ignore the effects of both Social Security and pensions; they could, however, be easily incorporated into the analysis.

In the model, as in the real world, workers who retire without health insurance coverage have two options: they may purchase an individual policy, or go uninsured. In either case, their

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[11] In reality, most retiree health insurance plans do "top off" Medicare to some extent. This does not alter the main conclusions of this section.

out-of-pocket medical expenditures will be significantly higher than if they receive retiree coverage or have the option of continuing their group coverage. For a worker with retiree health insurance, the slope of the budget constraint will be the after-tax wage, which is depicted by line line AB in Figure 1. Since medical expenditures are insured, there is no uncertainty about future consumption.

For the worker without retiree coverage, the relative position and slope of the budget constraint depends on two factors. First, because individuals are risk averse, those without retiree health insurance will have a lower level of CE consumption; this places the no insurance budget constraint below that of an insured worker.[12] Second, because both the mean and the variance of medical expenditures rise with age, a year of health insurance coverage is worth more at older ages. The cumulative reduction in CE consumption will be greater at younger retirement ages, but the incremental effect will smaller. This latter effect gives curvature to the no health insurance budget constraint, line CD in Figure 1. At age 65 there is a jump in the no insurance budget constraint as Medicare equalizes the opportunities of all individuals.

If leisure is a normal good, retiree health insurance will lead to earlier retirement, at age $R_1 < R_0$, because such coverage makes individuals wealthier. As individuals are more risk averse, the wealth effect will increase as the both the level of the no health insurance budget constraint falls and its slope becomes steeper.

Now consider the effect of a continuation mandate that provides one year of subsidized insurance coverage relative to having no health insurance. For the risk-neutral worker, this is simply equivalent to an increment to wealth equal to expected medical costs for a year minus the

[12] Risk aversion in this model operates in a similar fashion to higher expected medical costs.

cost of the group policy.[13] This increment rises in value as the worker ages since expected medical expenditures increase with age. Thus, the budget constraint with a continuation option, line EF, lies above the no health insurance constraint but below the retiree coverage constraint. At younger ages, it is very close to the no insurance constraint; at age 64, it differs from the retiree coverage constraint by the cost of the group coverage. As workers become more risk averse and the no health insurance constraint becomes steeper, the distance between the no health insurance and the continuation coverage constraints will increase, and this increase will be greater at older ages. In this case, the value of one year of coverage will equal expected medical costs minus the cost of the group policy plus the increase in CE consumption implied by eliminating uncertainty in that year.

The value of both retiree health insurance and continuation benefits will rise with the cost of being uninsured or the cost of buying individual insurance in the private market. The important difference between these two sources of coverage, however, is their age patterns: while retiree insurance coverage is of highest value to very early retirees, continuation benefits are more valuable at older ages. Because of this, we might expect continuation benefits to be used primarily by older workers seeking a "bridge to Medicare" which allows them to retire a certain number of months before age 65 without losing group coverage. If this is the case, we would expect the effect of continuation coverage on retirement to be greatest at older ages.

There are, however, a number of complications which cloud this basic intuition. The first is the empirical violation of one of our assumptions, namely that retirement is permanent. Diamond and Hausman (1984) report substantial reentry rates for early retirees; among 55-64

[13] Once again, this amount is presumably positive even for a risk-neutral worker due to cross-subsidization of the group policy by younger co-workers.

year olds, the one-year reentry rate is approximately 15%. Sueyoshi (1989) finds that one-third of the elderly "partially retire", moving from permanent employment to less than full-time work. To the extent that continuation mandates facilitate movement across jobs, rather than permanent retirement, they may have larger effects at younger ages than was depicted above.[14]

In this analysis, we have assumed that retiree health insurance offers pure rents to workers in the firms that offer this type of coverage. In labor market equilibrium, presumably at least a portion of these rents will be reflected in lower wages for workers with retiree coverage. The extent to which these compensating differentials offset the benefits of retiree health insurance at each age will be a function of the employer's ability to set relative age-specific wages freely,[15] the mobility of workers across firms at different ages, and the excess of the cost of continuation benefits over the group premium paid by the early retiree.[16] The existence of compensating differentials may affect both the location and the shape of the budget constraint facing the potential retiree; the net effect on retirement age will be a function of the nature of the compensating differential.[17]

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[14] One important consideration, of course, is whether this reentry is to jobs that offer health insurance; unfortunately, there is little evidence on this question.

[15] See Rosen (1986) for a discussion of the theory of compensating differentials. Gruber (1992) provides some evidence that shifting the costs of employer-provided benefits to distinct demographic groups in the workplace is feasible.

[16] Huth (1991) reports that the health insurance claims of COBRA recipients exceed those of active employees by 50%. This difference in costs is attributed to adverse selection; it is the sickest individuals who will find continuation coverage most attractive and they will therefore be the ones most likely to take it up. Similar evidence is provided in Long and Marquis (1992).

[17] For example, if the entire cost of the benefits is shifted to older workers, this will lower the slope of the budget constraint with continuation benefits (Figure 1) relative to the budget constraint without benefits (because wages for those with benefits fall) which will have both income and substitution effects on the retirement decision.

Finally, we have ignored the possibility that workers may be liquidity constrained in making their retirement decisions. The fact that most retirees have few liquid assets (Diamond and Hausman 1984) implies that such liquidity constraints may be empirically important in determining retirement dates. This explanation is suggested in both Diamond and Hausman (1984) and Burtless and Moffitt (1984) in their discussion of why Social Security benefits do not seem to affect retirement until they actually become available at age 62. Samwick (1993) finds that much of the estimated increase in retirement probabilities attributed to Social Security occurs among those with pensions, suggesting that all workers would like to take advantage of these benefits early, but that only those with pensions can afford to do so. The presence of liquidity constraints could increase the effect of continuation benefits at younger ages, as the wealth increment which these benefits represent could be loosening these constraints.

IV. Data and Regression Framework

Data

The data for this study must meet two key criterion. First, in order to exploit the variation in state and federal continuation of coverage legislation they must extend over a number of years before and after 1986. Second, there must be a large sample size so that the effects of state law changes on older workers can be identified. The data which best meet these two criterion is the Merged Outgoing Rotation Group (MORG) sample of the Current Population Survey (CPS). The CPS is a nationally representative survey which interviews over 50,000 households each month. The MORG file contains information on demographic characteristics and labor force attachment during the survey week for one-quarter of each month's sample for each month of the year. This is the largest available annual data set on individual labor force behavior in the United States.

Recent studies of retirement behavior have focused on dynamic modelling of the transition into retirement. In this paper, we instead use a static model of whether or not an individual is currently retired since the only labor force information we have in the MORG is for the week of the interview. Evidence on the stock of retired persons can still be useful for considering the effect of continuation mandates on retirement; if the laws are affecting flows, they should affect stocks as well.[18] Furthermore, dynamic modelling strategies for retirement decisions using survey data sets often suffer from an important econometric difficulty, known as dynamic sample selection bias (Diamond and Hausman 1984). In the case of continuation coverage laws, this bias arises from the fact that the set of individuals observed actually working after the law has been in place for a number of years will be less likely to retire in response to the law than would the entire population because those most likely to respond will have already retired. When the sample is selected on the basis of those who are still working, the results will therefore be biased against finding an effect of the law.[19] In a multivariate setting, the bias cannot be signed a priori, and with time varying covariates in the model, such as months of continuation coverage, it

[18] This is not strictly true if the mandates affect the number of persons who decide to work at all; in this case, both the numerator and denominator of the labor force participation rate would be increasing, and the effect on the stock would be ambiguous. This is not likely to be a problem for the sample of older males on which we focus.

[19] An alternative way to see this point is to imagine a law that applied to a cohort rather than to an age group. The individuals who are most likely to respond to this law will do so in the first year. In the next year, by selecting on the set of individuals who have not yet retired, we will bias the results against finding an effect of the law. When the law applies to an age group, rather than a cohort, this effect is attenuated by the fact that new members arrive into the age group.

is impossible to correct for this "left-censoring."[20] Our static regressions, which include all 55-64 year old males regardless of initial work status, do not suffer from this bias.

On the other hand, the major disadvantage of our static framework is that we cannot control for the characteristics of the job from which the individual has retired. This will be important if, for example, there is a systematic correlation between the passage of these mandates and the nature of the jobs in the states where they are passed. In the regression analysis, we attempt to reduce any bias which results from this potential correlation by controlling for the time invariant characteristics of the states that pass these mandates. In Section V we will contrast our findings from this static regression with those from dynamic models which allow us to better control for the types of jobs held by individuals.

Regression Framework

We focus on two definitions of retirement: whether or not an individual reports being retired, and whether or not an individual is out of the labor force. Both are based on a CPS question which asks about the major activity in which an individual was engaged during the week before the survey. The latter definition is useful because retirement may be a subjective term which takes on different meanings for different individuals. These retirement definitions are clearly problematic along at least two dimensions. First, we are unable to contrast the effect of these regulations on both "full" and "partial" retirement as is done in Burtless and Moffit (1984) or Sueyoshi (1989). Second, we are unable to account for reentry into the labor market as discussed in Diamond and Hausman (1984). Nevertheless, these measures should provide

[20] It is possible, however, to test for the magnitude of the bias; see Gruber and Madrian (1993).

reasonable estimates of the effect of continuation mandates on the propensity of older workers to remain employed.

Our sample consists of men between the ages of 55 and 64. Overall, 20% of the sample report being retired and 35% are out of the labor force. The average level of education is 12 years, and 9.5% of the sample is nonwhite.

We estimate the following probit model of retirement:

$$Pr(Retired_{ijt}) = \Phi(\alpha + \beta_1 \cdot X_{ijt} + \beta_2 \cdot State_j + \beta_3 \cdot Time_t + \beta_4 \cdot Law_{jt}) \tag{1}$$

where i indexes individuals, j indexes states, and t indexes time. X_{ijt} is a set of individual demographic characteristics, $State_j$ is a set of state dummies, $Time_t$ is a set of year and month dummies, and Law_{jt} is the number of months of continuation coverage available in state j at time t.[21] The state fixed effects control for any time invariant characteristics of a state which may be correlated with the state's propensity to pass continuation legislation. We include a set of year dummies to control for national trends in retirement behavior which may be correlated with the passage of these laws, and month dummies to control for seasonal patterns in retirement behavior. Thus, the effect of the laws is identified in this model by changes in retirement behavior in states which passed the laws (or which were affected by the Federal law), relative to those which did not, during the period after the laws were passed. Further identifying variation comes from differences across states in the number of months of eligibility which these laws allow. Since we have monthly data, we phase-in the federal law in 12 equal increments between July 1986 and June 1987.

[21] We exclude individuals from two states from our sample: Hawaii, which has mandated health insurance for all employees, and West Virginia, for which we were unable to definitively date the effective date of their continuation mandate.

V. Results

The basic regression results are reported in Table 9. The first column reports the probit coefficients from the self-reported retirement equation while the second column gives the marginal probabilities implied by these coefficients.[22] The same is done in the third and fourth columns using not in the labor force as the definition of retirement. More education is associated with a slightly lower probability of being retired and a much lower probability of being out of the labor force. Being non-white is associated with a lower probability of retirement but a significantly higher probability of being out of the labor force. Individuals who are married are less likely to be either retired or out of the labor force. The age pattern of retirement propensities is familiar from the previous literature; there is a large jump in the probability of being retired at age 62, and individuals age 64 are 25% more likely to be retired than individuals age 55. This pattern is even more pronounced for being out of the labor force, as the probability at age 64 is 40% greater than the probability at age 55.

The availability of continuation coverage has a sizeable and significant effect on the probability of being retired. One year of coverage raises the probability that an individual is retired by 1.1 percentage points which is 5.4% of the baseline probability of being retired in this sample. For the not in the labor force regressions, the estimated effect of a year of continuation coverage is of approximately the same magnitude as in the retired equation (although the

[22] For dummy variables, the marginal probabilities are calculated by predicting the probability of retirement with the dummy equal to one for the entire sample, predicting the probability with the dummy set equal to zero for the entire sample, and taking the average of the difference in these predictions across all individuals. For continuous variables, the marginal probability is calculated by predicting the probability at the current level of the variable, predicting the probability by adding one to the variable, and once again taking the average of the difference in these predictions across individuals. The marginal probability on months of coverage is the probability increase associated with going from 0 to 12 months of coverage.

coefficient is only significant at the 10% level), and suggests an increase in the baseline probability of being out of the labor force of 2.8%

The model described in Section III suggests the possibility that the effect of continuation mandates on retirement could vary with age; intuitively, it seemed that this effect should be strongest at older ages. In Table 10, therefore, we free-up the effect of months of continuation coverage by age. The second and fifth columns present the marginal probability derivatives of the probits. The third and sixth columns express these percentage point increases in retirement propensities as a fraction of the baseline retirement rate at each age. This allows for a more natural interpretation of the percentage effects of continuation benefits on retirement at each age.

In both equations, the coefficients rise with age and are statistically significant at ages 62 and above. The pattern of effects as a fraction of baseline retirement probabilities, however, is not uniformly supportive of the hypothesis suggested in Section III. For the retirement equation, there is actually a declining pattern of effects by age; for the not in the labor force equation, the effects are slightly increasing with age.

There are several possible explanations for this counterintuitive finding that the effects are not proportionately greatest at the ages near Medicare eligibity. The first is the set of theoretical issues we raised in Section III, such as the possibility that individuals may face liquidity constraints which are loosened by this temporary health insurance. The second reason is statistical: we may not have enough power in these probits to distinguish true larger effects at older ages from the effects at younger ages. Given the precision of our estimates, this seems an unlikely explanation for the unexpected age pattern of our results.

Alternatively, it may be that our result is spurious. One potential problem with our identification strategy is that the passage of these laws could be correlated with some other change in retirement behavior in these states. Alternatively, it could be that the laws themselves are endogenous responses to changes in retirement propensities among the population; that is, if more individuals are retiring, states may respond by mandating benefits that cover individuals after their retirement.

One form of potential endogeneity could be that the propensity of legislatures to mandate continuation coverage is correlated with long term within-state trends in retirement behavior. In this case, even with state fixed effects included in the regression, there will be a spurious correlation between changes in retirement behavior within a state and the passage of a continuation mandate. One possible control for such spurious causation is to include in the regression not only state effects but state-specific trend terms; that is, we interact each state effect with a trend for the ten year period.[23] The results from this specification check are presented in Table 11. For the not in the labor force regression, the age-specific coefficients are virtually unchanged from those in Table 10; in the retirement equation, the coefficients are slightly larger, but once again the effects are very similar.

A further potential problem with these findings is that it may not be appropriate to compare the effects of the state and federal mandates. As we noted earlier, these mandates differ along a number of dimensions, the most important being that the state mandates do not apply to self-insured firms, while the federal mandate does not apply to small firms. In results not reported, we have rerun these regressions for the period prior to July 1986 in order to restrict our anlaysis to the effects of the state laws. The results are somewhat stronger than those in Tables 10 and 11, although the age patterns are similar.

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[23] The trend is monthly, taking on values of 1 to 132. This type of "random growth" or "fixed trend" estimator is suggested by Heckman and Hotz (1988) and is used by Jacobson, Lalonde, and Sullivan (1992) and Gruber (1993).

Comparisons with Findings from Dynamic Models

In related work (Gruber and Madrian 1993), we consider the effect of continuation benefits on transitions into retirement using two different data sets--the March files of the CPS, and the Survey of Income and Program Participation (SIPP). These data sources allow us to estimate dynamic retirement models and to control for some characteristics of the jobs from which individuals retire. The sample sizes are much smaller than we have with the MORG data, however, and we are confronted with the dynamic sample selection issue discussed above. Nevertheless, this study confirms the two key findings of the research reported above. First, there is a sizeable and significant effect of continuation coverage on retirement behavior. Using one-year retirement transitions in the March CPS, we find that one year of continuation coverage raises retirement propensities by 1.4 percentage points. This is quite similar to the 1.1 percentage point effect estimated in this paper using the MORG data. Furthermore, the implied effect on the hazard rate in both the March CPS and SIPP data is identical.

Second, despite the presumption that these laws should act as a "bridge to Medicare," the estimated effects in these dynamic models do not rise with age either. Figure 2 graphs the change in the propensity to be retired from having a year of continuation coverage estimated from the MORG regressions (column 3 of Table 10) along with the percentage increase in retirement probabilities estimated using transition data from the March CPS (Gruber and Madrian 1993).[24] To facilitate comparability, the two series are each normalized to take on a value of one at age 55. While the pattern of effects differs somewhat at the early ages, both series show a similar decline after age 59, and the effect at age 64 is approximately 1/3 as large as that at age

[24] These latter coefficients comes from the model which is most comparable to that used in this paper. See Gruber and Madrian (1993) for a number of extensions to this basic dynamic model.

55. Thus, our two main findings from the static framework employed in this paper are borne out in the dynamic model that we employ elsewhere.

It is also interesting to consider what the magnitudes of these findings imply about individual valuation of continuation benefits by comparing them to the estimated increase in retirement propensities following an increase in post-retirement income. The results from a static probit model of retirement in Samwick (1993) suggest that a $5000 increment to Social Security wealth increases the retirement hazard by approximately 8%. In a dynamic stochastic programming model employed by Stock and Wise (1990a and 1990b) and Lumsdaine et al. (1992a and 1992b), they find that a $5000 increase in the value of pension wealth leads to an increase in the retirement hazard of between 10 and 13% for individuals between the ages of 55 and 64.[25]

The basic specification of Gruber and Madrian (1993) finds that one year of continuation coverage raises the retirement hazard by 19%. This implies that a year of continuation benefits is valued at between $7,300 and $12,000 in terms of post-retirement wealth. Based on the cost information reported in Section II, a COBRA policy would save an older worker approximately $4,500 per year on the price of family coverage. Taken at face value, this results suggest that workers value the insurance received from continuation policies at a somewhat higher level than its associated cost savings. This may reflect the fact that the individual policy we priced, as with most individual policies, excluded pre-existing conditions for some period. Alternatively, it may be that a number of early retirees must pay substantially more for individual policies or are unable to obtain such policies at all.

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[25] We are grateful to Andrew Samwick, Robin Lumsdaine and Jim Stock for performing these calculations for us.

VI. Insurance Coverage

In this section, we consider the effects of continuation mandates on the insurance coverage of early retirees. If continuation mandates are having an effect on the retirement decisions of older workers, then, by definition, they should be affecting their insurance coverage as well. Thus, evidence that such mandates increase insurance coverage among early retirees provides a necessary (but not sufficient) specification check of our result that these mandates affect retirement behavior. Furthermore, it is interesting to contrast the direct effects of these mandates on insurance coverage with their indirect effects on retirement behavior. To what extent do continuation mandates affect the "inframarginal" individual, who would have retired in their absence, relative to the "marginal" individual whose retirement decision is made in response to their presence?

In order to investigate the effect of continuation mandates on insurance coverage, we use data from the Survey of Income and Program Participation (SIPP).[26] The SIPP is a nationally representative survey of households designed to collect information on the economic and demographic characteristics of individuals and their families. We use data from the 1984, 1985, 1986 and 1987 panels of the SIPP. Sample members are interviewed every four months for roughly 2½ years and asked to provide information about their labor market activity, income, and participation in welfare and transfer programs over the previous four months. The first interviews of the 1984 Panel were conducted in October of 1983, while the initial interviews for subsequent panels commenced in February of the corresponding calendar year. For previously

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[26] To keep the sample of individuals comparable to the MORG data used in this paper, one could in principal use the March CPS to look at insurance coverage over a similar time period. Unfortunately, a 1988 change in the questionnaire which altered the reported coverage rates of older individuals who were not working (precisely the group of interest) precludes performing a reliable analysis with this dataset.

cited reasons, we exclude individuals living in West Virginia and Hawaii. We also drop individuals from several other small states because, out of concern for confidentiality, the SIPP has grouped these states together thereby making it impossible to assign the appropriate state laws to individuals in these states.[27]

We restrict our sample to men aged 55-64 who retire during the sample period. The SIPP does not ask individuals directly whether they have retired. We therefore use a measure of retirement based on length of time out of the labor force. This has the advantage, relative to point in time self-reported measures, of capturing transitions to non-work rather than partial (but perceived) retirement. It has the disadvantage, however, of not allowing us to disentangle retirement from other reasons for a temporary absence from the labor force. Following Rogowski and Karoly (1992), we define retirement as a departure from the labor force of 5 or more months.[28] Individuals who are not in the labor force for at least the first four months for which we observe them are excluded from the sample, and individuals who report being out of the labor force in the last 5 months of the panel are censored at the last month for which they are in the labor force.

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[27] These states are Alaska, Idaho, Iowa, Maine, Mississippi, Montana, New Mexico, North Dakota, South Dakota, Vermont and Wyoming. The CPS results are similar if we restrict our CPS sample in the same fashion. See Gruber and Madrian (1993) for more detail on our SIPP sample.

[28] Rogowski and Karoly (1992) actually impose a 6-month rule for departure from the labor force. It turns out that almost all of the individuals who are out of the labor force for 5 months are actually out for 6 or more months. This definition of retirement helps alleviate the problem of measurement error in the reporting of individual labor force status; since individuals are interviewed every four months, they must report that they are out of the labor force in two consecutive interviews to be counted as retired. See Gruber and Madrian (1993) for a further discussion.

Table 12 presents the results from a probit equation for whether or not an individual is covered by employer-provided health insurance after retirement. The key independent variable is the number of months of continuation coverage available at the time of retirement. The results suggest that an extra month of continuation coverage increases the probability of being insured after retirement by 0.5%. This implies that one year of coverage would increase the probability of being insured by 6%, while 18 months would increase the probability of coverage by 9%, a result consistent with that found by Rogowski and Karoly (1992). The results of Table 13 corroborate the evidence on take-up rates presented in Section II. As mentioned, Zedlewski (1993) estimates that 5.2% of retired individuals between the ages of 55 and 64 are covered by COBRA. This fraction is very similar to our 6% estimated increase in coverage from one year of continuation coverage which is the average length of time for which older individuals receive COBRA (Flynn 1992).

Furthermore, we can reconcile this finding with our estimates of the effect of continuation mandates on retirement. Our findings imply that one year of coverage raised the probability of being retired by about 1.1 percentage points, but that it raises the probability of being insured by 6 percentage points. This suggests that the primary effect of these mandates is "inframarginal". That is, they provide insurance coverage for individuals who would have retired in the absence of these mandates even though they would not have been covered by employer-provided health insurance. Thus, continuation mandates may be policies with a sizeable "bang for the buck": they have a large and significant effect along their intended dimension, increased insurance coverage, with a relatively small effect along their unintended dimension, increased retirement.

VII. Conclusion

A number of current policy proposals in the U.S., such as increasing the age of Medicare eligibilty to 67 or providing guaranteed health insurance coverage for all citizens, would affect the health insurance coverage of early retirees. Thus, it seems especially important at this time to understand the interaction between insurance coverage and the retirement decision. If retirement is very sensitive to insurance coverage, for example, it could have important public finance implications for policies which provide universal health insurance coverage; a spate of retirement may non-trivially lower the tax base on which new policies can be financed.

Our strategy for estimating the effect of health insurance on retirement has been to examine the effect of state and federal continuation coverage mandates on retirement propensities. We do this in a static regression framework which allows us to exploit a very large data set and to avoid the problems of dynamic sample selection which plague other studies based on survey data. Our results suggest that continuation mandates have a sizeable and significant effect on retirement. However, contrary to our basic intuition, the effects are not necessarily the strongest at older ages. Rather, taken in conjunction with evidence from dynamic models, we appear to find declining effects by age. We also found that one year of continuation benefits is associated with a 6% increase in insurance coverage levels, suggesting that these policies are not only inducing retirement, but are "inframarginally" covering those who would have retired anyway.

Our use of continuation of coverage regulations as the source of variation for identifying the effect of insurance coverage on retirement has both advantages and disadvantages relative to looking directly at workers with and without employer-provided retiree health insurance. One potential problem with the latter strategy is that the researcher is unable to control for job characteristics which may be correlated with both the generosity of retiree health coverage and the incentives that these jobs offer for retirement. An obvious example is pensions (which are accounted for in both Lumsdaine, Stock and Wise (1992a) and Gustman and Steinmeier (1992)). There may be a number of other ways in which firms encourage or discourage retirement, however, such as through the tasks that they assign older workers or the wage profile that these workers are offered. Furthermore, there may be sorting of workers by retirement propensities into the types of firms that do or do not offer retiree health insurance. To the extent that these are unobserved to the econometrician but correlated with both the offering of retiree coverage and the retirement decision, they will bias the estimated effect of such coverage on retirement. What is needed to identify the effect of retiree health insurance is exogenous assignment of such coverage to individuals that is independent of these other job characteristics. Continuation mandates potentially provide such exogenous assignment.

The primary disadvantage of our strategy is that continuation benefits are more expensive to the early retiree than retirement health insurance and only provide coverage for a limited number of months. These differences may make it unreasonable to extrapolate our results to infer the effects of full retiree health insurance coverage. Future research should focus on combining a study of true employer-provided retiree coverage with an identification strategy that overcomes the ommitted variable bias problems described above.

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TABLE 1

Self-Reported Health Status by Age

	Health Status			
Age	Excellent	Good	Fair	Poor
25-34	36.4%	53.1%	9.5%	1.1%
35-44	32.0	54.6	11.9	1.5
45-54	27.8	52.5	15.6	4.1
55-64	18.0	50.7	24.9	6.4
65+	9.3	43.1	36.1	11.4

Source: Authors' calculations using data from the 1987 National Medical Expenditure Survey. The numbers in the table give the fraction of individuals who report having the given health status.

TABLE 2

Incidence of Health Problems by Age

	Age				
Condition	25-34	35-44	45-54	55-64	65+
Stroke	0.4%	0.8%	1.6%	3.6%	7.4%
Cancer	1.6	2.4	4.7	9.7	13.3
Heart Attack	0.3	1.1	3.8	7.7	13.3
Gallbladder disease	1.6	3.6	7.3	9.4	14.6
High blood pressure	10.1	18.2	29.1	41.9	49.8
Arteriosclerosis	0.2	0.6	2.8	6.1	16.3
Rheumatism	0.8	1.6	5.2	8.2	16.4
Emphysema	0.4	1.0	2.6	5.2	8.0
Arthritis	5.1	11.6	24.9	41.2	54.9
Diabetes	1.7	3.0	5.7	9.8	14.7
Heart disease	0.8	2.2	6.1	11.9	22.2
Any of the above	18.2	31.7	51.8	72.3	84.2

Source: Authors' calculations using data from the 1987 National Medical Expenditure Survey. The numbers in the table give the fraction of individuals who report ever having had the listed medical condition.

TABLE 3

Annual Medical Care Utilization by Age

	Age				
	25-34	35-44	45-54	55-64	65+
A. Fraction admitted to hospital	9.2%	6.8%	8.7%	11.0%	20.1%
Number of admissions (if ever admitted)	1.17	1.24	1.39	1.5	1.5
Nights in hospital (if ever admitted)	5.5	6.8	9.3	11.8	13.8
B. Fraction with prescribed medicines	52.9%	55.6%	61.1%	71.1%	81.9%
Number of prescribed medicines (if any prescribed medicines)	5.2	6.6	11.5	14.7	18.5
C. Fraction who visited a doctor	64.1%	67.1%	71.1%	77.9%	85.8%
Number of doctor visits (if visited a doctor)	4.6	4.6	5.5	6.0	7.4

Source: Authors' calculations using data from the 1987 National Medical Expenditure Survey.

TABLE 4

Average Annual Medical Expenditures by Age ($1980)

	Age					
	25-34	34-44	45-54	55-64	65-74	75+
A. Average Expenditures						
Hospital/Inpatient	$794 ($3763)	$744 ($3186)	$894 ($3648)	$1526 ($6211)	$2142 ($6567)	$3700 ($10,811)
Physician/Outpatient	$334 ($716)	$330 ($653)	$391 ($890)	$473 ($1176)	$543 ($1582)	$560 ($916)
Prescription Medication	$477 ($98)	$65 ($154)	$111 ($208)	$163 ($299)	$195 ($271)	$221 ($276)
Total	$1176 ($4025)	$1135 ($3537)	$1395 ($4001)	$2144 ($6532)	$2877 ($7070)	$4481 ($11,045)
B. Average Expenditure if Expenditure > 0						
Hospital/Inpatient	$2103 ($5900)	$2350 ($5323)	$2289 ($5557)	$3945 ($9502)	$4747 ($9151)	($7482) ($14,218)
Physician/Outpatient	$467 ($807)	$458 ($731)	$543 ($1002)	$592 ($1287)	$668 ($1732)	($662) ($959)
Prescription Medication	$80 ($117)	$111 ($189)	$178 ($243)	$230 ($332)	$258 ($284)	($269) ($282)
Total	$1454 ($4431)	$1428 ($3913)	$1699 ($4357)	$2461 ($6944)	$3270 ($7450)	($4820) ($11,383)

Source: Authors' calculation using data from the 1980 National Medical Care Utilization and Expenditure Survey (inflated to $1990 using the Medical Care Component of the Consumer Price Index). Standard deviation of expenditures is given in parentheses.

TABLE 5

Insurance Coverage by Age and Employment Status

	Employment-Based		Other	Nongroup	CHAMPUS/	Medicare/	Uninsured
	Any	Own Name	Group		CHAMPVA	Medicaid	
A. All Individuals							
25-54	71.6%	51.1%	1.2%	5.9%	5.7%	5.6	15.4%
55-64	64.5	44.8	4.1	14.5	7.7	10.4	12.0
B. Employed							
25-54	78.5%	62.7%	1.1%	5.8%	4.9%	1.2	13.5%
55-64	76.3	63.1	4.0	12.6	6.8	0.8	10.1
C. Not Employed							
25-54	44.2%	4.2%	1.3%	6.2%	8.8%	23.4	23.0%
55-64	51.6	24.7	4.3	16.6	9.2	20.9	14.1

Source: Authors' calculations using data from the 1987 National Medical Expenditure Survey.

TABLE 6

Group and Nongroup Health Insurance Benefits

	Fraction of Individuals with Specified Benefit	
	Group Plans	Nongroup Plans
A. Primary benefits		
Major medical coverage	86.9%	39.1%
Hospital room and board	98.4	91.4
Surgery	97.6	91.6
Physician office visit	87.9	40.4
B. Other benefits		
Ambulance	89.0	54.0
Outpatient diagnostic services	95.9	66.0
Prescribed medicines	87.3	30.3
Mental health	92.2	66.0
C. Generosity of benefits (conditional on having benefit)		
Major medical deductible < $100	94.3	61.6
Full semi-private room charge	77.8	38.2
80-100% of UCR surgical charge	70.6	60.0
80-100% of UCR physician charge	91.8	81.3

Source: Farley (1986), Tables 45-58.

TABLE 7

State Continuation of Coverage Laws

State	Effective Date	Months of Coverage	State	Effective Date	Months of Coverage
Arkansas	7/20/79	4	North Carolina	1/1/82	3
California	1/1/85	3	North Dakota	7/1/83	10
Colorado	7/1/86	3	New York	1/1/86	6
Connecticut	10/1/75	10	Oklahoma	1/1/76	1
	1/1/87	20	Oregon	1/1/82	6
Georgia	7/1/86	3	Rhode Island	1/1/88	18
Illinois	1/1/84	6	South Carolina	1/1/79	2
	8/23/85	9		1/1/90	6
Iowa	7/1/87	9	South Dakota	7/1/84	3
Kansas	1/1/78	6	South Dakota	3/3/88	18
Kentucky	7/15/80	9	Tennessee	1/1/81	3
Minnesota	8/1/74	6	Texas	1/1/81	6
	3/19/83	12	Utah	7/1/86	2
	6/1/87	18	Vermont	5/14/86	6
Missouri	9/28/85	9	Virginia	4/17/86	3
Nevada	1/1/88	18	Wisconsin	5/14/80	18
New Hampshire	8/22/81	10			
New Mexico	7/1/83	6			

Sources: Hewitt (1985), Thompson Publishing Group (1992), and state statutes.

TABLE 8

Health Insurance Coverage Before and After COBRA

	All Individuals		Employed		Not Employed	
	25-54	55-64	25-54	55-64	25-54	55-64
I. Insurance Coverage in 1984						
Any private health insurance	82.1	83.7	89.1	92.5	60.1	74.1
Health insurance in own name						
Employment-based	52.1	47.4	66.7	68.9	5.9	23.6
Not employment-based	5.1	12.5	5.1	10.4	5.2	14.7
Covered as a dependent	24.2	23.4	16.8	12.8	47.7	35.0
II. Insurance Coverage in 1989						
Any private health insurance	82.4	84.3	88.6	92.1	57.3	74.9
Health insurance in own name						
Employment-based	54.7	49.2	66.4	68.1	7.1	26.6
Not employment-based	5.3	12.9	5.2	9.6	5.2	16.8
Covered as a dependent	22.0	21.8	16.4	14.4	43.7	30.6

Source: Authors' calculations using data from the Survey of Income and Program Participation, 1984 Wave 3 and 1987 Wave 7.

TABLE 9

The Effect of Continuation Coverage on the Probability of Being Retired

	Definition of Retired			
	Report Being Retired		Not in the Labor Force	
Independent Variable	Coefficient (st. error)	Marginal Probability	Coefficient (st. error)	Marginal Probability
Months of Coverage	.0036 (.0017)	.0107	.0025 (.0015)	.0098
Married	-.0154 (.0010)	-.0037	-.0577 (.0009)	-.0187
Education	-.0655 (.0092)	-.0162	-.3427 (.0081)	-.1173
Non-white	-.1204 (.0121)	-.0282	.0918 (.0104)	.0305
55 Years Old	-1.205 (.0503)	-.1950	.1180 (.0443)	.0392
56 Years Old	-1.097 (.0502)	-.1853	.1935 (.0443)	.0646
57 Years Old	-1.016 (.0501)	-.1770	.2435 (.0443)	.0816
58 Years Old	-.9251 (.0499)	-.1669	.3157 (.0442)	.1063
59 Years Old	-.8115 (.0498)	-.1525	.4094 (.0442)	.1385
60 Years Old	-.6254 (.0496)	-.1254	.5302 (.0441)	.1804
61 Years Old	-.4903 (.0496)	-.1024	.6394 (.0442)	.2187
62 Years Old	-.0854 (.0494)	-.0203	.9977 (.0441)	.3441
63 Years Old	.1033 (.0494)	.0260	1.161 (.0442)	.3996
64 Years Old	.1938 (.0494)	.0504	1.262 (.0442)	.4324

The table gives estimates from a probit equation for whether or not an individual is retired using data from the 1980-1990 Merged Outgoing Rotation Groups of the CPS. The sample is comprised of 214,508 men aged 55-64. Coefficients year, month and state dummies are not reported.

TABLE 10

The Age-Specific Effect of Continuation Coverage on the Probability of Being Retired

	Definition of Retired					
	Report Being Retired			Not in the Labor Force		
Independent Variable	Coefficient (st. error)	Marginal Probability	Percent of baseline	Coefficient (st. error)	Marginal Probability	Percent of Baseline
55*Months	.0028 (.0023)	.0083	13.3%	.0012 (.0019)	.0047	2.4%
56*Months	.0013 (.0023)	.0037	4.8	.0022 (.0019)	.0088	4.1
57*Months	.0021 (.0022)	.0061	6.8	-.0005 (.0019)	-.0021	0.9
58*Months	.0027 (.0022)	.0080	7.6	.0008 (.0019)	.0031	1.2
59*Months	.0046 (.0021)	.0135	10.6	.0019 (.0019)	.0074	2.6
60*Months	.0024 (.0021)	.0071	4.2	.0018 (.0018)	.0069	2.1
61*Months	.0020 (.0020)	.0060	2.9	.0021 (.0018)	.0085	2.3
62*Months	.0048 (.0020)	.0143	4.2	.0040 (.0018)	.0161	3.2
63*Months	.0041 (.0020)	.0121	2.9	.0045 (.0018)	.0179	3.2
64*Months	.0067 (.0020)	.0202	4.5	.0063 (.0018)	.0251	4.1

The table gives estimates from a probit equation for whether or not an individual is retired using data from the 1980-1990 Merged Outgoing Rotation Groups of the CPS. The sample is comprised of 214,508 men aged 55-64. Coefficients for year, month, age and state dummies are not reported. Education, race, and marital status are also included.

TABLE 11

The Effect of Continuation Coverage on the Probability of Being Retired (Fixed-Trend Included)

	Definition of Retired					
	Report Being Retired			Not in the Labor Force		
Independent Variable	Coefficient (st. error)	Marginal Probability	Percent of baseline	Coefficient (st. error)	Marginal Probability	Percent of Baseline
Age Effects Equal						
Months of coverage	.0045 (.0020)	.0133	6.7%	.00027 (.0017)	.0105	3.0%
Age-Specific Effects						
55*Months	.0037 (.0025)	.0108	17.3%	.0014 (.0021)	.0054	2.8%
56*Months	.0021 (.0025)	.0061	8.0	.0024 (.0021)	.0094	4.4
57*Months	.0030 (.0024)	.0087	9.7	-.0003 (.0021)	-.0013	0.5
58*Months	.0035 (.0024)	.0105	10.0	.0010 (.0021)	.0039	1.5
59*Months	.0054 (.0023)	.0160	12.6	.0020 (.0020)	.0081	2.9
60*Months	.0033 (.0023)	.0097	5.7	.0020 (.0020)	.0077	2.4
61*Months	.0028 (.0023)	.0084	4.1	.0023 (.0020)	.0091	2.5
62*Months	.0056 (.0022)	.0169	5.0	.0042 (.0020)	.0167	3.3
63*Months	.0049 (.0022)	.0146	3.6	.0046 (.0020)	.0184	3.2
64*Months	.0075 (.0022)	.0227	5.1	.0065 (.0020)	.0257	4.2

The table gives estimates from a probit equation for whether or not an individual is retired using data from the 1980-1990 Merged Outgoing Rotation Groups of the CPS. The sample is comprised of 214,508 men aged 55-64. Coefficients for year, month, age and state dummies are not reported. Education, race, and marital status and state-specific trends are also included.

TABLE 12

Continuation Coverage and the Probability of
Being Insured After Retirement

Independent Variable	Coefficient (st. error)	Marginal Probability
Married	.0820 (.1764)	.026
Black	-.8403 (.2227)	-.293
Education	.0381 (.0122)	.012
Age	.1788 (.0903)	-.020
Age^2	-.0001 (.00006)	–
Months of Coverage	.0163 (.0084)	.005

The table gives estimates of the probability of being insured after retirement using data from the Survey of Income and Program Participation. The sample is comprised of 527 men aged 55-64 who retire over the sample period. Coefficients for industry and occupation dummies are not reported.

Figure 1: Modelling the Effect of Continuation Coverage on Retirement

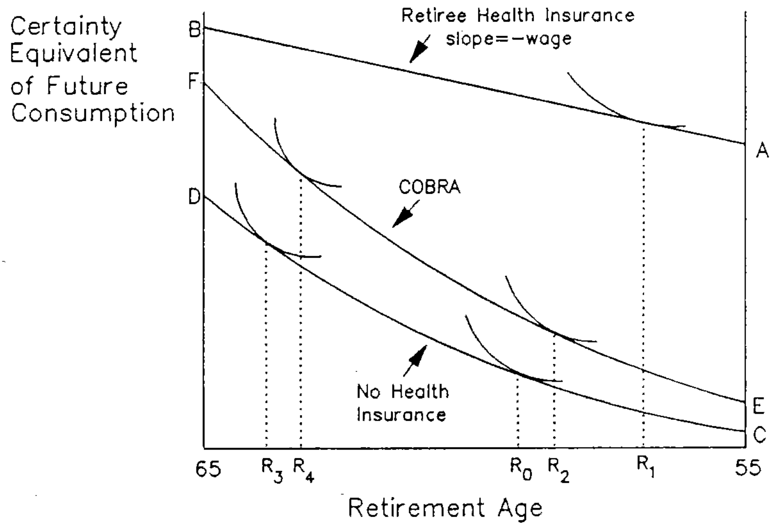

Figure 2: Relative Age-Specific Effects of Continuation Coverage on Retirement