Health-Insurance Availability and the Retirement Decision

By Jonathan Gruber and Brigitte C. Madrian*

There is an extensive literature that examines the retirement behavior of older workers. This research has focused primarily on the effects of social security, private pensions, disability insurance, and health status. One factor that has not received much attention is the role of health insurance, even though the high and uncertain medical costs of older individuals should make the availability of health insurance an important determinant of the timing of retirement.

In this paper, we address this gap by examining the effect of state and federal "continuation of coverage" mandates on the retirement decision. These mandates grant individuals the right to continue purchasing health insurance through a previous employer for a specified number of months after leaving the firm. Although individuals must pay the average employer cost of their group insurance, the price to the early retiree is typically well below that of a policy purchased in the individual market. By reducing the cost of health insurance after retirement, continuation-of-coverage benefits may increase the likelihood of early retirement for those individuals whose employers do not already provide postretirement health insurance. The differences across states and over time in the availability and generosity of these laws provide the variation needed to assess the sensitivity of

retirement behavior to this type of subsidized insurance.1

1Four recent papers have studied the effect of postretirement health insurance on retirement behavior, finding mixed results. Two use information on actual firm provision of postretirement coverage (Alan L. Gustman and Thomas L. Steinmeier, 1994; Madrian, 1994), one imputes such provision based on firm size and industry (Lynn A. Karoly and Jeannette A. Rogowski, 1994), and the last examines the effect of Medicare (Robin L. Lumsdaine et al., 1994). One problem faced by all of these studies is that they may be unable to control for job characteristics that are correlated with both the generosity of retiree health coverage and the incentives that these jobs offer for retirement (i.e., pension incentives or firm-specific age/earnings profiles). Furthermore, there may be sorting of workers by retirement propensities into the types of firms that do or do not offer retiree health insurance. What is needed to identify this effect of retiree health insurance is an exogenous assignment of such coverage to individuals that is independent of these other job characteristics. Continuation mandates potentially provide such an exogenous assignment.

1. Background

The importance of health insurance for the potential early retiree is underscored by the high medical costs of older individuals. Individuals aged 55–64 are almost three times as likely as those aged 25–54 to report themselves in fair or poor health; the likelihood of having cancer or a heart attack is four times greater for those aged 55–64 than for those 25–54; and the medical expenses of older individuals are almost twice as large, and twice as variable, as those of their younger counterparts.2

Despite their higher medical costs, the extent of insurance coverage among 55–64-year-olds is similar to that of 25–54-year-olds. Overall, 12 percent of 55–64-year-olds are uninsured, compared to 15.4 percent of

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25–54-year-olds. Half of nonworking older individuals are covered by employer-provided insurance, either in their own name or a spouse’s, which reflects the fact that 45 percent of individuals work in firms that provide retiree health insurance (Madrian, 1994). However, 31 percent of older non-workers are either uninsured (14 percent) or purchase health insurance in the individual market (17 percent). It is this group of nonworking older individuals whom we would expect to benefit from the availability of continuation coverage.4

Continuation-of-coverage mandates require 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 first such law was implemented by Minnesota in 1974. More than 20 states passed similar laws over the next decade before the federal government, in 1986, mandated such coverage at the national level under COBRA. The various state statutes are summarized in Table 1.5 The length of coverage is often quite short, usually 3–6 months, although 12 states mandate coverage of nine months or more.

Both the state and federal laws stipulate that the employee must pay the full cost of coverage. At the federal level, this is defined specifically as 102 percent of the average (not age-specific) employer cost of providing coverage. Although 102 percent of the average employer cost is typically much more than individuals pay as active employees, it is substantially less than the cost of buying equivalent coverage in the private market for older workers for at least two reasons. First, workplace pooling, by reducing administrative expenses and the possibilities for adverse selection, is estimated to lower the cost of health insurance in large (10,000 or more employees) firms relative to small (1–4 employees) firms by 35 percent (Congressional Research Service, 1988). For older individuals, the cost differential between employer-provided and individual health insurance is further exacerbated by

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3Rates of insurance coverage are tabulations by the authors, using the 1987 National Medical Expenditure Survey.
4This figure constitutes a lower bound on the fraction of individuals who would benefit from continuation coverage, since a number of those reporting employment-based insurance coverage may already be availing themselves of continuation benefits.
5Details on state laws are from Hewitt Associates (1985) and Thompson Publishing Group (1992) and have been cross-checked against the actual statutes. 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; we do not include these states. For a more detailed discussion of these laws, see Gruber and Madrian (1994).
the fact that policies in the individual market are typically age-rated while continuation coverage is not, so that younger and more healthy workers subsidize its cost. Moreover, individual policies almost always place exclusions on preexisting conditions, and many are medically underwritten as well; as a consequence, those with health problems may find themselves liable for many of their medical expenses that were previously covered by insurance.

Continuation coverage is therefore likely to be an attractive option for many early retirees. This attractiveness is reflected in the high take-up of this coverage among those eligible. We estimate that roughly two-thirds of eligible early retirees whose employers do not offer postretirement health insurance take up continuation coverage.6

What effect should we expect continuation benefits to have on the retirement decision? Essentially, continuation benefits provide a subsidy whose value is the certainty equivalent of the difference between the cost of 18 months of group health insurance and 18 months of self-insuring, be it through purchasing private health insurance or actually going without. If leisure is a normal good, this subsidy will lead individuals to retire earlier than they would otherwise.

The age pattern of these retirement effects is unclear ex ante. If individuals were infinitely risk-averse or if coverage in the private market were infinitely expensive, then these laws would necessarily act only as a "bridge to Medicare," allowing individuals to retire a certain number of months before age 65 without exposing themselves to any period without group coverage. However, as noted earlier, over 30 percent of early retirees appear willing to face the vagaries of the individual insurance (or self-insurance) market. For these individuals, the continuation subsidy may simply serve to push forward their retirement dates. Fur-

thermore, younger retirees may use continuation benefits not as a bridge to Medicare, but as a bridge to another job; Peter A. Diamond and Jerry A. Hausman (1984) and Glenn T. Sueyoshi (1989) report high rates of reentry to the labor force from retirement and partial retirement.

II. Empirical Results from the Current Population Survey

Our primary data source for modeling the transition to retirement is the Census Bureau's March Current Population Survey (CPS) for the years 1980–1990. Our sample consists of roughly 5,000 55–64-year-old men from each year. Each March, the CPS asks not only about current labor-force attachment, but about labor-force attachment in the previous year as well, including a number of questions about the characteristics of the longest job held during that year. We can therefore observe transitions into retirement over a one-year interval. We define individuals as retiring if they worked at least one week in the previous year but report being retired at the time of the survey.7 We restrict the sample to those whose previous jobs provided health insurance, since individuals without health insurance have nothing to continue when they retire. The retirement rate in our sample is 6.9 percent, and the average length of continuation coverage is 6.8 months.

We begin with a simple "difference-in-difference" analysis of the effect of the federal continuation mandate. If continuation coverage encourages retirement, we should see an increase in the retirement rate of individuals living in states affected by the federal law. Because several states had already enacted their own continuation mandates before the federal law was passed, individuals in these states provide a natural

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6This calculation uses data on COBRA take-up rates from Patrice Flynn (1992); details of the calculation are available upon request.

7The estimated coefficients are very similar if we restrict the sample to individuals who worked at least 10 weeks in the previous year. The results are also similar for all persons who report themselves to be out of the labor force, rather than only retired, in the March interview.
control group for measuring the effect of the federal law. However, to the extent that the number of months of coverage available affects the retirement decision, the federal law, which allows for 18 months of coverage, may have some effect on retirement rates in the control states, most of which provided continuation benefits for a shorter period of time (before the federal law was passed, only Wisconsin allowed for 18 months of coverage). But for states with continuation mandates in place before 1986, the federal law increased the duration of continuation coverage by less than 13 months; for states without such laws, duration rose by the full 18 months.

Using the CPS, we calculate the average retirement probability for the years 1983–1985 (before the federal law change) and for 1988–1990 (after the federal law change) for individuals in states with and without pre-COBRA continuation mandates. These results are presented in Table 2. For individuals in states without a continuation mandate prior to the federal law, we see a sizable 2-percentage-point increase in the average retirement rate in the post-COBRA period (column 1). At the same time, the retirement rate was roughly constant in the states that already had a mandate in place (column 2). On net, the difference-in-difference estimate of the effect of the federal law is a 1.7-percentage-point increase in the retirement rate, an effect that is significant at the 7-percent level. This represents 27.6 percent of the baseline retirement rate for our sample, and it suggests an important role for continuation mandates in determining retirement behavior.

While straightforward, this simple estimate does not exploit all of the available variation in continuation mandates, which arises both from changes in state and federal laws over time and from cross-sectional differences in the length of coverage available. We therefore use the entire 1980–1990 CPS sample to estimate the following probit model of retirement:

\[ \Pr(RETIRE_{ijt}) = \Phi[\alpha + \beta_1(X_{ijt}) + \beta_2(STATE_j) + \beta_3(TIME_t) + \beta_4(LAW_{jt})] \]

where \( i \) indexes individuals, \( j \) indexes states, and \( t \) indexes time. The vector \( X_{ijt} \) contains a set of individual demographic and job characteristics: race, marital status, education, and single-year age dummies; eight industry and six occupation controls; a control for having health insurance on the job; and a control for being covered by a pension on the job. \( STATE_j \) is a set of state dummies, and \( TIME_t \) is a set of year dummies. \( LAW_{jt} \) is the number of months of continuation coverage available in state \( j \) at time \( t \). The state fixed effects control for any fixed state-specific omitted variables correlated with the propensity to retire. The year dummies control for national trends in retirement behavior which may be correlated with the passage of continuation-of-coverage laws. Thus, the effect of the laws is identified in this model by changes in retirement behavior in states that passed the laws (or were affected by the federal law) relative to those that 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 that these laws allow.
The results from this "flow" probit are presented in Table 3. The sample in the first column is restricted to individuals covered by employer-provided health insurance. Although the coefficients are not shown, neither marital status nor education has a significant effect on the probability of retiring; nonwhites, however, have significantly lower retirement rates. Having a pension significantly raises the probability of retiring.\textsuperscript{9} The age coefficients imply that the probability of retiring increases with age, with a spike at age 62 that is generally attributed to the eligibility for social-security early-retirement benefits at that age. The coefficient on months of coverage (the coefficient of interest) is reported in the first row. It is statistically significant and suggests that one year of continuation coverage increases the probability of retiring by 2.2 percentage points. We also present (in square brackets) the implied effect of one year of continuation coverage on the retirement hazard; we find that one year of coverage raises the hazard by 32.4 percent. This is a sizable effect for such limited insurance coverage.

We have attempted a number of specification checks to explore the robustness of this finding. One alternative hypothesis for this result is that of "legislative endogeneity": states where more individuals are retiring respond by mandating benefits that cover individuals after their retirement. There is a natural test for this alternative hypothesis, however. Note that only individuals who have health insurance on the job are eligible to receive continuation benefits. Thus, workers without health insurance on the job provide a control group for assessing the effects of continuation mandates. If these mandates are due to exogenous changes in retirement propensities, then the laws should be correlated with the retirement propensities of workers both with and without health insurance on the job. However, if the laws are causing changes in retirement propensities, this should only affect workers with employer-provided health insurance.

In the right-hand columns of Table 3, we report the results for those without health insurance. The coefficient on months of coverage is negative and insignificant. Thus, to the extent that uninsured workers in a given state/year provide a valid control group for other factors determining retirement behavior, this suggests that the continuation mandates were having a causal effect.\textsuperscript{9}

Earlier we discussed the possibility that continuation mandates may have different effects at different ages. We investigate this empirically in Table 3 (model 2) by allowing the effect of continuation mandates to differ by age. For those with health insurance, the effects are very strong at each age and are significant at the 10-percent level for nine of the ten age interactions. However, there is no obvious pattern to the age coefficients, and we cannot reject the restriction that the coefficients are the same at all ages between 55 and 64. Expressed as hazard

\textsuperscript{9}One potential problem with the estimation is that the state continuation mandates do not apply to firms that self-insure, the federal mandate does not apply to firms with fewer than 20 employees, and neither of the mandates is binding on firms that provide retiree health insurance. Correcting the estimates for state-specific variation in self-insurance, retiree insurance, and firm size does not affect our inferences, although it increases the magnitude of the estimated effect. Another potential problem is that our sample selection rule, that individuals be working in the previous year, introduces a potential left-censoring bias into our estimation, since the older individuals still working after a mandate has been put into place have revealed themselves to be less sensitive to this incentive. When we test for the importance of this bias by restricting our sample to the first year that a mandate is in place, we find almost precisely the same coefficients as in Table 3.

\textsuperscript{8}One potential weakness of our study is that we do not have better data on pension incentives. This will only affect our results if pension incentives changed systematically in a way that was correlated with the passage of continuation mandates. While we are unable to investigate empirically the changes in pension incentives across states and time, there is an insignificant negative relationship between pension coverage and continuation mandates, suggesting that, if anything, changes in pension coverage bias our results downward.
<table>
<thead>
<tr>
<th>Independent variables</th>
<th>Employer-provided health insurance</th>
<th>No employer-provided health insurance</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Coefficient (SE)</td>
<td>Marginal probability $\times 10^2$ [hazard ratio]</td>
</tr>
<tr>
<td><strong>Model 1:</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Months of coverage</td>
<td>0.0143 (0.0057)</td>
<td>2.225 [1.324]</td>
</tr>
<tr>
<td><strong>Model 2:</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Age 55 × months</td>
<td>0.0089 (0.0072)</td>
<td>1.337 [1.562]</td>
</tr>
<tr>
<td>Age 56 × months</td>
<td>0.0132 (0.0072)</td>
<td>2.042 [1.804]</td>
</tr>
<tr>
<td>Age 57 × months</td>
<td>0.0156 (0.0072)</td>
<td>2.444 [1.921]</td>
</tr>
<tr>
<td>Age 58 × months</td>
<td>0.0170 (0.0071)</td>
<td>2.694 [1.872]</td>
</tr>
<tr>
<td>Age 59 × months</td>
<td>0.0138 (0.0070)</td>
<td>2.140 [1.596]</td>
</tr>
<tr>
<td>Age 60 × months</td>
<td>0.0166 (0.0067)</td>
<td>2.608 [1.363]</td>
</tr>
<tr>
<td>Age 61 × months</td>
<td>0.0133 (0.0067)</td>
<td>2.040 [1.271]</td>
</tr>
<tr>
<td>Age 62 × months</td>
<td>0.0107 (0.0063)</td>
<td>1.617 [1.084]</td>
</tr>
<tr>
<td>Age 63 × months</td>
<td>0.0167 (0.0065)</td>
<td>2.636 [1.146]</td>
</tr>
<tr>
<td>Age 64 × months</td>
<td>0.0181 (0.0068)</td>
<td>2.881 [1.174]</td>
</tr>
</tbody>
</table>

**Model 3:**
- Continuation dummy: 0.1357 (0.0722) [1.595 (1.233)]
- Months of coverage: 0.0005 (0.0094) [0.077 (1.011)]

<table>
<thead>
<tr>
<th>Coefficient (SE)</th>
<th>Marginal probability $\times 10^2$ [hazard ratio]</th>
<th>Coefficient (SE)</th>
<th>Marginal probability $\times 10^2$ [hazard ratio]</th>
</tr>
</thead>
</table>

**Notes:** The table shows estimates of a probit model of retirement using a sample of 56,180 men aged 55–64 from the 1980–1990 Current Population Surveys. Of these, 39,123 have employer-provided health insurance, and 17,057 do not. Coefficients for marital status, education, race, health-insurance and pension coverage, age, year, state, industry, and occupation dummies are not reported. Model 2 gives coefficients from a second specification which instead interacts the months of coverage available with each age separately; model 3 includes a dummy for the presence of a continuation mandate along with the number of months of coverage available. Marginal probabilities are given in the second and fourth columns; hazard ratios are given in brackets.

rates, in fact, the effect of continuation mandates actually declines with age. On the other hand, the results for those without health insurance are uniformly insignificant and generally wrong-signed, supporting our contention that continuation mandates are having a causal effect. This pattern of age effects runs counter to the simple presumption that these laws act as a bridge to Medicare, and it may reflect a uniform shift toward early retirement noted above. We discuss this further below.

### III. Empirical Results from the Survey of Income and Program Participation

Several previous studies have suggested that retirement is more naturally modeled in a hazard framework (see e.g., Diamond and Hausman, 1984; and Hausman and...
David A. Wise, 1985). We have therefore also estimated a hazard model of retirement using data from the 1984–1987 panels of the Survey of Income and Program Participation (SIPP). The SIPP is a longitudinal data set in which sample members are interviewed every four months for roughly 2.3 years and asked to provide information about their labor-market activity, income, and participation in welfare and transfer programs over the previous four months. 

As with the CPS data, we restrict our sample to men aged 55–64. Individuals are included in the sample upon turning 55 and are censored upon reaching their 65th birthday. We exclude individuals who are in the sample for only one wave and individuals already retired when we first observe them. Unfortunately, the SIPP does not ask individuals directly whether they have retired. We therefore define retirement to occur when we first see an individual leave the labor force. The set of covariates that affect the hazard include marital status, race, education, industry, occupation, and whether or not an individual has health insurance on the job. We include a set of dummies for the year of the SIPP panel in order to capture time trends in retirement, and we include a dummy for each calendar month in order to model seasonality. We also include a dummy for whether a particular month is an interview (or “seam”) month, since there is some tendency in the SIPP for individuals to propagate their current employment status back through the entire quarter covered by the interview. Since the much smaller sample size and shorter time period in the SIPP make identification from state deviations more tenuous, we do not include state dummies; the results are somewhat strengthened if such dummies are included. Once again, we restrict the basic sample to those who had health insurance on the job at the start of the sample.

Table 4 gives the coefficient estimates and corresponding hazard ratios from our basic hazard specification using data from the SIPP. The hazard ratio gives the effect on the retirement hazard of one year of continuation coverage. Although the coefficients are not presented, those who are married, nonwhite, and more educated and who have health insurance are less likely to retire, but none of the estimated coefficients is significant. As shown in the first two columns of Table 4, however, the coefficient on months of coverage is statistically significant for those who have employer-provided health insurance. It indicates that one year of continuation coverage raises the probability of retirement by 32.1 percent, an effect which is almost exactly the same as that measured from the CPS. In the right-hand columns, when we restrict the estimation to those without health insurance on the job, the effect of months of coverage is substantively and statistically insignificant; this result is also consistent with that found in the CPS.

Table 4 also explores the age pattern of the continuation-coverage coefficients in the SIPP (model 2). Once again, the presumption that the effects should be stronger near the age of Medicare eligibility is not borne out by the data. In fact, as in the CPS data, the hazard rates decline with age, and they are even negative at the oldest ages (although insignificant). As with the overall results, the effects are much stronger for those with health insurance than for those without health insurance at most ages. Thus, the SIPP confirms the age pattern found in the CPS. Continuation mandates are clearly not acting simply as a “bridge to Medicare.”
### Table 4—The Effect of Continuation Coverage on Retirement
(Survey of Income and Program Participation Data)

<table>
<thead>
<tr>
<th>Independent variable</th>
<th>Employer-provided health insurance</th>
<th>No employer-provided health insurance</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Coefficient (SE)</td>
<td>Hazard ratio</td>
</tr>
<tr>
<td><strong>Model 1:</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Months of coverage</td>
<td>0.0232 (0.0075)</td>
<td>1.321</td>
</tr>
<tr>
<td><strong>Model 2:</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Age 55 × months</td>
<td>0.0688 (0.0180)</td>
<td>2.283</td>
</tr>
<tr>
<td>Age 56 × months</td>
<td>0.0689 (0.0179)</td>
<td>2.286</td>
</tr>
<tr>
<td>Age 57 × months</td>
<td>0.0526 (0.0180)</td>
<td>1.880</td>
</tr>
<tr>
<td>Age 58 × months</td>
<td>0.0716 (0.0156)</td>
<td>2.361</td>
</tr>
<tr>
<td>Age 59 × months</td>
<td>0.0303 (0.0179)</td>
<td>1.438</td>
</tr>
<tr>
<td>Age 60 × months</td>
<td>0.0039 (0.0166)</td>
<td>1.048</td>
</tr>
<tr>
<td>Age 61 × months</td>
<td>0.0210 (0.0152)</td>
<td>1.287</td>
</tr>
<tr>
<td>Age 62 × months</td>
<td>-0.0040 (0.0129)</td>
<td>0.953</td>
</tr>
<tr>
<td>Age 63 × months</td>
<td>0.0073 (0.0144)</td>
<td>1.091</td>
</tr>
<tr>
<td>Age 64 × months</td>
<td>-0.0143 (0.0169)</td>
<td>0.842</td>
</tr>
<tr>
<td><strong>Model 3:</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Continuation dummy</td>
<td>0.0309 (0.1288)</td>
<td>1.031</td>
</tr>
<tr>
<td>Months of coverage</td>
<td>0.0216 (0.0102)</td>
<td>1.296</td>
</tr>
</tbody>
</table>

**Notes:** The table shows estimates of a proportional-hazard model of age at retirement for 4,399 men aged 55–64 using data from the 1984, 1985, 1986, and 1987 panels of the Survey of Income and Program Participation. Coefficients for marital status, education, race, health-insurance coverage, industry, occupation, month, panel, and seam dummies are not reported. Model 2 gives coefficients from a second specification which instead interacts the months of coverage available with each age separately; model 3 includes a dummy for the presence of a continuation mandate along with the number of months of coverage available.

One reason for this, noted in Section II, is that younger "retirees" may be using continuation benefits as a bridge to further employment. If so, our estimates may reflect both the effects of continuation coverage on job turnover and their effects on retirement. In fact, calculations from our SIPP data indicate that of those who leave the labor force for at least one month between the ages of 55 and 62, 40 percent return to the labor force within the next 12 months. Between the ages of 62 and 65, this reentry rate drops to 20 percent. This raises the question of whether our findings are actually effects on retirement, or just effects on turnover. In other research we have found that one year of continuation benefits increases the likelihood of job turnover among younger workers by about 10 percent (Gruber and Madrian, 1994). Given that, on average, 32 percent of our SIPP sample reenters the labor force, this implies that at
most 16 percent of our estimated retirement effect is accounted for by job changers.12

Other reasons for this finding could be that retirees are myopic in their perceived need for coverage in future years or that they do not understand that the continuation mandates provide coverage for a limited time only. In order to assess this view, we include both a dummy for the presence of a continuation mandate and the number of months of eligibility in our regressions. These results are presented in the bottom panel of Table 3 for the CPS and the bottom panel of Table 4 for the SIPP. If individuals respond rationally and with full information about the generosity of these mandates, the presence of a law should be irrelevant when compared to the number of months of coverage provided.

In the CPS, we find that the coefficients on months of coverage becomes insignificant when the dummy is included. The coefficient on the continuation dummy itself is also insignificant, but the implied marginal effect of having coverage is roughly two-thirds of the earlier estimated effect of one year of coverage. In the SIPP, on the other hand, we find that the continuation dummy is completely insignificant, and the months coefficient is virtually unchanged from the first row of the table. Thus, while the CPS data suggest that there may either be some myopia or misinformation about law generosity, the SIPP data suggest that individuals are responding in the expected fashion.

IV. Implications and Conclusions

To place our findings in context, it is useful to compare them to the estimated increase in retirement propensities induced by an increase in postretirement income. In the dynamic stochastic programming model employed by James H. Stock and Wise (1990) and Lumsdaine et al. (1994), a $5,000 increase in the value of pension wealth is associated with an increase in the retirement hazard of 11.1 percent for individuals between the ages of 55 and 64.13 Our basic specification suggests that one year of continuation coverage raises the retirement hazard by 30 percent. This implies that one year of continuation coverage is valued at $13,600, an amount much higher than the typical cost differential between purchasing health insurance in the individual market and continuing coverage through one's employer.14 This may reflect the fact that almost all individual policies exclude preexisting conditions for some period, thereby understating the true cost difference. Alternatively, it may be that a number of early retirees are unable to obtain individual insurance at all.

We can assess the implied risk aversion of a 55–64-year-old worker who would pay $13,600 for continuation coverage rather than go uninsured, using data on the distribution of family medical expenditures for this age group from the 1987 National Medical Expenditure Survey (NMES). Under the assumption of a constant-relative-risk-aversion utility function that is separable in consumption (net of medical expenditures) and its other arguments, this premium is consistent with a coefficient of relative risk aversion of approximately 2.15 This is within

12Furthermore, our definition of retirement in the SIPP is not as stringent as that in the CPS, which requires that individuals actually label themselves as retired. The reentry rates estimated from the SIPP, and the bias to our retirement effect from job changes, is therefore most likely an upper bound. Using the 15-percent reentry rate estimated by Diamond and Hausman (1984), the bias from job changes is only 8 percent.

13We are grateful to Robin Lumsdaine for performing these calculations for us.

14We estimate that, for a 58-year-old male in Massachusetts, continuation coverage provides a subsidy on the order of $3,600 per year (based on the average price of group coverage and individual coverage in this state).

15This calculation is done as follows. We begin with a sample of 55–64-year-olds in the NMES with private group insurance. We then pose a constant-relative-risk-aversion utility function that is a function of net nonmedical consumption (income minus medical spending). We use this function to find the certainty-equivalent premium for the representative person, which would equate his utility from purchasing insur-
the range of estimated coefficients found in the macroeconomics literature (Stephen P. Zeldes, 1989).

We can also use our estimates to evaluate how much of the observed decline in the labor-force participation of older men during the 1980's can be accounted for by the availability of continuation coverage. In 1980, 16.4 percent of 55–64-year-old men called themselves retired. By 1990, this fraction had increased to 21.7 percent. Using data from the merged outgoing rotation groups of the 1980 CPS, we first calculated the age-specific retirement rates of older men and the implied age-specific retirement hazards. We then increased the retirement hazard by 30 percent at all ages. The resulting retirement rate is 19.7 percent. This suggests that the availability of continuation coverage explains about 60 percent of the net rise in retirement for 55–64-year-old males during the 1980's. This figure represents an upper bound on the true effect of this subsidy, however, as continuation coverage was phased in during the decade; our calculation assumes that 18 months of coverage was made available in 1980.

Overall, our findings imply that policies to provide universal health-insurance coverage (with no out-of-pocket spending) with his utility if he faced an out-of-pocket spending distribution equal to the actual distribution of spending in this population. We search over parameters of risk aversion that yield a certainty-equivalent premium of $17,600 ($13,600 plus $4,000, the average cost of employer-provided health insurance (Employee Benefit Research Institute, 1993 [table 18]). One problem with this calculation is that utility will not be defined for individuals for whom either spending or the continuation premium is greater than income. We therefore drop individuals with income less than $17,600 from the sample, and reset spending to 99 percent of income for those spending more than their income.

Eighteen months of continuation coverage increases the retirement hazard among those with health insurance by 45 percent; however, only about two-thirds of individuals receive health insurance through their own employment. The overall effect of 18 months of continuation coverage is therefore about a 30-percent increase in the probability of retirement.

On the other hand, a few states did have continuation mandates already in place by 1980.

age could lead to a large increase in the rate of early retirement. This factor should be accounted for in considering the potential financing of such policies. For example, in 1989, 55–64-year-old men earned $193 billion. Our results imply that the 18 months of continuation coverage provided under COBRA increased the fraction of retired 55–64-year-olds by 3.3 percentage points in steady state. If national health insurance has an effect on retirement that is twice as large, then the revenues from financing such health insurance by (for example) a 10-percent payroll tax would be $1.3 billion smaller than would be apparent from a static revenue calculation. Perhaps more important, however, is the effect on output of such a large drop in the productively employed labor force (minus the gain in household production). Future research to assess the magnitude of the resulting output loss would be useful for evaluating insurance policy.

REFERENCES


Authors' tabulations from the March 1990 CPS.


